Paying a Premium on your Premium? Consolidation in the U.S. Health Insurance Industry

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Abstract

We examine whether and to what extent consolidation in the U.S. health insurance industry is leading to higher private insurance premiums. We make use of a proprietary, panel dataset of employer-sponsored healthplans enrolling over 10 million Americans annually between 1998 and 2006 to estimate the relationship between premium growth and changes in market concentration. We exploit the differential impact of a large national merger of two insurance firms across local markets to estimate the causal effect of concentration on market-level premiums. We estimate real premiums increased by 2 percentage points (in a typical market) due to the rise in concentration during our study period. We also find evidence that consolidation facilitates the exercise of monopsonistic power vis a vis physicians, whose absolute employment and relative earnings decline in its wake.

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Although the vast majority of healthcare expenditures in the U.S. are funneled through the health insurance industry, few researchers have examined whether the industry itself is contributing to rising health insurance premiums. This possibility has become ever more salient as consolidations continue in this highly-concentrated sector. In 2001, the American Medical Association (AMA) reported nearly half of the 40 largest Metropolitan Statistical Areas (MSAs) were "highly concentrated," using the *Horizontal Merger Guidelines* cutoff of HHI > 1,800. By 2008, the AMA expanded its annual report to include 314 geographic areas (mainly MSAs), 94 percent of which were found to be highly concentrated.[1] During the same period (corresponding to data years 2000 and 2006), the average premium for a family of four receiving coverage through an employer rose 81 percent, reaching $11,480 in 2006.[2]

This study examines whether there is a causal link between changes in market concentration and growth in health insurance premiums. From a theoretical standpoint, both the sign and the magnitude of the effect of concentration on insurance premiums are ambiguous. On the one hand, increases in market concentration may allow health insurers to raise their markups, leading to higher premiums. On the other hand, increases in market share may strengthen insurers' bargaining positions vis a vis healthcare providers, leading to reduced outlays and lower premiums. In addition, there are many potential sources of efficiency gains from consolidation, including economies of scale in IT investing and disease management programs, which would also reduce costs and optimal premiums.[3] The net effect on insurance premiums is an empirical question.

The key challenges to empirically estimating such a link are adequate data and exogenous variation in market concentration. Comprehensive data on healthplans is extremely difficult to obtain because contracts are customized for each buyer across many different dimensions, renegotiated annually, and considered highly confidential. In addition, premiums vary based on the demographics, health risks, and expenditure history (or "experience") of the insured

[1] "Competition in Health Insurance: A Comprehensive Study of U.S. Markets," American Medical Association, 2001 and 2008. HHI is calculated for the combined HMO and PPO product market. Estimates are not strictly comparable over time due to changes in methodology and sample selection. For example, self-insured HMOs are generally included in 2001 but excluded in 2008.

[2] Premiums include both employer and employee contributions. Source: *Employer Health Benefits Summary of Findings*, 2000 and 2006, Kaiser Family Foundation/Health Research and Educational Trust Survey, http://ehbs.kff.org/.

[3] Rent transfers from providers to insurers are not efficiency gains, although they may reduce premiums.

population. Thus, it is difficult to calculate a standardized premium to enable comparisons across employers and/or over time.

Our study exploits a rich, proprietary panel dataset of healthplans offered by a large sample of U.S. firms. The data, which we call the "Large Employer Health Insurance Dataset" or LEHID, span the period from 1998 to 2006 (inclusive), representing coverage for over 10 million nonelderly Americans each year. By focusing on the *growth* in average health insurance premiums for the same employer in a specific market over time, we account for unobserved, time-invariant differences that might influence health insurance premium levels. We also make use of time-varying data such as employee demographics, the types of plans offered (e.g. HMO, POS, etc.), and the level of copays. These variables serve as controls in models that aim to estimate changes in price for a fixed product, and as dependent variables in models exploring employer reactions to changes in local market concentration.

We begin with a descriptive analysis of the relationship between changes in employer premiums and changes in HHI within the 139 geographic markets defined by our source (and described in detail in the data section). We do not find evidence that premiums are rising more quickly in markets that are becoming more concentrated during the 1998 to 2006 period. Although these estimates are useful for descriptive purposes, they do not provide causal estimates of the impact of market structure on premiums. Changes in HHI are likely to be driven by many factors that are arguably not exogenous to premium growth. These include changes in consumer preferences, changes in product offerings and pricing strategies, and changes in the healthcare provider landscape. For example, consider a market with a struggling local economy. In such a market, consumers may flock to low-priced carriers, bringing about an increase in local market concentration and a simultaneous reduction in average premium growth. This pattern does not imply consolidations in such a market would reduce premium growth, *ceteris paribus.*

To obtain an estimate of the causal impact of concentration on premiums, we exploit sharp and heterogeneous increases in local market concentration generated by the 1999 merger of two industry giants, Aetna and Prudential Healthcare. Both were national firms, active in most local insurance markets, and thus the merger had widespread impact. However, the pre-merger market shares of the two firms varied significantly across local markets, resulting in very different shocks to post-merger competition. For example, in our sample the pre-merger market shares of Aetna and Prudential in Jacksonville, Florida were 19 and 24 percent, respectively, versus 11 and 1 percent, respectively, in Las Vegas, Nevada. Holding all else constant, this implies an increase in post-merger HHI of 892 points in Jacksonville, but only 21 in Las Vegas. (HHI is measured as the sum of squared market shares for each carrier, multiplied by 10,000.) Focusing on the years immediately surrounding this merger, we examine the relationship between premium growth and HHI changes using these predicted changes as instruments for actual changes.

The point estimates indicate that rising concentration in local health insurance markets accounts for a small share of premium growth in recent years. Specifically, our instrumental variables estimates imply that the mean increase in local market HHI during 1998-2006 generated a premium increase of 2.1 percent. Given private insurance premiums of approximately $850 billion in 2009, *if* this result generalizes to all private insurance the "premium on premiums" amounts to $18 billion per year.[4]

Although our focus is on the exercise of market power by insurers in the *output* market, consolidation may also have important effects on *input* prices. Using data on earnings and employment of healthcare personnel, we exploit the Aetna-Prudential merger to examine a causal link between concentration and these outcomes. Our analysis indicates that the growth in insurer bargaining power curtailed growth in earnings and employment for physicians in markets experiencing consolidation, and facilitated substitution toward nurses.

The paper is organized as follows. Section 1 discusses prior related research. Section 2 describes the data in detail. We examine the association between changes in local market concentration and premium growth in Section 3. In Section 4 we estimate a causal relationship between these two variables using the variation in HHI induced by the Aetna-Prudential merger. Section 5 extends the analysis in Section 4, exploiting the Aetna-Prudential merger to estimate the impact of changes in HHI on other outcomes of interest such as the percent of enrollees in HMOs. Section 6 describes our analyses of the relationship between changes in concentration and healthcare employment and earnings. Section 7 concludes.

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[4] Source: National Health Expenditure Data provided by the Center for Medicare and Medicaid Services; available online at <http://www.cms.hhs.gov/NationalHealthExpendData/>

I. Related Research

Our study builds on research from two distinct streams of literature: studies of the relationship between market concentration and competitive outcomes in the empirical industrial organization literature, and studies of the health insurance industry, mainly from the health services literature. In this section, we summarize the key insights of each, and identify our contributions at the end.

A. Price-Concentration Studies in Industrial Organization

The structure-conduct-performance paradigm in industrial organization triggered a wave of empirical studies of the relationship between market concentration and profitability.[5] Using cross-sectional data for a large number of industries, many of these studies documented a positive relationship between profits and concentration. This approach was famously critiqued by Harold Demsetz (1973), who argued that the observed relationship could also be explained by differences in efficiency across firms.[6] Subsequent studies focus on price, an outcome less influenced by this "efficiency critique."

Recent studies in this literature rely on within-industry variation in concentration and price, e.g. by using observations on different geographic markets. Most document higher prices in more concentrated markets. Examples include Morrison and Winston (1990) in airlines, Hannan (1992) in banking, and (Cotterill 1986) in grocery retailing. However, much of this work assumes market structure is exogenously determined with respect to price. Given many of the same unobservable factors determine both, regressions of price on concentration and observable controls likely yield biased estimates.

[5] Although our discussion focuses on studies of horizontal consolidation, researchers have also investigated the impact of vertical consolidation on price (as well as other outcomes). Recent examples of such studies include Cuellar and Gertler (2005) on physician-hospital integration and Hortacsu and Syverson (2007) on integration in the cement and ready-mixed concrete industries.

[6] This approach was also criticized on other fronts, particularly on the failure to control for differences in economic factors across industries, and on the use of accounting measures for profitability.

Recent studies have pursued two distinct approaches to surmount this endogeneity issue. One solution relies on a two-step estimation procedure. In the first step, the authors estimate an equilibrium model predicting the number of competing firms in a market. This model is used to generate a correction term to include in the second-stage regression of price on concentration, much in the same way selection correction terms are included in wage regressions (Heckman 1979). Some recent examples include Manuszak and Moul (2008), who use this method to evaluate the prospective impact of the Staples-Office Depot merger, and Singh and Zhu (2006), who study auto rental markets. Mazzeo (2002) extends this approach to account for the impact of product differentiation by specifically allowing for differences in the competitive effects of firms with different product characteristics. This approach lends itself to estimating welfare changes and performing counterfactual experiments (such as estimating the effects of a merger), but it requires strong assumptions about the behavior of firms to enable an accurate characterization of market structure in the first-stage equation.

The second solution requires variables that can serve as instruments for market structure, i.e. measures that are correlated with market structure but uncorrelated with unobservable factors affecting price. Two of the best-known studies in this vein use lagged market structure as an instrument for current market structure, e.g. Evans, Froeb and Werden (1993) (airlines) and Davis (2004) (movie theaters). For example, Davis explores the relationship between within-theater variation in pricing and geographic market structure, using lagged counts of movie screens owned by own and rival chains within various distances as instruments for their current levels. He finds ownership structure has a statistically significant but economically small effect on admission prices charged to consumers. Unfortunately, using lags of an endogenous variable as an instrument is only valid under relatively strong assumptions.

We also pursue an instrumental variables approach to estimate the causal relationship between market structure and price. Our instrument consists of market-specific shocks induced by a large national merger. To the extent these shocks are both correlated with observed changes in market structure and orthogonal to other determinants of premium growth, our estimates will be unbiased. We are unaware of other studies that explicitly use mergers to instrument for changes in market concentration, although there is certainly a related literature on merger effects. Although this literature is too vast to be summarized here, we note that most of the reduced-form estimates of merger effects suffer from selection bias. Markets or industries in which mergers occur are unlikely to be randomly selected, or to be more precise, to be selected in a way that is unrelated to other determinants of the outcomes of interest. Some merger analyses contend with this selection problem by exploiting a temporal shock that induces additional mergers, e.g. Kim and Singhal (1993) on airlines and Berry and Waldfogel (2001) on radio stations, or by using an instrument to predict which institutions merge (e.g. Dafny 2009 on hospital mergers).

Rather than exploiting multiple exogenously-induced mergers, this study exploits a single merger with different impacts across geographic markets. We carefully consider whether the merger we examine generates plausibly exogenous variation in market concentration. This identification strategy is similar to that of Gilbert and Hastings (2005), who use an acquisition of a West Coast refinery as a source of exogenous variation in the degree of vertical integration across retail gasoline markets in 13 West Coast metropolitan areas. However, the Aetna-Prudential merger affected many more markets, and we are able to examine its effect using a larger sample of micro data.

B. Studies Focusing on the Health Insurance Industry

Several studies published in health economics or health services journals examine the relationship between industry structure and insurance price (i.e., premiums). Robinson (2004) uses a database of state regulatory filings to study the state-level market structure of commercial insurance carriers in 2003. He finds the largest firm controls at least a third of the market in almost 40 states in 2002-03. The top 3 insurance firms control over 50 percent of total enrollment in almost all states. Using a variety of other sources, Robinson also documents a sharp increase in insurer revenues and profits over the time period 2000-2003. There is, however, no attempt to establish a causal relationship between these two phenomena.

Wholey, Feldman and Christianson (1995) examine the effects of HMO market structure (measured by the number of HMOs) on HMO premiums from 1988 to 1991. Their analysis uses the HMO (which may be national, regional, or local) as the unit of observation. Premiums are estimated as average premium revenue per member, and the market structure facing each HMO is a weighted average of the number of competitors in the geographic markets in which the HMO is active. The results suggest premiums decrease when entry occurs. However, the specifications do not include HMO fixed effects, so the results are subject to the usual biases arising from cross-sectional sources of identification.

Key to our study design is a unique, proprietary dataset containing detailed information on the healthplans of roughly 10 million Americans in every year from 1998 to 2006, inclusive. This dataset affords us a number of advantages over other studies of the industry. It includes the actual premium charged to every sampled employer for each healthplan they offer. Several details are available for each healthplan as well, including the identity of the insurance carrier, the plan type, and a summary measure of enrollee demographics. The micro-level data enables us to avoid the noise and error associated with high-level aggregation. We also make use of geographic market definitions supplied by the industry, as opposed to arbitrary geographic units that may correspond poorly to actual markets. Finally, the panel nature of the dataset permits us to eliminate cross-sectional differences across markets and employers as a source of identification for the relationship of interest.

Our research complements recent work by Dafny (2008). Using the same dataset employed here, Dafny evaluates whether health insurance markets are competitive by examining whether health insurers engage in "direct" price discrimination, charging higher premiums to firms with deeper pockets, as measured by operating profits. She finds they do, and this result is not driven by cross-sectional differences across firms or plans: firms with positive profit shocks subsequently face higher premium growth, even for the same healthplans. Moreover, this relationship is strongest in geographic markets served by a small number of insurance carriers. This evidence of price discrimination by health insurers implies insurers possess and exercise market power in some local markets. Here, we focus on whether insurers use their market power to raise premiums *overall*, and by how much.

II. Data

Our primary source is the Large Employer Health Insurance Dataset (LEHID). LEHID contains information on all of the healthplans offered by a large and non-random sample of employers between 1998 and 2006, inclusive. Descriptive statistics for each year of data are presented in **Table 1.** LEHID is gathered and maintained by a leading benefits consulting firm, and the employers included in the dataset have some past or present affiliation with the firm. The unit of observation is the healthplan-year. A healthplan is defined as a unique combination of employer, market, insurance type, insurance carrier, and plantype, e.g. Company X's Chicago-area fully-insured Aetna HMO. We now discuss each of the components that jointly identify this unit of observation in turn.

The full dataset includes observations from 813 *employers*. Employers may enter or exit the sample at any time. The median number of years an employer is present in the sample is two. One-quarter of employers appear in the sample for 4 or more years. A non-trivial number of employers reappear after exiting. Most employers are large, multi-site, publicly-traded firms, such as those included on the *Fortune 1000* list. The leading industries represented include manufacturing (110 employers), finance (101), and consumer products (73), although nonprofit and government sectors are also represented (43 in the "government/education" category).

Geographic *markets* are defined by the source using 3-digit zipcodes. The 139 markets reflect the geographic boundaries used by insurance carriers when quoting prices, and collectively cover all of the continental U.S., with the exception of a few rural areas. Large metropolitan areas are separate markets, and non-metropolitan areas are lumped together within state boundaries, e.g. "New Mexico – Albuquerque" and "New Mexico – except Albuquerque."[7] To match county-level data to these markets, we allocate zipcodes within the markets to counties, and use zipcode population data to weight the county data appropriately when aggregating to the market level. The two county-year measures we use are the unemployment rate (from the Bureau of Labor Statistcs), and the average Medicare costs per capita (known as the AAPCC, from the Center for Medicare and Medicaid Services). We also calculate the general, acute-care hospital HHI at the market-year level using hospital-year data on the number of beds for all general hospitals included in the American Hospital Association Annual Surveys of Hospitals. To create this measure, we assign hospitals with the same "system ID" to a common owner.

[7] There is only one market that crosses state boundaries, "Massachusetts – Southern and Rhode Island." A map of the markets is available in Dafny (2009).

The sample includes both *fully-insured* and *self-insured* plans. As these terms suggest, only the former is "classical" insurance in which the insured pays the carrier to bear the risk of realized healthcare outlays. 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 may purchase stop-loss insurance to limit their remaining exposure. Per ERISA (the Employee Retirement Act of 1974), these plans are also exempt from state regulations (such as specific benefit mandates) and state insurance premium taxes. In our sample, the fraction of plans that are fully-insured declines from 45 to 20 percent between 1998 and 2006. The decline is somewhat less precipitous when calculated using the fraction of enrollees – 42 to 25 percent – but clearly remains an important phenomenon in the data. The reasons for this decline are the subject of a current research project. Here we note the decline is not particular to our data source: it has been corroborated in the Kaiser Family Foundation/Health Retirement Education Trust Annual Survey of Employer Benefits and the Medical Expenditure Panel Survey-Insurance Component (MEPS-IC), and appears to be especially pronounced among the very largest firms.[8] The substitutability of fully-insured and self-insured plans (at least for employers within our sample) is one reason we aggregate the data to the employer-market-year level. In Section 5 ("Extensions"), we also examine the impact of market concentration on the decision to offer fully-insured plans.

Each firm that administers any plan in the data is labeled an "insurance *carrier*."[9] During the entire study period, there are 357 carriers that serve at least one employer, and 195 that serve 5 or more. The smaller carriers tend to be local or regional firms, or sometimes "third party administrators" who pay claims and contract with another firm to "rent" its network of providers and associated discounts. The industry is highly concentrated and becoming more so over time. **Figure 1** presents the four-firm concentration ratio for the nation as a whole, estimated using the LEHID sample. This measure increased from an impressive 58 percent in 1998 to 79 percent in 2006. Concentration ratios within local markets - arguably where most of the competition takes place - are much higher.[10]

[8] We are grateful to Kosali Simon for tabulating the MEPS-IC data to investigate this trend.
[9] Blue Cross and Blue Shield (BC/BS) affiliates are all assigned the same carrier ID. (Note: both Wellpoint and Anthem (before it was acquired by Wellpoint) own BC/BS affiliates, so they also have the BC/BS carrier ID. Given we calculate concentration within each market, and there are only a handful of markets in which BC/BS affiliates complete, the uniform coding of these affiliates is unlikely to be consequential for our analysis.
[10] The notable exception is the market for multisite employers interested in a uniform plan across all sites. Our data do not include an identifier for jointly-negotiated plans.

The *plan types*, ordered from most to least restrictive in terms of provider choice, are Health Maintenance Organization (HMO), Point of Service (POS), Preferred Provider Organization (PPO), and Indemnity. HMOs and POS plans control utilization of care through primary-care physicians ("gatekeepers"). Only in-network providers are covered by HMOs, while POS plans provide some coverage for out-of-network providers (once the gatekeeper has approved the service in question). PPOs engage in less utilization management, and like POS plans, typically cover out-of-network care at a reduced rate. Finally, indemnity plans are traditional fee-for-service arrangements in which benefits do not depend on the network status of the provider. As Table 1 reveals, the composition of plan types fluctuated during the study period, with a clear resurgence of PPOs toward the end of the study period.

In addition to the elements that jointly define a plan, we have the following variables: *premium, demographic factor, plan design factor*, and *number of enrollees*. *Premium* is expressed as an average amount per enrollee (i.e. a covered employee); it therefore increases with the average family size of enrollees in a given plan. Premium combines employer and employee contributions, and for self-insured plans it is a projection of expected costs per enrollee (including estimated administrative fees paid to an insurance carrier). These projections may include a partial risk premium if the employer purchases stop-loss coverage; whether stop-loss coverage is purchased is not captured in the data. Because the forecasts are used for budgeting and to establish employee premium contributions, they are carefully developed and vetted. Employers often hire outside actuaries and benefits experts (such as our source) to assist in formulating accurate projections.

Demographic factor is a measure that reflects family size, age, and gender composition of enrollees in a given plan. *Plan design factor* captures the generosity of benefits within a particular carrier-plan type, with an emphasis on the degree of coinsurance and copays. Both factors are calculated by the source, and the formulae were not disclosed. The *number of enrollees* in each plan refers to the number of enrolled employees, i.e. it does not reflect dependents. The total number of enrollees in all LEHID plans averages 4.7 million per year. Given an average family size above 2, this implies over 10 million Americans are represented in the sample in a typical year.

As noted above, we perform most analyses using data aggregated to the employer-market-year level. **Table 2** presents descriptive statistics for this unit of observation. Because our primary outcome is *growth* in health insurance premiums (in order to avoid cross-sectional identification of the coefficients of interest), aggregating the data to the employer-market-year level enables us to use a much larger proportion of the data. With the healthplan-level data, growth in premium is undefined when an employer terminates a particular plan. Analogously, new plans can only enter into the analysis after multiple observations are available. Changes to plan offerings are quite common in our data. Moreover, changes in market concentration may affect both the likelihood that an employer switches to a different insurer as well as the type of insurer (e.g. low or high cost) and plan type that the employer selects, so we do not want *a priori* to eliminate this substitution from our sample.[11]

Before proceeding to the analyses, we evaluate the representativeness of the LEHID data. The best source for nationally-representative estimates of employer-sponsored health insurance premiums is the annual Employer Health Benefits Survey, sponsored jointly by the Kaiser Family Foundation (KFF) and the Health Research and Educational Trust (HRET).[12] Using these data, KFF/HRET report the average growth in premiums for a family of four. Although we would not expect premium *levels* to be similar for this sample and the LEHID sample (both because the selection of firms is nonrandom and because family sizes differ across plans), if growth rates are similar this would suggest the results of our study are applicable to a broader sample of employers because all specifications rely on premium growth over time.

Appendix Figure 1 graphs the annual growth rate for employee-weighted premiums against that reported by KFF/HRET. The trends in both samples are very similar over time. Dafny (2008) also reports that the ratio of sampled enrollees to total insured lives (available at the county-level from the US Census of 2000) varies little across geographic markets. In the appendix, we describe our efforts to compare the LEHID-based estimates of market structure

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[11] As an example of the frequency with which this occurs, consider employer-market pairs that are present in both 1999 (the year of the Aetna-Prudential merger) and 2002. More than half of the plans offered by these firms in 1999 are no longer present in 2002, either because the employer switched to different carriers or because it changed the type of plan with the same carrier.

[12] 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.

with those obtained by other researchers using the proprietary InterStudy database, specifically Scanlon, Chernew, and Lee (2006). Scanlon et al (2008) use these data to show that increased levels of HMO competition do not lead to increases in plan quality. Interstudy reports some enrollment and premium figures at the insurer and MSA level, but for reasons outlined in the Appendix, it is not an ideal source for our purposes.

III. Are Increases in Local Market Concentration Associated with Increases in Premiums?

In this section, we examine the relationship between premium growth and changes in local market concentration. We begin by describing the distribution of market-level HHI and how this has changed over time. Next, we estimate OLS regressions relating premium growth at the employer-market level to changes in the corresponding market HHI. These results reveal whether premiums are growing more (or less) quickly in markets that are becoming more concentrated. By estimating our models in long differences, we control for unobservable, time-invariant differences across employer-market cells (e.g. a high average premium for employees in the San Francisco office of Firm Y). The richness of the data also permits us to control for important time-varying differences (such as the percent of enrollees in HMOs and the degree of copays). Although interesting as a descriptive exercise, this analysis does not yield estimates of the *impact* of changes in market structure on premium growth, as observed changes in market structure are unlikely to be exogenous. In Section IV, we estimate this causal relationship by using the Aetna-Prudential merger to instrument for changes in market concentration.

A. Market Structure of the Group Insurance Market, 1998-2006

During the 9-year study period, the average market-level HHI (estimated using our sample, on a scale from 0 to 10,000) increased from 2,286 to 2,984. Using the categorization from the Horizontal Merger Guidelines, the fraction of markets falling into the top "highly concentrated" category (HHI > 1,800) rose from 68 to 99 percent. As illustrated by the histograms presented in **Figure 2**, the biggest increases occurred during the second half of the study period, but sizeable increases are present in the first half as well. Between 1998 and 2002, 53 percent of markets experienced increases in HHI of 100 points or more, and 25 percent saw increases of 500+ points. The corresponding figures for 2002 to 2006 are 78 and 53 percent, respectively. The Merger Guidelines provide a helpful frame of reference for interpreting these changes. According to the Guidelines, mergers resulting in an increase of 100+ points are "presumed…likely to create or enhance market power or facilitate its exercise."[13] Importantly, there is wide variation in the magnitude of changes in HHI across markets, notwithstanding the fact that most are positive.

The reasons for these changes in HHI (apart from sample composition, which we discuss below), can be subdivided into "structural" (related to entry, exit, and consolidation of insurance carriers) and "non-structural" sources. Using data on fully-insured HMOs only, Scanlon et al (2006) report that 61 to 65 percent of the variation in HHI between 1998 and 2002 is attributable to changes in market structure. Structural changes (primarily due to consolidation or exit) are also important in our sample: the mean number of carriers per market declined from 18.9 in 1998 to 9.6 in 2006. (As the data on HHI suggests, many of these carriers are quite small. This is due to the presence of many small self-insured plan administrators, particularly in the earlier part of the study period. Some of these administrators may not be active participants in a given market, i.e. they "rent networks" from other carriers so as to offer a particular client a consistent plan across all geographies.) **Figure 3** contains histograms for changes in the number of carriers. Whereas in 1998-2002 the modal net loss is 1 to 3 carriers, between 2002 and 2006 the modal net loss is 4 to 6 carriers. Of course, neither structural nor non-structural sources of changes in HHI can be presumed exogenous to other determinants of premiums.

B. OLS Estimates of the Relationship between Market Structure and Premiums

In our baseline OLS model, we regress premium growth at the employer-market level on changes in HHI at the market level, controlling for changes in other market-level covariates and employer-market demographics:

$$\textit{(1)}\quad \Delta \ln premium_{em} = \alpha + \gamma \Delta lagged\ HHI_m + \Delta lagged\ X_m \beta + \varphi \Delta demographics_{em} + \varsigma_e$$
$$\left[+\, \omega \Delta plan\ type\ shares_{em} + \vartheta \Delta plan\ design_{em}\right] + \varepsilon_{em}$$

[13] *Horizontal Merger Guidelines,* Federal Trade Commission and Department of Justice, issued in 1992 and revised in 1997. Accessed at http://www.usdoj.gov/atr/public/guidelines/hmg.htm.

The subscripts *e* and *m* refer to the employer and market, respectively. Because premiums are set prospectively (e.g. premiums for 2006 are determined in 2005), we lag HHI and the other market-level covariates (denoted X_m) by one year before taking differences. These covariates include the unemployment rate (to capture local economic conditions), the log of per-capita Medicare costs (to capture trends in healthcare utilization), and the general, acute-care hospital Herfindahl index (to capture concentration in the provider market, which could independently lead to premium increases). Differencing the data eliminates cross-sectional variation in market concentration as a source of identification for the coefficient of interest, and also reduces the possibility that any results are affected by changes in the composition of employers in the sample over time. We report standard errors clustered by market.

We begin by considering the entire study period (corresponding to 1999-2006 given the lag in HHI) and subsequently subdivide this into the earlier (1999 to 2002) and later (2002 to 2006) segments. Because the panel is unbalanced, the number of observations differs by study period; results are similar if we restrict attention to employer-markets present in all years. In all specifications, HHI is measured on a scale from 0 to 1. Results for each period are presented in panels A, B, and C of Table 3, respectively. Standard errors are clustered by market. The first column in each panel corresponds to the baseline specification (i.e. without terms in brackets). Columns 2 and 3 add controls for changes in the generosity of plans, namely the change in the percent of enrollees in each plan type (excluding POS, the omitted category) and the change in *plan design*. Relaxing constraints on provider choice and utilization (i.e. moving toward PPOs) should be associated with higher premiums. Increases in plan design should also result in higher premiums. Because substitution across plan types and modifications to plan design may constitute a response to changes in HHI, controlling for these terms is akin to using a Laspeyres price index as a dependent variable, i.e. using the change in price for a given product type and design.

All specifications include employer fixed effects (denoted by ς_e in equation 1). These terms control for differences in average premium growth for different employers. To the extent employers with particularly high or low premium growth are systematically located in markets with particularly large or small changes in HHI, omitting these terms will bias our estimate of the relationship between premiums and HHI.

Before turning to the results, we elaborate on the shortcomings of this analysis. Most important is the inability to ascribe causality to the coefficient of interest. Changes in HHI are likely to be determined in part by expected changes in premiums. For example, exit by carriers (and hence increases in HHI) may be more likely in markets where premiums are expected to grow most slowly. Non-structural changes in HHI may also generate a downward bias in the HHI coefficient, e.g. if HHI increases because employers in markets with dim economic prospects substitute toward a low-priced "Walmart-style" carrier. Indeed, most plausible sources of endogeneity suggest the OLS coefficient will be downward-biased. A second shortcoming of this analysis stems from measurement error in HHI. Although the sample size is large, particularly as a share of large-firm employment in the U.S., it is non-random and estimated market shares will be noisy. Any systematic, time-invariant error (e.g. due to the absence of small carriers in the sample) should be eliminated by differencing the data, but the coefficient of interest will still suffer from attenuation bias. For these two reasons, we expect the OLS estimates to understate the actual impact of market concentration on premiums.

The OLS estimates offer no support for the hypothesis that increasing consolidation of health insurance markets is associated with higher premiums. Across all study periods and specifications, the point estimates on changes in HHI are negative, small in magnitude and imprecisely estimated. The results in Panel A imply the mean market-level increase in HHI of 698 points between 1999 and 2006 (corresponding to .0698 given the scale used in our estimation) is associated with premium decreases of 0.2 to 0.4 percent. As discussed above, all of the OLS coefficients are likely downward-biased. Nevertheless, comparing the estimates across models provides some insights.

First, we note that adding controls for changes in plan generosity (as measured by the percent of enrollees by plan type and the average plan design factor) generally reduces the coefficient of interest. This implies that employers in consolidating markets are shifting toward cheaper plans. The strongest evidence of such "benefit buybacks" appears in Panel A, where adding controls for the percent of enrollees in each plan type halves the HHI coefficient and renders it statistically indistinguishable from zero.

For the most part, the coefficient estimates on the market-level control variables are statistically insignificant. The coefficient estimates on the employer-market controls indicate demographic factor and plan type are particularly important predictors of premiums. As expected, increasing values of the former are associated with higher premiums, and increasing management of care is associated with lower premiums. Finally, increases in plan design are associated with higher premiums, except in the 1999-2002 study period.

C. *Robustness*

As a robustness check, we also re-estimated our specification of interest in levels. These models utilize the entire sample of employer-market-year observations from 1999-2006 and regress ln(premium) on lagged HHI, lagged market covariates, employer-market-year covariates such as demographic factor and plan design, employer-market fixed effects, employer-year fixed effects[14] and year fixed effects. As with the long-difference models, this specification also exploits within-employer-market variation in premiums, but employers need only be present in any two years to help identify the coefficients of interest. The results were very similar: employers in consolidating markets do not end up paying higher increases in premiums.[15]

IV. **Do Increases in Local Market Concentration *Cause* Increases in Premiums?**

In this section, we attempt to estimate the causal effect of changes in market concentration on premium growth by exploiting shocks to market concentration produced by mergers and acquisitions (M&A). Because M&A activity in local or regional markets may itself be motivated by expected trends in premium growth, we considered only large, non-local mergers as candidates for this analysis. We also ruled out mergers with insufficient pre or post periods (e.g. Aetna and NYLCare in 1998), few overlapping markets, or very small shares in our sample for one of the merging parties (e.g. United Healthcare and MAMSI). Only one merger

[14] The employer-year fixed effects control for differences in premium growth rates across employers, as is effectively done by including employer fixed effects in the long-difference models.
[15] Results available upon request.

remained: the Aetna-Prudential merger of 1999. Post-merger, the new firm (known as "Aetna") was widely reported to be the national's largest insurer, covering 21 million individuals.[16] Importantly, and as we describe in detail below, there was substantial overlap in the local market participation of Aetna and Prudential prior to the merger, generating the potential for sizeable post-merger changes in local market concentration.

Our analysis is subdivided into four sections. First, we discuss the context for the merger, paying special attention to whether the timing was affected by anticipated, market-specific changes in premium growth trends. Second, we estimate the impact of the merger on market concentration (the "first stage" analysis). In so doing, we document the range of pre-merger market shares for Aetna and Prudential, as well as the degree of pre-merger overlap. Third, we perform a reduced-form analysis, in which we examine the impact of the merger on premium growth. Fourth, we combine these analyses to produce our estimate of the causal impact of concentration on premiums.

A. The Aetna-Prudential Merger of 1999

In December 1998 Aetna Inc. announced its intention to purchase Prudential Health Care (hereafter Prudential) for $1 billion. Prudential had been publicly searching for an acquirer since at least October of the year prior; it was widely reported to be losing money and its parent firm, Prudential Insurance Company of America, had decided to exit the health insurance business. Importantly, Aetna was an unlikely suitor, as it had recently closed another $1 billion acquisition (of NYLCare), and had publicly stated that future acquisitions would not occur "for at least a year."[17] In announcing the deal, Aetna's CEO claimed Prudential had 'made an offer we can't refuse.'[18] The deal closed in July 1999, after Aetna signed a consent decree to address concerns raised by the Department of Justice (DOJ). DOJ alleged that after the merger, Aetna would have a market share for fully-insured HMOs of 63 percent (in Houston), and 42 percent (in Dallas). As a precondition to approve the merger, it required the divestment of all Houston and Dallas-area plans Aetna had acquired in the 1998 NYLCare purchase.

[16] Sanders, Alain L., "Will the Aetna-Prudential Merger Hurt the Patient?" *TIME* magazine, June 22, 1999.
[17] Freudenheim, Milt, "Aetna to Buy Prudential's Health Care Business for $1 Billion," *The New York Times*, December 11, 1998, Section C, page 1.
[18] *Ibid*

According to industry analysts, Aetna's acquisition of Prudential was part of a strategic bet on the long-term viability of managed care. Originally focused on providing fee-for-service plans to large, self-insured employers, Aetna gambled on the rising popularity of HMOs with the 1996 purchase of U.S. Healthcare, which offered fully-insured HMOs to small groups. The acquisitions of NYLCare (New York Life's healthcare unit) and Prudential soon followed; managed plans were also the dominant segment for these units. At its peak (after the Prudential acquisition in 1999), the firm covered 21 million lives. However, enrollment fell rapidly thereafter, plateauing at 13 million in 2002.[19] According to a 2004 *Health Affairs* article by James Robinson, "Aetna was the poster child for the aspirations and failures of managed care, channeling patients and physicians into HMOs; holding down premiums so that enrollment would grow; acquiring competitors to penetrate new markets; and then floundering in adverse publicity, economic shortfalls, and investor disenchantment."

Given this history, the Aetna-Prudential merger does not appear to raise ex ante concerns about endogeneity. We corroborate this conjecture empirically below, by examining whether premium growth in the pre-merger period was systematically different in markets where both firms had significant pre-merger overlap.

B. The Effect of the Aetna-Prudential Merger on Market Concentration

In our sample from 1999, Aetna and Prudential were the third and fifth largest insurers in terms of the number of enrollees. All 139 markets included plans offered by both firms. There was significant variation across markets, however, in the pre-merger shares of each firm. We hypothesize that markets served by both firms experienced increases in market concentration immediately following the merger of Aetna and Prudential, and that these increases varied by the pre-merger shares of the two firms. Specifically, for every market we calculate "simulated HHI change" ($sim\,\Delta HHI_m$) as follows:

[19] This history, together with the enrollment figures, is summarized in "From Managed Care to Consumer Health Insurance: The Fall and Rise of Aetna," (Robinson 2004).

$$(2) \quad sim\,\Delta HHI_m = \left[Aetna\ 1999\ share_m + Pru\ 1999\ share_m\right]^2 - \left[\left(Aetna\ 1999\ share_m\right)^2 + \left(Pru\ 1999\ share_m\right)^2\right]$$
$$= 2 * Aetna\ 1999\ share_m * Pru\ 1999\ share_m$$

$sim\,\Delta HHI_m$ represents the merger-induced increase in market *m*'s HHI that would have occurred from 1999 to 2000 absent any other changes in carriers' market shares. For example, if Aetna and Prudential were two of four firms with equal market share in 1999, $sim\,\Delta HHI_m$ would equal $0.125 = (0.5)^2 - ((0.25)^2 + (0.25)^2)$ or 2*0.25*0.25. **Figure 4** provides detail on the actual distribution of $sim\,\Delta HHI_m$.

We propose to use $sim\,\Delta HHI_m$ as an instrument for ΔHHI_m. In this section, we present first-stage models to confirm that $sim\,\Delta HHI_m$ is indeed correlated with ΔHHI_m. We also take two approaches to confirm the merger is orthogonal to other determinants of HHI. First, we investigate whether HHI is trending differently just prior to the merger in those markets predicted to be most affected by the merger. Second, we investigate whether the relationship between ΔHHI_{mt} and $sim\,\Delta HHI_m$ is severed in Texas, where the DOJ attempted to mitigate the impact of the merger by requiring the divestiture of NYLCare plans.

Because HHI and $sim\,\Delta HHI_m$ vary at the market-year level, we perform the analyses in this section using the market-year as the unit of observation. We begin by estimating the following specification on data from all non-Texas markets:

$$(3) \quad \text{HHI}_{mt} = \alpha + \lambda_m + \tau_t + \beta sim\,\Delta HHI_m * \tau_t + X_{mt} + \varepsilon_{\text{mt}}$$

The vectors denoted by λ_m and τ_t represent a full set of market and year fixed effects, respectively. As in earlier specifications, X_{mt} is the vector of market-year covariates (log of per capita Medicare costs, unemployment rate, and hospital market HHI). By interacting $sim\,\Delta HHI_m$ with separate dummies for each year (except 1998, the omitted category), this model investigates the possibility that trends in market concentration may have been different prior to the merger in markets differentially impacted by the merger. The merger was effectively cleared in July 1999, when the Department of Justice submitted its Proposed Final Judgment. Given insurance premiums are set a few months prior to the start of the calendar year, the impact of the merger should become apparent in 2000 or later.

Figure 4 graphs the coefficient estimates on the yearly interactions with $sim\,\Delta HHI_m$, together with the 95% confidence intervals. The sample includes data from 1998 to 2003. Estimates are presented in numerical form in Appendix Table 1. Relative to the omitted interaction term, $sim\,\Delta HHI_m * (year == 1998)$, only the interactions with indicators for 2000 and 2001 are statistically significant. The coefficient estimate for β in 1999 is small and negative (-0.10), whereas estimates for β in 2000 and 2001 are large (0.50 and 0.47, respectively) and significant at $p<0.05$. Notably, these coefficients are significantly smaller than 1, suggesting employers substituted away from Aetna and Prudential in the wake of the merger. This is the first suggestive evidence that the merged firm may have been pricing in an uncompetitive fashion post-merger.

The coefficient estimates on β in 2002 are 2003 are both noisy and negative. These estimates reveal that the effect of the merger on market concentration declined sharply after 2001. This finding is consistent with reports from industry experts. According to a 2004 *Health Affairs* article by James Robinson, "[G]ossip speculates [Aetna] would be lucky to still have 30,000 of the 5 million it acquired from Prudential." Robinson claims NYLCare (acquired by Aetna in 1998) and Prudential (acquired by Aetna in 1999) were restraining their premium increases to maximize membership in anticipation of a sale. "Once Aetna reset premiums on NYLCare and Pru accounts at sustainable levels," he writes, "most former customers simply went elsewhere." We revisit these observations in Section 4. None of the market-year covariates explains the variation in HHI over time.

Given the results in Figure 4, we focus our attention on the period from 1998-2001. Table 4 reports results of specifications that take the following form:

$$\text{(4)}\quad \begin{aligned} HHI_{mt} = {} & \alpha + \lambda_m + \tau_t + \beta_0 sim\Delta HHI_m * post_t \\ & \left[+ \beta_1 sim\Delta HHI_m * post_t * Texas_m + \psi post_t * Texas_m\right] \\ & + X_{mt}\vartheta + \varepsilon_{mt} \end{aligned}$$

A single interaction between $sim\Delta HHI_m$ and a "post" dummy replaces the individual year interactions in this parsimonious specification. After estimating the baseline model (which excludes the terms in brackets), we add the Texas markets to the sample and include a triple-interaction, $sim\Delta HHI_m * post_t * Texas_m$, to explore whether the post-merger impact of $sim\Delta HHI$ differs in these markets. We also add the term $post_t * Texas_m$ to control for average changes in Texas as compared to other states during the post-period. As in prior specifications, we include market-year covariates.

The estimates of the baseline model, presented in Column 1, confirm the conclusion reached earlier: markets predicted to be most affected by the Aetna-Prudential merger did indeed experience significantly greater increases in market concentration during the first 2 post-merger years. The coefficient $sim\Delta HHI_m * post_t$ is 0.50, with a standard error of 0.14. The results in Column 2 reveal that the federal government achieved its objective of neutralizing the merger's effect on market concentration in Texas markets: the triple-interaction term for Texas markets is negative and statistically-significant. The estimated magnitude (-1.2) exceeds that needed to offset the impact of the merger, although we cannot reject the hypothesis that the sum of the double and triple-interaction terms is equal to zero. (Note the coefficient on $post_t * Texas_m$ is small and noisily estimated.) Observations from Texas may therefore constitute a useful comparison group for our later analyses involving health insurance premiums and related outcomes of interest.

Taken together, the results in this section demonstrate that the merger of Aetna and Prudential, two of the nation's largest health insurers, resulted in substantial increases in market concentration in markets differentially served by both firms. The effect of the merger dissipated quickly, with no lingering effect on market concentration by 2002. We also find no effect of the merger on concentration within Texas markets, where the DOJ consent decree more than offset the predicted effects of the merger on market structure.

C. The Effect of the Aetna-Prudential Merger on Health Insurance Premiums

To investigate the effect of the merger-induced increases in local market concentration on plan premiums, we estimate models of the following form:

$$
\begin{aligned}
(5) \qquad \Delta \ln(\mathit{premium})_{emt} = \alpha + \kappa_1 \mathit{sim}\Delta \mathit{HHI}_m * \mathit{post}_t + \\
\left[+ \kappa_2 \mathit{sim}\Delta \mathit{HHI}_m * \mathit{post}_t * \mathit{Texas}_m\right] \\
+ \Delta \mathit{demographics}_{emt} + \Delta X_{mt\text{-}1}\vartheta + \lambda_m + \tau_t + \varsigma_e \\
\{+ \omega\Delta \mathit{plan\ type\ shares}_{emt} + \vartheta\Delta \mathit{plan\ design}_{emt}\} + \varepsilon_{emt}
\end{aligned}
$$

To maximize the number of observations in our sample, we estimate the model in first differences, i.e. using annual changes in all variables (except $\mathit{sim}\,\Delta \mathit{HHI}_m$, which is determined using 1999 market shares). In light of the results from the preceding section, we focus on the period between 1998 and 2002 (i.e. 1998-99, 1999-2000, 2000-01, 2001-02). Note that in this model post_t takes a value of one for 2000-2001 and 2001-2002 changes. When we expand the sample to include observations from Texas, we add the triple-interaction term in square brackets. Standard errors for all models are clustered by market to allow for correlation in the error terms among employers in the same market.

As in the premium regressions presented in Section III, we begin with a parsimonious specification that controls for changes in employer demographics, lagged market covariates, and fixed differences across employers and markets in average annual premium *growth* (captured respectively by employer and market fixed effects, denoted ς_e and λ_m). The results are reported in Column 1 of **Table 5**, under the subtitle "Reduced Form Estimates." The coefficient estimate on $\mathit{sim}\Delta \mathit{HHI}_m * \mathit{post}_t$ is positive and statistically significant. Given the mean $\mathit{sim}\,\Delta \mathit{HHI}_m$ of 0.014 (across all 139 geographic markets), the point estimate of 0.169 implies that, in a typical market, the merger induced market-wide premium increases of 0.24 percent. Columns 2 and 3 add controls for changes in the generosity of plans, namely the annual change in the percent of enrollees in each plan type (excluding POS, the omitted category) and the annual change in *plan design*. We do not find evidence of "benefit buybacks" in the wake of the merger – the coefficient on $\mathit{sim}\,\Delta \mathit{HHI}_m * \mathit{post}_t$ actually declines somewhat when these controls are included.

The final column in the "Reduced Form" row in Table 5 presents the results of the falsification test enabled by the divestiture requirement in Texas. To execute this test, we add Texas observations to the sample and estimate the full model (as in Column 3) with the addition of a triple interaction term, $sim\Delta HHI_m * post_t * Texas_m$.[20] The estimated coefficient on this term will reveal whether the post-merger impact of $sim\Delta HHI$ differs in Texas markets. In fact, it is highly significant and negative (-0.132), and almost perfectly offsets the main effect of $sim\,\Delta HHI_m$ in this specification (0.146). An F-test confirms the sum of these terms cannot be statistically distinguished from zero. Thus, the market power effect of the merger in Texas was indeed neutralized by the DOJ's actions.[21]

D. IV Estimates

In order to obtain an IV estimate of the effect of changes in the local market HHI on growth in premiums, we estimate specifications similar to the reduced form specifications described above, but using $sim\,\Delta HHI_m$ as an instrument for lagged HHI. The results from these models are presented in Table 5 in the row labeled "IV Estimates". The coefficient on lagged HHI in the base model is positive, statistically significant, and roughly twice as large as the reduced form estimate. This is anticipated given the coefficient of 0.5 reported in Table 4.[22] On adding controls for changes in plan generosity, the coefficient remains positive and statistically significant with only a small decline in magnitude. Applying the estimate from the full model in column 3 to the mean market-level increase in HHI of .0698 over the period 1998-2006, we predict a premium increase of exp(.021) = 2.15 percent due to changes in market concentration.

For the sake of comparison, we also present coefficient estimates obtained using OLS models, in which lagged HHI is the predictor of interest. As noted before, OLS estimates are likely to be downward-biased, understating the actual impact of changes in market concentration on premiums. Indeed, the coefficients from the OLS models are close to zero, and Hausman

[20] Note a second-order interaction (i.e.$post_t$*$Texas_m$) is not appropriate in this model as state fixed effects already control for differences in annual growth rates across states. Given the short time period, the coefficient on such a term would be difficult to separately identify from the triple interaction term.
[21] As an additional (and separate) extension of the reduced-form analysis, we examined whether the impact of the merger was greater in markets with higher initial levels of concentration. Unfortunately, coefficient estimates on $sim\Delta HHI_m * post_t *$initial HHI_m (and variants thereof) were very imprecise.
[22] The IV estimate is not exactly equal to the ratio of the reduced-form estimate to the "first stage estimate" reported in Table 4, as Table 4 uses the market-year as the unit of observation.

specification tests reject the null assumption of consistency for these models (with p-values between .01 and .02), underscoring the need for instrumental variables estimation.[23]

Collectively, the results presented in this section show that consolidation does result in "premiums on premiums." We arrive at this conclusion by exploiting arguably exogenous increases in local market concentration caused by the nationwide merger between two large insurance firms, Aetna and Prudential. We show that the merger led to large and varying increases in HHI across local health insurance markets which in turn resulted in higher insurance premiums. Two key results indicate this finding is not driven by unobserved factors correlated with the pre-merger market share of Aetna and Prudential. First, there is no evidence that premiums in markets with higher $sim\,\Delta HHI$ were trending differently before the merger took effect. Second, we find the opposite response in Texas, where the merger was effectively blocked by the Department of Justice. These tests support the use of $sim\,\Delta HHI_m$ as an instrument for $lagged\ HHI_m$.

V. Extensions

In this section, we assess the impact of insurer consolidation on healthplan characteristics other than price. We begin by looking at the effect of the merger on *plan design*. We regress annual changes in *plan design* on the same independent variables as in the baseline reduced-form model presented in equation (5) (corresponding to Column 1 of Table 5). The results are presented in Table 6. For parsimony, all models in Table 6 are estimated on the sample including Texas (and the concomitant interaction term).[24] We find that employers reduce the generosity of benefits in the wake of the Aetna-Prudential merger, and this effect is almost perfectly offset in Texas markets. Thus, increasing consolidation not only leads to higher prices, holding constant observable plan characteristics such as plan design (which was controlled for in the reduced-form specifications), but also to "benefit buybacks" as employers try to reduce the burden of higher insurance premiums.

[23] p-values for rejection of the null assumption of consistency of OLS models in Columns 1, 2 and 3 are.0065, .0252 and .0171, respectively.

[24] Results change little when Texas is excluded or additional controls added. Note that most of the dependent variables we consider appear as controls in equation (5), hence it is not possible to estimate the exact same set of specifications for each dependent variable in Table 6.

Columns 2 through 4 examine the impact of the merger on the share of employees enrolled in HMOs, PPOs, and Indemnity plans, respectively. We find employers in markets heavily impacted by the merger move away from HMOs and toward Indemnity plans. The estimated impact on PPO enrollment is positive but noisily estimated. Although we might have anticipated a shift toward *cheaper* plan types following a major consolidation, ceteris paribus, given the specifics of the merger in question these findings are unsurprising. The patterns suggest employers switching away from Aetna-Prudential (which was heavily pushing its HMO product) preferred to return to more traditional, unmanaged plans. We underscore that the potential impact of these changes on premiums is controlled for in most of our premium models (columns 2 through 4 of Table 5).[25] In the interest of space, we do not report results for two additional outcomes (percent of enrollees in POS plans and percent of enrollees in fully-insured plans), for which the coefficient estimates were not precise.

VI. Evaluating the Effects of Insurer Consolidation on Providers

Thus far, we have examined the impact of market structure in the insurance industry on downstream buyers, specifically of group plans. However, the extent of competition in the insurance industry will also affect upstream suppliers, such as healthcare providers, pharmaceutical firms, and medical device manufacturers. To the extent suppliers have few outside options, a lack of vigorous competition among insurers may lead to monopsonistic practices, i.e. insurers reducing the quantity of purchased inputs and acquiring these at a reduced "subcompetitive" price. In this section, we consider this possibility explicitly by estimating the impact of our HHI instrument ($sim\Delta HHI_m$) on the employment and compensation of healthcare personnel (such as physicians and nurses). As in the premium analysis, if variation in the impact of the merger on different geographic localities can be assumed exogenous to other determinants of employment and compensation trends, our results can be interpreted as causal estimates of the impact of consolidation on these outcome measures.

[25] We also considered the following dependent variables: change in percent of enrollees in self-insured plans, and z. WE find....

Concern about insurers' monopsonistic practices has emanated not only from provider organizations (such as the American Medical Association and the American Hospital Association) but also from state and federal regulatory authorities. As previously noted, the DOJ's challenge of the Aetna-Prudential merger in two Texas markets was based in part on concern over post-merger monopsony power. The formal complaint alleged the merger "would enable Aetna to exercise monopsony power against physicians, allowing Aetna to depress physicians' reimbursement rates in Houston and Dallas, likely leading to a reduction in quantity or degradation in quality of physicians' services".[26] More recently, as a precondition for the merger between UnitedHealth Group Inc. and Pacificare Health Systems Inc., the DOJ required the divestment of portions of Pacificare's commercial health insurance business in Tucson and Boulder in order to alleviate concerns about reduction in competition for physician services in those markets.[27]

A number of recent studies find evidence that insurer bargaining power depresses hospital prices (e.g. Feldman and Wholey 2001; Sorensen 2003; Shimazaki, Vogt and Gaynor 2008; and Ho 2009). Of these, only Feldman and Wholey explicitly consider the impact on quantity transacted, which should decline in the textbook monopsony case. They find HMOs' buying power (measured by the percentage of all hospital days in its enrollment area that the HMO reimbursed) is associated with lower hospital prices, but higher utilization of hospital services.[28]

Our analysis complements existing research by using a different subset of the provider industry (personnel rather than hospitals), and an identification strategy that mirrors the approach for estimating the causal impact of consolidation in the downstream premium market. We supplement the LEHID data with data from the Occupational Employment Statistics (OES) survey on income and employment in healthcare-related occupations. The OES survey is conducted semi-annually and provides estimates of employment and wages in over 800 occupations representing all full-time and part-time wage and salary workers in nonfarm

[26] See Complaint, U.S. vs. Aetna Inc. (ND TX, 21 June 1999)
[27] See Complaint, U.S. vs. UnitedHealth Group Inc. (20 Dec 2005)
[28] Other studies that focus on insurer-hospital bargaining include Brooks, Dor and Wong (1997), Town and Vistnes (2001) and Capps, Dranove and Satterthwaite (2003)

industries.[29] The survey description specifically notes that physicians are included in the survey, apart from the 15 percent who are self-employed. Approximately 200,000 establishments are surveyed every six months, and estimates are provided by geography (MSA) and by industry.

The OES data are organized by the North American Industry Classification System (NAICS), which groups establishments into industries based on the activity in which they are primarily engaged. We restrict attention to NAICS Sector 62 – Health Care and Social Assistance - and within this sector to occupations that are classified under the Standard Occupational Classification (SOC) system as "Healthcare Practitioner and Technical Occupations." These include 43 occupation categories such as dentists, registered nurses, anesthesiologists, surgeons, and pharmacy technicians. To facilitate a comparison of impacts on physicians versus nurses, we pool together the eight occupation categories referring to physicians and the two referring to nurses.[30]

The unit of observation for this data (as well as all analyses in this section) is the occupation-MSA-year and the variables of interest are the mean annual wage and estimated employment. Using a crosswalk that matches MSAs to LEHID markets, we merge this data with our measures of insurer concentration (including $sim\,\Delta HHI_m$). Table 7 provides annual summary statistics for the entire sample between 1999 and 2002, and separately for "Physicians" and "Nurses," as defined above. There is steady growth in average income over time for all occupation categories, with physicians experiencing a large jump between 2001 and 2002.[31] Nurses make up the largest employment category in the dataset by far, accounting for more than half of the estimated employment in healthcare-related occupations in all years.

[29] The employment and wage estimates for all occupations do not include the self-employed. The OES survey data is available online at <http://www.bls.gov/OES/>

[30] The categories pooled under "Physicians" are Dentists, Family and General Practitioners, General Internists, Obstetricians and Gynecologists, General Pediatricians, Psychiatrists, Podiatrists and Surgeons. Some of the individual physician categories have low estimates for employment and are present in only a handful of markets in both years. The "Nurses" category includes Registered Nurses and Licensed Practical Nurses.

[31] This is partly due to changes in the survey methodology between 2001 to 2002, when the Bureau of Labor Statistics revised the low-end of the highest wage range from "$60 and over" to "$70 and over" per hour. When we re-estimate our specifications using 1999 to 2001 as the study period, our conclusions remain unchanged.

We estimate parsimonious specifications using the change in log average earnings or employment between 1999 and 2002 as the dependent variable, and $sim\,\Delta HHI_m$ as our main predictor:

$$(6)\quad \Delta \ln y_{om,99-02} = \alpha + \gamma sim\,\Delta HHI_m + \omega Physician_o * sim\,\Delta HHI_m + \vartheta Nurse_o * sim\,\Delta HHI_m + \varsigma Physician_o + \theta Nurse_o + \upsilon \Delta HospitalHHI_m + \left[\Delta \ln y_{om,97-98} + \varsigma_o\right] + \varepsilon_{om}.$$

The subscripts *o* and *m* denote occupation and MSA, respectively. Our baseline specification includes indicators for the physician and nurse occupation categories as well as interactions between these indicators and $sim\,\Delta HHI_m$. The indicators capture differences in earnings and employment growth for each category (relative to other healthcare occupations) nationwide, while the interactions reflect the differential impact of insurer consolidation on earnings and employment in these categories. In all specifications, we control for the change in hospital concentration (as measured by the HHI) in each market. As specification checks, we progressively add each of the terms in brackets. The first term, $\Delta \ln y_{om,97-98}$, represents the change in earnings or employment between 1997 and 1998, and serves as a control for pre-existing trends in earnings (or employment) growth. The second term represents a full set of fixed effects for the 35 occupation categories. We necessarily restrict the sample to occupation-markets present in both 1999 and 2002, and we weight each observation by the average estimated employment in that occupation-market. Standard errors are robust and clustered by market.

The results are summarized in Table 8. Columns 1 through 3 pertain to models using the change in log average earnings from 1999-2002 as the dependent variable, while columns 4-6 use the change in log employment as the dependent variable. The coefficient estimate on $sim\,\Delta HHI_m$ in columns 1 through 3 is positive but imprecisely estimated, implying no significant impact of the merger on average earnings across all healthcare occupations. The coefficient on the physician indicator in columns 1 and 2 demonstrates that physicians experienced an increase of around 21 percent in average earnings between 1999 and 2002. However, the coefficient estimate on $Physician_o * sim\,\Delta HHI_m$ is negative and significant, revealing that earnings growth for physicians was lower in markets affected by the merger. Given the average value of .01 for $sim\,\Delta HHI_m$, the point estimate implies that the merger restrained growth in physician earnings by around 2 percent in a typical market. The coefficient on the nurse indicator reveals that nurses experienced a small decrease (around 1.5 percent) in earnings over the same time period. However, the interaction term for nurses is positive and statistically significant, implying this decrease was offset at least in part in markets where Aetna and Prudential had pre-merger overlap. Changes in hospital concentration do not appear to impact earnings growth of healthcare personnel, and the results are robust to the specification checks.

Columns 4 through 6 present estimates from specifications examining the impact of the merger on employment. The coefficients are again similar across all models. Relative to other healthcare occupations, employment of physicians increased, while that of nurses decreased, during the study period. The point estimate on $sim\,\Delta HHI_m$ is negative and significant: in a typical market, the merger led to a drop in healthcare-related employment of 2.7 percent. The interaction between the physician indicator and $sim\,\Delta HHI_m$ is negative but noisily estimated, whereas the interaction between the nurse indicator and $sim\,\Delta HHI_m$ is large, positive and significant. The smaller merger-induced decline in nurse employment implies there was some substitution toward nurses in markets impacted by the merger. This explanation is buttressed by the earnings regressions, which found the merger depressed growth in physicians' earnings while modestly boosting nurses' earnings.

In summary, we find that increases in market concentration predicted to occur in the wake of the Aetna-Prudential merger resulted in pronounced declines in healthcare-related employment. These declines were smaller for nurses than for other occupations on average (including physicians), and nurses also enjoyed wage increases relative to other occupations (and physicians in particular). The point estimates imply that post-merger market power facilitated a net reduction in payments to healthcare professionals. The results are consistent with the exercise of monopsony power by insurers vis a vis healthcare workers. We caution, however, that this conclusion is based upon the aftermath of one merger, albeit the largest merger to date (in terms of membership) and one with different impacts across the 139 geographic markets in the U.S. (implying 139 small experiments).

Paired with the results of the previous section, we conclude that in markets where Aetna and Prudential had substantial pre-merger overlap, insurers were able to exercise market power simultaneously in input *and* output markets post-merger. Thus, the premium increases documented in the previous section understate the increase in insurer profits due to consolidation.

VII. Discussion and Conclusions

Both the private and public sectors of the U.S. economy have struggled with soaring healthcare costs for the past few decades. The annual growth in private health insurance premiums has exceeded the annual growth in earnings in all but one of the last 20 years, and by a wide margin at that. In this study, we investigate whether and to what extent increasing consolidation in the U.S. health insurance industry is responsible for growth in employer-sponsored health insurance premiums over the past several years.

The scope of the private health insurance industry is difficult to overstate. Over 160 million non-elderly Americans are privately-insured, and this figure does not include publicly-insured individuals whose coverage is outsourced to private insurers (as is the case for the majority of Medicaid beneficiaries). In addition, most of the elderly purchase private supplemental insurance, also known as "Medigap" plans. Finally, most healthcare reform proposals would expand the reach of this $850 billion industry.

Our research focuses on employer-sponsored group health insurance plans, which during our study period (1998-2006) accounted for slightly less than 90 percent of the privately-insured non-elderly.[32] Our data includes the healthplan offerings, enrollment, and premiums for an unbalanced panel of 800+ large U.S. employers, and appears to be fairly representative of large employers nationwide. We include both fully-insured and self-insured plans in our analysis, as both options are viable for the firms in our sample.

[32] Source: EBRI Issue Brief, October 2007; Kaiser/HRET Survey of Employer Sponsored Health Benefits, 2007, Exhibit 10.1.

We arrive at four main conclusions. First, most Americans live in markets dominated by a small number of insurers, and most markets are becoming more concentrated over time. We estimate that the fraction of local markets falling under the "highly concentrated" category (per the Department of Justice's Horizontal Merger Guidelines) increased from 68 to 99 percent between 1998 and 2006. Second, premiums are *not* rising more quickly in markets experiencing the greatest increases in concentration, even controlling for a rich set of observable characteristics of plans. Third, when we account for the fact that changes in concentration are not orthogonal to other determinants of premium growth, we find that increases in concentration do raise premiums. Our instrumental variables estimates, which exploit plausibly exogenous shocks to local market structure generated by the 1999 merger of Aetna and Prudential, imply the average market-level change in HHI between 1998 and 2006 of 698 points produced a market-wide increase in premiums of 2.1 percent. Fourth, we find evidence that consolidation results in lower employment of healthcare workers, and facilitates the substitution of nurses for physicians. Using data from the Occupational Employment Statistics survey between 1999 and 2002, we find the Aetna-Prudential merger reduced physician earnings in a typical market by 2 percent and raised nurse earnings 0.4 percent. Employment of all healthcare workers in such a market declined by 2.4 percent on average, and of nurses by 0.7 percent. Of course, all of these magnitudes were amplified in markets with larger pre-merger market shares of Aetna and Prudential.[33]

Our results confirm that Americans are indeed paying a premium on their premiums. However, consolidation explains very little of the steep increase in health insurance premiums in recent years. While 2.1 percent is large in absolute terms – it translates into ~$17 billion in extra annual profits – it pales in comparison to the doubling in real premiums for our sample during the same 1998-2006 time period.[34] These findings do not imply insurance markets are competitive, however, only that consolidation has not raised premiums much. The industry was sufficiently concentrated even before the recent wave of consolidations to sustain supra-competitive prices. To the extent insurance carriers behaved as a "disciplined" oligopoly by the late 1990s, there may have been little room to optimally raise premiums in the wake of further

[33] To be more precise, the market shares need be large *and* overlapping. The predicted change in local market HHI associated with the merger equals 2*Aetna share * Prudential share. See equation 2 for the derivation.

[34] To calculate real premium growth during the study period, we divide premium by demofactor and convert to 2000 dollars using the annual CPI. This yields the weighted average premium per "effective enrollee" in our sample, which rose from $1,772 to $3,601 (in $2000, between 1998 and 2006).

consolidation. While it is beyond the scope of this paper to assess the welfare implications of the earnings and employment effects we document, we note these findings confirm the exercise of monopsonistic power in some markets.

We caution that our analysis relies on a single merger, albeit one that effectively generated 139 experiments (one per geographic market) that we exploit to generate our estimates. Additional research that utilizes other exogenous sources of variation in market structure would be invaluable to assessing conduct in this important industry. There has also been a great deal of consolidation across (as opposed to within) markets, and the effects of such consolidation are not reflected in our estimates.

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Appendix: Representativeness of the LEHID Dataset

This appendix compares the LEHID data to the two leading alternative sources of insurance data: the Kaiser Family Foundation/Health Retirement Education Trust (KFF/HRET) Annual Survey of Employer Benefits, and the proprietary Interstudy database of insurer data. The KFF/HRET survey randomly samples public and private employers to obtain national statistics on employer-sponsored health insurance; approximately 2000 employers respond each year. The data are not publicly available, nor is the sample designed to provide estimates at the market level. However, the survey is designed to yield representative estimates of national trends. Appendix Figure 1 below reports the annual growth rate in premiums for a family of four in an employer-sponsored plan. As in LEHID, both employer and employee premium contributions are combined, and both fully and self-insured plans are included. However, LEHID does not report premiums for a standard family size. Thus, to obtain a comparable measure from the LEHID sample, we divide the average annual premium in LEHID by the demographic factor. According to our source, this yields the premium per “person equivalent.” Annual growth rates for this “individual” premium are reported in Appendix Figure 1 as well. The trends are quite similar throughout the period.

We also compare our measures of market concentration with measures constructed by other researchers using the proprietary InterStudy database. InterStudy reports enrollment and premium figures at the insurer and MSA level. We compare the HHI and number of carriers tabulated by Scanlon et al (2008) to the corresponding figures from the LEHID data. [35]

Before describing the results, we note the InterStudy data is not directly comparable to LEHID for several reasons. The InterStudy data includes only fully-insured HMO plans for the time period we consider, and the allocation of enrollment across geographic markets is fairly noisy. In addition to these issues, the LEHID geographic markets, which generally correspond to MSAs (but may include multiple MSAs), are often larger than the Interstudy markets.[36]

To compare measures of insurer market structure derived from the two sources, we begin by mapping MSAs to the corresponding LEHID markets.[37] When multiple MSAs comprise one LEHID market, we weight the InterStudy MSA measures of market structure by the population of that MSA (obtained from the 2000 Census) to create measures of insurer market concentration (HHI, number of carriers) for each geographic market defined in the LEHID dataset.

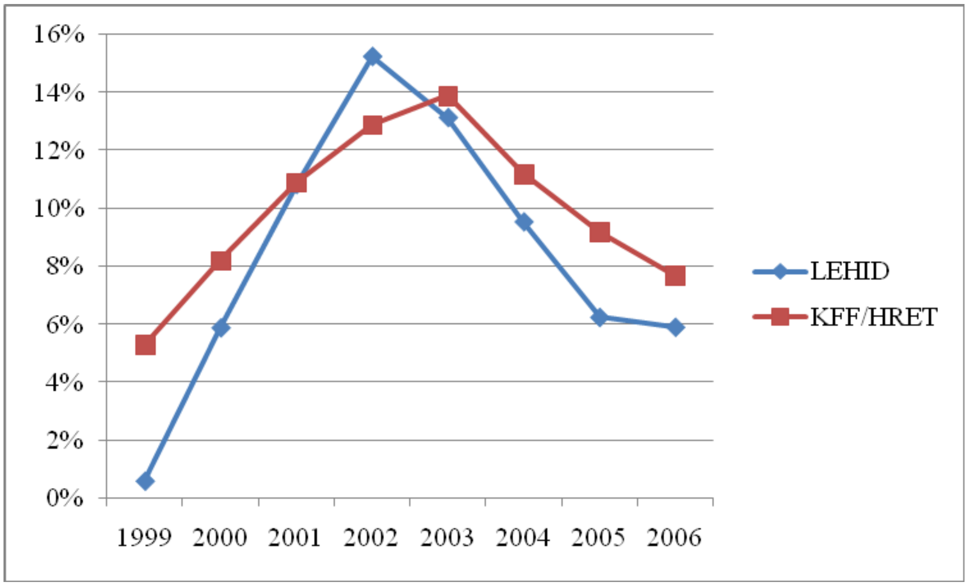
When we use all plans in the LEHID dataset to construct HHI (as in our regression models), the correlation coefficient between the two measures is 0.18 (N=139). This figure rises to 0.31 when we restrict attention to HMO plans only.[38] As is apparent in Appendix Figure 2, there are also some differences between the two estimates when we compare trends over time. The LEHID HHI exhibits fairly steady growth in the latter half of the study period while the Interstudy HHI peaks in 2003. Unfortunately, there are no obvious explanations for these discrepancies.

We use the LEHID-based HHI estimates for theoretical and practical reasons. First, the set of carriers that serve large, multisite firms such as those included in LEHID may differ from the set of carriers at large. Thus, LEHID itself likely offers the best estimate of the relevant insurance market structure. Second, the InterStudy data does not consistently include PPO enrollment during our study period, and PPOs account for a large share of our data. Third, as

[35] Our sincere thanks to Mike Chernew, Dennis Scanlon and Woolton Lee for sharing their estimates of market structure. For details on the construction of the InterStudy HHIs, see Scanlon et al (2006).
[36] For example, the entire state of Maine, is a single geographic market in the LEHID data.
[37] We were able to find a match for 284 out of a total of 328 MSAs present in the Interstudy dataset
[38] Note that the InterStudy estimates include only fully-insured plans, while the LEHID estimates include both fully-insured and self-insured plans. If we construct LEHID HHIs using only fully-insured plans, the corresponding correlation coefficients are .27 and .32 respectively.

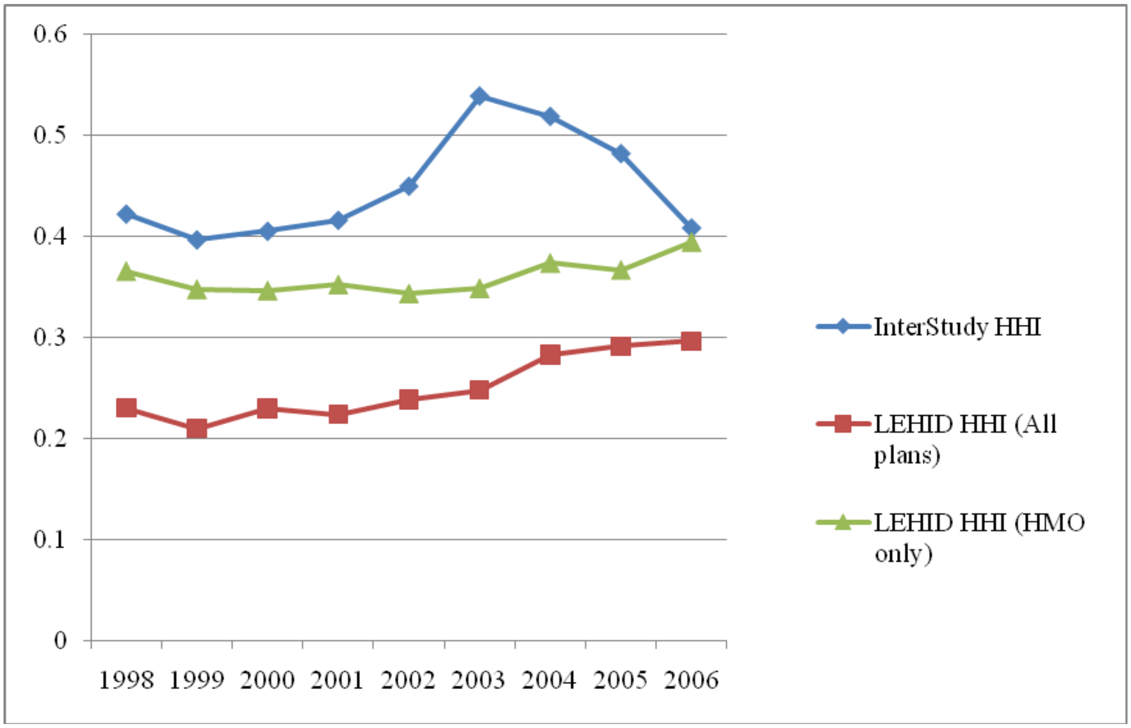
noted above, researchers have documented serious concerns about the way in which InterStudy allocates enrollment across MSAs. Finally, the InterStudy data is quite expensive to acquire.

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



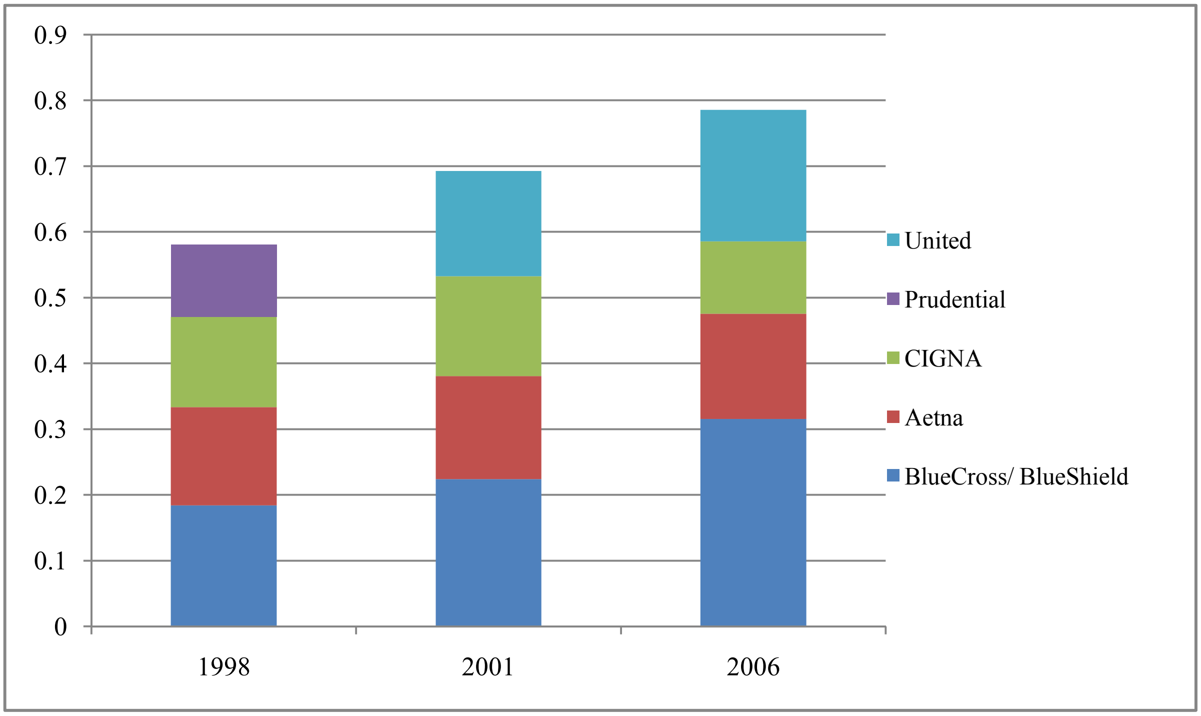
Sources: LEHID sample (all plans), and *2007 Kaiser/HRET Annual Survey of Employer-Sponsored Health Benefits* Annual growth rates for the LEHID sample are calculated using employee-weighted average premiums/demographic factor for each year. Both sources combine fully insured and self-insured plans.

Appendix Figure 2: Trends in HHI, LEHID vs. Interstudy



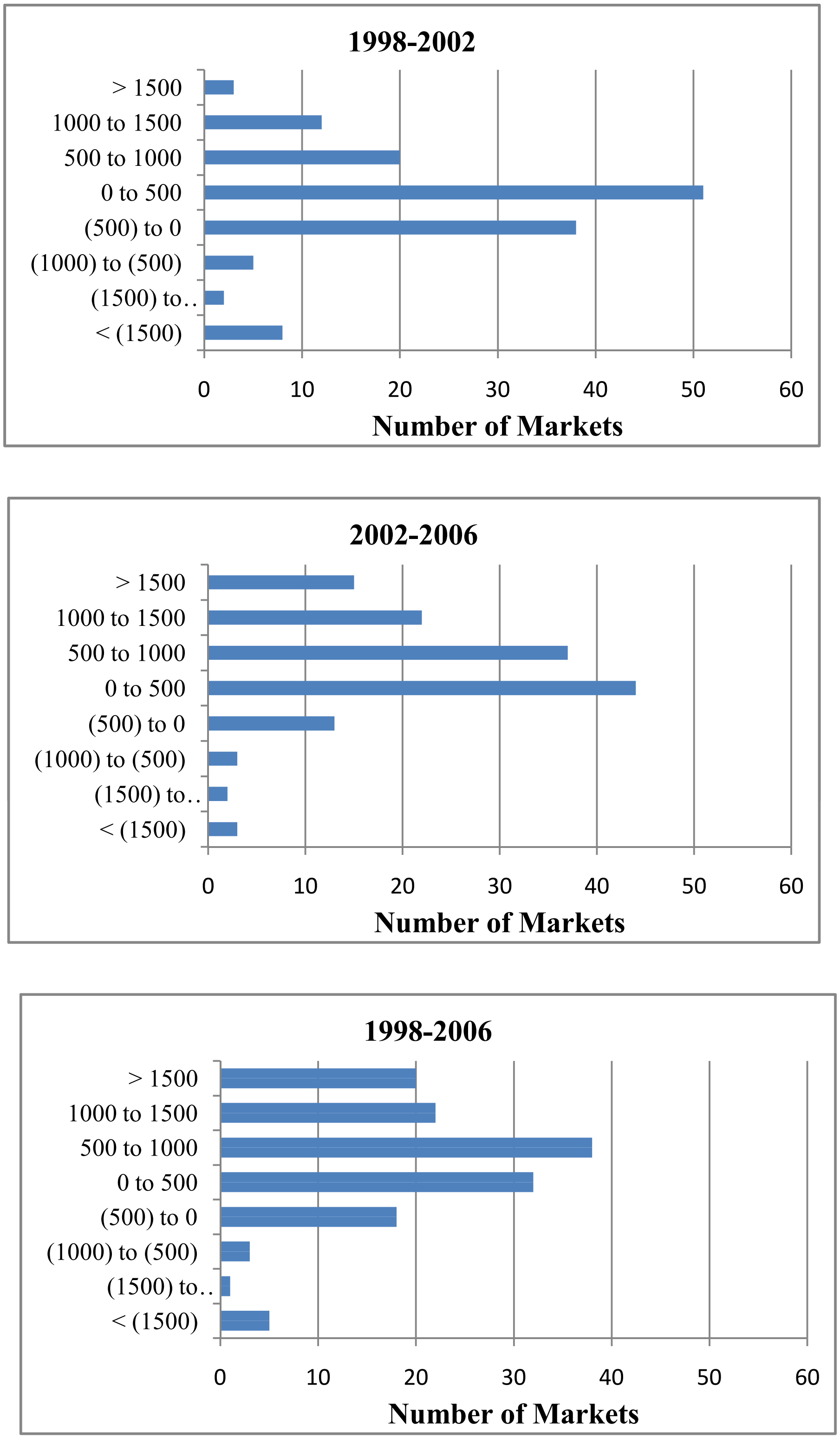
Sources: LEHID sample (all plans), and Scanlon et al. (2008)

Figure 1. Nationwide Four Firm Concentration Ratio, 1998-2006



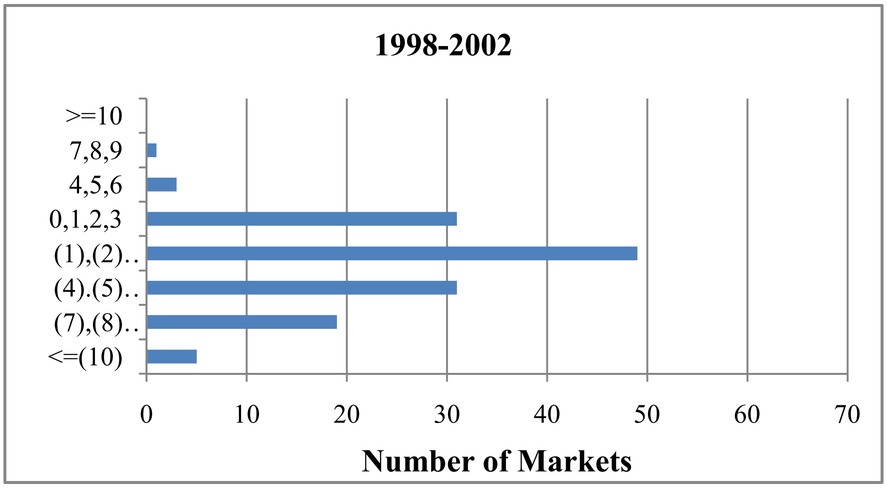
Note *: Market shares are computed using the entire sample (Table 1)*

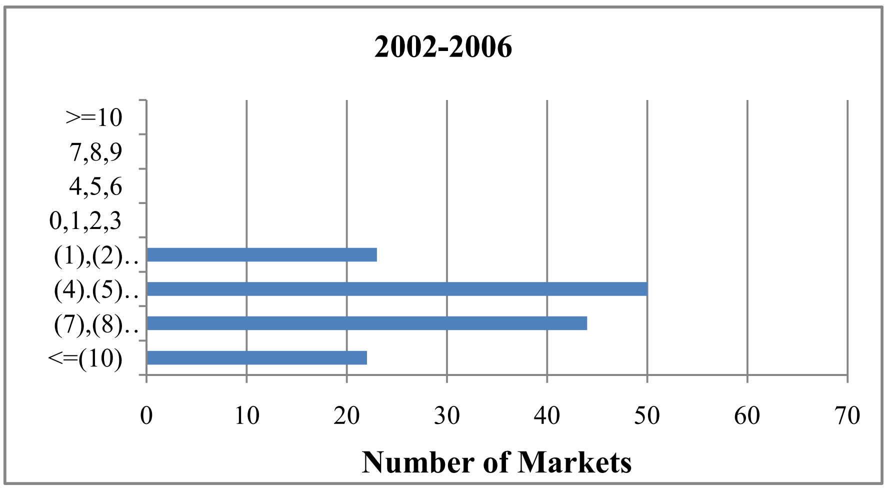
Figure 2. Change in Local Market Herfindahl



Note : HHI is scaled from 0 to 10,000

Figure 3. Change in Number of Carriers per Market





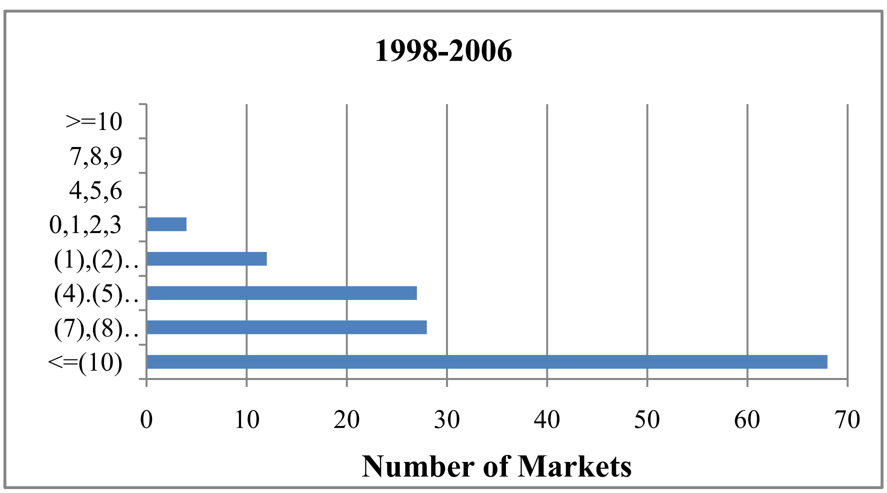
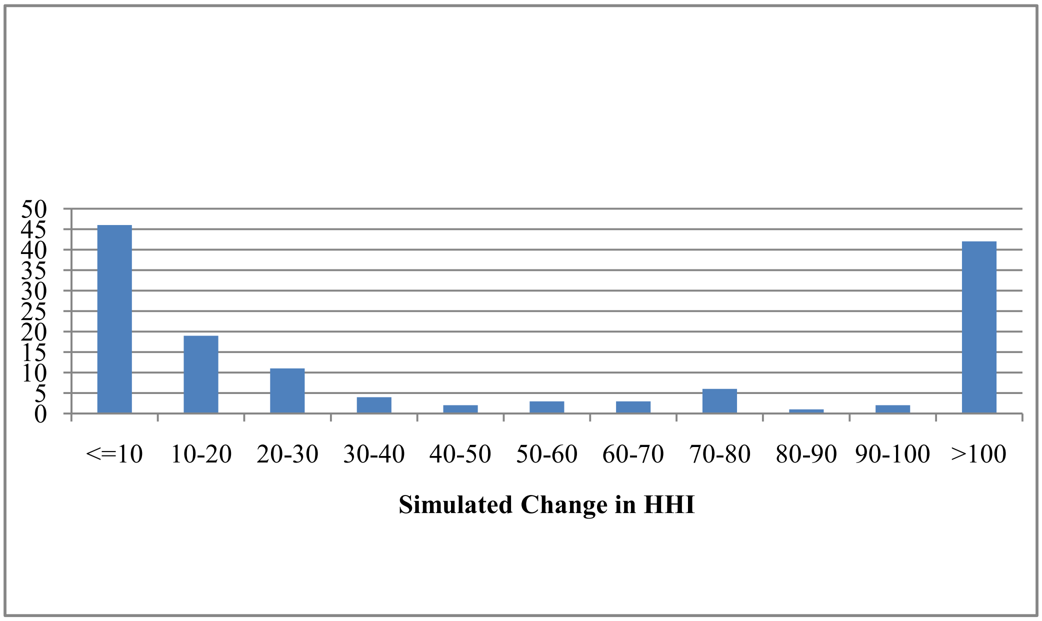


Figure 4. Distribution of Simulated Change in HHI Resulting from Aetna-Prudential Merger



<u>Note</u> : HHI is scaled from 0 to 10,000

Table 1. Descriptive Statistics

	1998	1999	2000	2001	2002	2003	2004	2005	2006
Premium ($)	3995.50	4125.50	4426.32	4868.92	5545.23	6338.24	6925.26	7400.19	7835.63
	1118.70	*1161.40*	*1222.23*	*1292.52*	*1425.18*	*1565.92*	*1734.47*	*1860.18*	*2014.87*
Number of Enrollees	181.70	165.40	156.30	173.03	174.42	178.65	171.32	196.42	190.16
	630.20	*553.57*	*475.18*	*545.77*	*577.56*	*619.76*	*523.98*	*828.83*	*640.60*
Demographic Factor	2.34	2.26	2.24	2.25	2.29	2.29	2.33	2.32	1.84
	0.50	*0.43*	*0.43*	*0.42*	*0.44*	*0.42*	*0.43*	*0.43*	*0.39*
Plan Design	1.06	1.06	1.04	1.06	1.06	1.04	1.03	0.99	0.99
	0.08	*0.08*	*0.09*	*0.08*	*0.08*	*0.08*	*0.09*	*0.09*	*0.09*
Plan Type									
HMO	41.1%	43.0%	40.4%	39.9%	39.4%	36.6%	33.8%	33.5%	33.4%
Indemnity	20.4%	17.8%	13.6%	10.6%	9.9%	7.7%	6.4%	4.8%	4.8%
POS	22.8%	18.1%	20.1%	17.8%	14.9%	14.4%	14.8%	13.6%	13.5%
PPO	15.5%	21.1%	25.8%	31.6%	35.7%	41.2%	44.9%	48.0%	48.2%
% Fully Insured	44.7%	45.0%	39.0%	36.6%	32.4%	26.2%	23.9%	21.3%	19.8%
<u>Market-Level Measures (counting each market once)</u>									
Herfindahl Index	0.23	0.21	0.23	0.22	0.24	0.25	0.28	0.29	0.30
	0.09	*0.08*	*0.08*	*0.07*	*0.08*	*0.08*	*0.09*	*0.10*	*0.11*
Four-firm Concentration	0.79	0.77	0.81	0.80	0.83	0.83	0.87	0.87	0.90
	0.09	*0.10*	*0.10*	*0.09*	*0.08*	*0.08*	*0.08*	*0.07*	*0.07*
Number of Carriers	18.88	20.07	15.80	17.67	16.10	16.38	13.16	13.14	9.63
	6.38	*6.17*	*5.38*	*5.42*	*4.64*	*4.60*	*3.87*	*3.39*	*2.82*
Lagged ln (Medicare costs)	*8.54*	*8.48*	*8.48*	*8.54*	*8.62*	*8.69*	*8.75*	*8.82*	*8.88*
	0.17	*0.17*	*0.16*	*0.16*	*0.16*	*0.15*	*0.15*	*0.14*	*0.14*
Lagged unemp rate	4.89	4.51	4.24	3.99	4.66	5.55	5.78	5.40	5.09
	1.65	*1.64*	*1.49*	*1.06*	*1.01*	*1.09*	*1.15*	*1.08*	*1.14*
Lagged Hospital HHI	0.12	0.20	0.13	0.13	0.13	0.13	0.14	0.13	0.14
	0.06	*0.10*	*0.07*	*0.07*	*0.07*	*0.07*	*0.08*	*0.07*	*0.08*
Number of Employers	194	205	199	242	255	330	246	262	229
Number of Markets	139	139	139	139	139	139	139	139	139
Number of Observations	22074	25678	23661	29114	31539	33692	26575	26473	21854

<u>Notes:</u> *All statistics are unweighted. The unit of observation is an employer-carrier-market-plantype-year combination, unless noted otherwise. Standard deviations are in italics. 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. Descriptive Statistics (Unit of Observation: Employer-Market-Year)

	1998	1999	2000	2001	2002	2003	2004	2005	2006
Premium ($)	4104.47	4185.45	4495.88	4914.50	5624.70	6443.94	6980.52	7455.44	7832.46
	1047.76	*1019.94*	*1100.30*	*1184.72*	*1280.61*	*1423.89*	*1583.40*	*1727.21*	*1807.98*
Number of Enrollees	399.86	368.17	333.68	364.29	370.42	368.85	334.76	371.10	361.47
	1465.47	*1289.57*	*1111.06*	*1303.26*	*1397.66*	*1317.26*	*1030.86*	*1803.23*	*1245.86*
Demographic Factor	2.35	2.27	2.26	2.26	2.29	2.32	2.34	2.33	1.84
	0.47	*0.40*	*0.40*	*0.40*	*0.41*	*0.40*	*0.40*	*0.41*	*0.38*
Plan Design	1.05	1.05	1.03	1.05	1.05	1.04	1.02	0.98	0.98
	0.06	*0.06*	*0.06*	*0.06*	*0.06*	*0.06*	*0.07*	*0.07*	*0.07*
Plan Type									
HMO	29.4%	32.8%	30.6%	29.6%	30.6%	28.7%	25.8%	25.1%	25.4%
Indemnity	22.4%	17.2%	12.2%	8.8%	7.2%	5.0%	3.9%	2.2%	2.8%
POS	28.1%	22.3%	24.6%	20.1%	16.8%	16.2%	16.3%	15.1%	14.1%
PPO	20.0%	27.7%	32.6%	41.6%	45.4%	50.0%	54.0%	57.6%	57.6%
% Fully Insured	33.0%	35.5%	30.0%	27.4%	24.2%	19.5%	17.1%	14.9%	14.4%
Number of Observations	10033	11536	11086	13829	14851	16318	13600	14012	11497

Notes: *All statistics are unweighted. The unit of observation is an employer-market-year combination. Standard deviations are in italics. 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 3. OLS Regression of Δ Log (Premium) on Δ Lagged HHI

Panel A. Dependent Variable = Δ Log Premium from 1999-2006

	(1)	(2)	(3)
Δ Lagged HHI	-0.0609*	-0.0260	-0.0236
	(0.0256)	*(0.0218)*	*(0.0225)*
Market controls			
Δ Lagged ln(Medicare costs per cap)	-0.0084	0.0276	0.0293
	(0.0459)	*(0.0442)*	*(0.0445)*
Δ Lagged Unemp rate	-0.2106	-0.3425	-0.2621
	(0.2706)	*(0.2791)*	*(0.3035)*
Δ Lagged Hospital HHI	-0.0091	-0.0037	-0.0084
	(0.0306)	*(0.0285)*	*(0.0292)*
Employer Market controls			
Δ Demographic factor	0.3307***	0.3236***	0.3201***
	(0.0104)	*(0.0111)*	*(0.0108)*
Δ % Insured by HMO		-0.0294*	-0.0786***
		(0.0114)	*(0.0133)*
Δ % Insured by PPO		0.0307**	0.0558***
		(0.0099)	*(0.0105)*
Δ % Insured by Indemnity		0.1095***	0.1515***
		(0.0113)	*(0.0128)*
Δ Plan Design			0.5863***
			(0.1075)
Employer fixed effects	Yes	Yes	Yes
Adjusted R2	0.6824	0.6982	0.7045
# Observations	3164	3164	3164

Notes: Unit of observation is the employer-market-year. Sample is restricted to employer-market observations present in both 1999 and 2006. Change in % Insured by POS is the omitted category. Standard errors are clustered by market. HHI is scaled from 0 to 1.

**** signifies p<.01, ** signifies p<.05 and * signifies p<.10*

Table 3. OLS Regression of Δ Log (Premium) on Δ Lagged HHI

Panel B. Dependent Variable = Δ Log Premium from 1999-2002

	(1)	(2)	(3)
Δ Lagged HHI	-0.0175	-0.0167	-0.0165
	(0.0297)	*(0.0265)*	*(0.0262)*
Market controls			
Δ Lagged ln(Medicare costs per cap)	-0.0715+	-0.0467	-0.0464
	(0.0409)	*(0.0391)*	*(0.0391)*
Δ Lagged Unemp rate	0.2947	0.2155	0.2069
	(0.3005)	*(0.2944)*	*(0.2917)*
Δ Lagged Hospital HHI	-0.0119	-0.0075	-0.0070
	(0.0268)	*(0.0236)*	*(0.0235)*
Employer Market controls			
Δ Demographic factor	0.3388***	0.3360***	0.3364***
	(0.0092)	*(0.0090)*	*(0.0090)*
Δ % Insured by HMO		-0.0584***	-0.0501***
		(0.0104)	*(0.0118)*
Δ % Insured by PPO		0.0471***	0.0410**
		(0.0110)	*(0.0126)*
Δ % Insured by Indemnity		0.1012***	0.0918***
		(0.0123)	*(0.0127)*
Δ Plan Design			-0.1206
			(0.1127)
Employer fixed effects	Yes	Yes	Yes
Adjusted R2	0.6159	0.6390	0.6392
# Observations	5537	5537	5537

Notes : Unit of observation is the employer-market-year. Sample is restricted to employer-market observations present in both 1999 and 2002 Change in % Insured by POS is the omitted category. Standard errors are clustered by market. HHI is scaled from 0 to 1.

**** signifies p<.01, ** signifies p<.05 and * signifies p<.10*

Table 3. OLS Regression of Δ Log (Premium) on Δ Lagged HHI

Panel C. Dependent Variable = Δ Log Premium from 2002-2006

	(1)	(2)	(3)
Δ Lagged HHI	-0.0692*	-0.0556*	-0.0609*
	(0.0269)	*(0.0262)*	*(0.0270)*
Market controls			
Δ Lagged ln(Medicare costs per cap)	-0.0112	-0.0028	0.0045
	(0.0372)	*(0.0367)*	*(0.0377)*
Δ Lagged Unemp rate	0.2888	0.2283	0.2495
	(0.2222)	*(0.2202)*	*(0.2379)*
Δ Lagged Hospital HHI	-0.0230	-0.0095	-0.0123
	(0.0586)	*(0.0577)*	*(0.0597)*
Employer-market controls			
Δ Demographic factor	0.3409***	0.3413***	0.3402***
	(0.0093)	*(0.0093)*	*(0.0090)*
Δ % Insured by HMO		-0.0246**	-0.0619***
		(0.0092)	*(0.0094)*
Δ % Insured by PPO		0.0184*	0.0402***
		(0.0093)	*(0.0099)*
Δ % Insured by Indemnity		0.0553***	0.0763***
		(0.0136)	*(0.0138)*
Δ Plan Design			0.5342***
			(0.0715)
Employer fixed effects	Yes	Yes	Yes
Adjusted R2	0.6703	0.6742	0.6812
# Observations	6031	6031	6031

Notes : Unit of observation is the employer-market-year. Sample is restricted to employer-market observations present in both 2002 and 2006. Change in % Insured by POS is the omitted category. Standard errors are clustered by market. HHI is scaled from 0 to 1.
**** signifies p<.01, ** signifies p<.05 and * signifies p<.10*

Table 4. Effect of the Aetna-Prudential Merger on Market Concentration

Dependent Variable = HHI

	(1) 1998-2001	(2) 1998-2001
Sim Δ HHI * POST	0.499***	0.486***
	(0.137)	*(0.136)*
Sim Δ HHI * POST * (Texas==1)		-1.20**
		(0.517)
Texas * POST		0.054
		(0.040)
Market controls		
ln(Medicare costs per cap)	0.077	0.091
	(0.131)	*(0.122)*
Unemp rate	-0.000	-0.001
	(0.005)	*(0.005)*
Hospital HHI	-0.058	-0.037
	(0.069)	*(0.065)*
Texas included?	No	Yes
# Observations	532	556
R-squared	0.560	0.563

Notes: The unit of observation is the market-year. All specifications include market and year fixed effects. Standard errors are robust. HHI is scaled from 0 to 1.

***** signifies p<.01, ** signifies p<.05, * signifies p<.10**

Table 5. Estimating Impact of HHI on Premiums, 1998-2002

Dependent Variable = Annual Change in ln(Premium)

Reduced Form Estimates				
Sim ΔHHI*POST	0.1686***	0.1367***	0.1442***	0.146***
	(0.0434)	*(0.0445)*	*(0.0440)*	*(.045)*
Sim ΔHHI*POST*(Texas = 1)				-0.132**
				(.055)
IV Estimates				
lagged HHI	0.3560***	0.2887**	0.3041**	---
	(0.1279)	*(0.1257)*	*(0.1254)*	---
OLS Estimates				
lagged HHI	0.0138	0.0114	0.0097	---
	(0.0210)	*(0.0178)*	*(0.0179)*	---
Employer-market controls				
Δ Plan Type Shares	No	Yes	Yes	Yes
Δ Plan Design	No	No	Yes	Yes
Texas Observations Included?	No	No	No	Yes
Number of Observations	28645	28645	28645	30493

Notes: The unit of observation is the employer-market-year. All specifications include change in demographic factor, change in lagged market covariates and employer, market and year fixed effects. Change in % Insured by POS is the omitted category. Standard errors are robust. HHI is scaled from 0 to 1.

Table 6. Estimating Impact of Consolidation on non-price variables, 1998-2002

Dependent Variable = Annual Change in

	Plan Design	***Fraction of HMO Enrollees***	***Fraction of Indemnity Enrollees***	***Fraction of PPO Enrollees***
Sim ΔHHI*POST	-0.0650***	-0.183*	0.234***	0.0868
	(0.0154)	*(0.1103)*	*(0.064)*	*(0.0637)*
Sim ΔHHI*POST*(Texas == 1)	0.0579***	0.1519	-.0986	0.1014
	(0.0216)	*(0.1325)*	*(0.088)*	*(.0848)*
Texas Observations Included?	Yes	Yes	Yes	Yes
Number of Observations	30493	30493	30493	30493

Notes: The unit of observation is the employer-market-year. All specifications include changes in lagged market covariates and employer, market and year fixed effects. Standard errors are robust. HHI is scaled from 0 to 1.

Table 7. Descriptive Statistics (OES Survey Data)

	1999	2000	2001	2002
All Occupation Categories				
Average Earnings	42251.23	43957.49	45445.53	49133.85
	21261.86	*21781.58*	*22030.18*	*29010.01*
No of Employees in Occupation-Market	1538.51	1240.81	1219.83	1193.92
	5805.37	*4909.60*	*4808.85*	*4680.37*
Physicians				
Average Earnings	113493.90	113301.12	116317.51	149584.12
	16654.64	*13630.17*	*13256.51*	*23923.14*
No of Employees in Occupation-Market	1154.46	1431.91	1414.39	1412.87
	2057.49	*2205.37*	*2253.86*	*1948.58*
Nurses				
Average Earnings	39601.00	41245.04	42981.85	44211.38
	5291.69	*5908.12*	*5895.71*	*6185.71*
No of Employees in Occupation-Market	16241.87	16113.73	16330.87	16405.28
	18780.68	*17812.83*	*17635.21*	*17248.18*
Totals				
Number of Employees	3398560	3657910	3758310	3771600
Number of Physicians	106210	173260	173970	172370
Number of Nurses	2030230	2030330	2057690	2050660
Number of Occupation Categories	35	35	35	35
Number of Markets	126	126	126	126
Number of Observations	2209	2948	3081	3159

__Notes:__ The unit of observation is an occupation-market combination. Sample does not include markets present in the state of Texas, where the DoJ imposed restrictions on the Aetna-Prudential merger. The OES survey collects hourly wage data in 12 intervals. The mean wage value for the upper open-ended wage interval is set at the lower end of the range (this practice has been modified in recent years). Standard deviations are in Italics.

Table 8. Effect of the Aetna-Prudential Merger on Healthcare Provider Earnings and Employment, 1999-2002

	Dependent Variable = Δ Log (Average Income) from 99-02			***Dependent Variable = Δ Log (Employment) from 99-02***		
Simulated Δ HHI	0.1106	0.0781	0.0909	-2.3723**	-2.7233**	-2.4365*
	(0.1800)	*(0.2153)*	*(0.2035)*	*(0.8088)*	*(0.9413)*	*(0.9776)*
Physician Indicator	0.1932***	0.1837***	N/A	0.5225**	0.4974**	N/A
	(0.0335)	*(0.0352)*		*(0.1700)*	*(0.1673)*	
Physician * Simulated Δ HHI	-2.0067*	-2.1795**	-2.1954**	-2.5065	-2.5823	-2.8580
	(0.8328)	*(0.8010)*	*(0.8109)*	*(7.9335)*	*(8.4406)*	*(8.4392)*
Nurse Indicator	-0.0131*	-0.0154**	N/A	-0.1537***	-0.1601***	N/A
	(0.0055)	*(0.0057)*		*(0.0253)*	*(0.0274)*	
Nurse * Simulated Δ HHI	0.4399*	0.4707+	0.4570+	1.7065*	2.0123+	1.7378+
	(0.2209)	*(0.2571)*	*(0.2541)*	*(0.8451)*	*(1.0711)*	*(1.0323)*
Δ Hospital HHI, 1999-2002	0.0234	0.0213	0.0237	-0.0240	-0.0270	-0.0674
	(0.0293)	*(0.0309)*	*(0.0315)*	*(0.2535)*	*(0.2469)*	*(0.2348)*
Trend in Dep Var, 1997-1998	No	Yes	Yes	No	Yes	Yes
Occupation Fixed Effects	No	No	Yes	No	No	Yes
# Observations	2110	1631	1631	2110	1631	1631

Notes : Unit of observation is the occupation-market-year. All physician occupations are lumped into one category. Specifications are restricted to ocupation- markets present in both 1999 and 2002. Simulated HHI is scaled from 0 to 1. Sample does not include observations from Texas where the DOJ blocked the merger in two markets. All specifications are weighted by average estimated employment in each occupation-market. Standard errors are robust and clustered by Market.

***** signifies p<.01, ** signifies p<.05, * signifies p<.10**