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Abstract

We examine whether and to what extent consolidation in the US health insurance industry has contributed to higher employer-sponsored insurance premiums. We exploit the differential impact across local markets of a national merger of two insurers to identify the causal effect of concentration on premiums. Using data for large groups, we estimate premiums in average markets were approximately seven percentage points higher by 2007 due to increases in local concentration from 1998-2006. We also find evidence consolidation facilitates the exercise of monopsonistic power vis-Ã -vis physicians, leading to reductions in their absolute employment and earnings relative to other healthcare workers

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Paying a Premium on Your Premium? Consolidation in the US Health Insurance Industry†

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Although the majority of health-care spending in the United States is funneled through the private 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," as defined by the *Horizontal Merger Guidelines* issued in 1997 by the US Department of Justice and the Federal Trade Commission. In 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 this seven-year period, the average, inflation-adjusted premium for employer-sponsored family coverage rose 48 percent (to $12,680 in 2008)[2] while real median household income *declined* by 2 percent to $50,303 (DeNavas-Walt, Proctor, and Smith 2009).

Prior studies point to the potential for insurer consolidation to raise premiums (e.g., Robinson 2004; Wholey, Feldman, and Christianson 1995; and

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[1] "Competition in Health Insurance: A Comprehensive Study of US Markets," American Medical Association, 2001 and 2008. These figures are based on the reported levels of the Herfindahl-Hirschman Index (HHI) for HMOs and PPOs combined. 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. The *Horizontal Merger Guidelines* issued in 1992 and updated in 1997 define markets with HHI $>$ 1,800 as "highly concentrated." A recent update adjusted this threshold to 2,500 (DOJ 2010), and as a result the share of markets in 2008 that would be highly concentrated is somewhat lower at 70 percent.

[2] The corresponding increase for single coverage was 44 percent (Kaiser Family Foundation/Health Research and Educational Trust *Employer Health Benefits Annual Survey 2009*). Premiums include both employer and employee contributions, and are adjusted to 2008 dollars by the authors using the CPI-U.

Dafny 2010); however, none attempts to quantify this effect.[3] From a theoretical standpoint, the effect of concentration on insurance premiums is ambiguous. On 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 concentration may strengthen insurers' bargaining positions vis-à-vis health-care providers, leading to reduced negotiated reimbursements and lower premiums. In addition, there are many potential sources of efficiency gains from consolidation, including economies of scale in investments in information technologies (IT) investing and disease management programs. Such efficiency gains would reduce optimal premiums.[4] The net effect on insurance premiums is ultimately an empirical question.

There are two key challenges to empirically estimating such a link: (i) adequate data and (ii) plausibly exogenous variation in market concentration. Regarding the first issue, comprehensive data on a large sample of health plans are extremely difficult to obtain because contracts are customized for each buyer across many dimensions, renegotiated annually, and considered highly confidential. In addition, premiums vary based on the demographics, health risks, and expenditure history of the insured population. Thus, it is difficult to calculate a standardized premium to enable comparisons across employers and/or markets. With respect to the second challenge, highly concentrated markets (or markets that are becoming more concentrated) are likely to differ from other markets in unobservable ways, making it difficult to separately identify the effect of concentration from other factors.

We address these challenges as follows. First, we utilize detailed longitudinal data on the health plans offered by a sample of more than 800 employers in 139 distinct geographic markets in the United States. The data span the nine years between 1998 and 2006 and represent approximately 10 million active employees and their dependents in each year. Rather than attempting to standardize premiums across different employee populations, products, and plan designs, we focus on the *growth rate* of health insurance premiums for the same employer in a specific geographic market over time and examine how this relates to the local market structure of health insurers. Focusing on growth alleviates concerns about time-invariant unobservable differences in the risk profiles of employee groups and the characteristics of plans they utilize that may be correlated with premium *levels*. We also control for the influence of time-varying measures such as employee demographics, the types of plans utilized (HMO, PPO, etc.), and the generosity of benefit design.

After documenting trends in the level and growth of concentration (as measured by the Herfindahl-Hirschman index (HHI), which is the sum of squared market shares) in 139 distinct geographic markets, we estimate OLS models of the relationship between premium growth and concentration levels. We do not find evidence

[3] Robinson (2004) shows that state-level insurance markets are dominated by a small number of firms and observes that insurer profits increased rapidly over 2000–2003. Wholey, Feldman, and Christianson (1995) report that premiums per HMO member are negatively related to the number of competitors facing the HMO in question, controlling for a host of HMO and market characteristics such as per capita income, Blue Cross affiliation and HMO ownership status. Last, Dafny (2010) finds health insurers engage in "direct" price discrimination, charging higher premiums to firms with deeper pockets, as measured by operating profits. This evidence of price discrimination implies insurers possess and exercise market power in some local markets but does not yield an estimate of the contribution of imperfect competition in this market to premium growth.

[4] Of course, rent transfers from providers to insurers are not true efficiency gains, although they may reduce premiums.

that premiums are rising more quickly in markets that are becoming more concentrated. While these estimates are useful for descriptive purposes, they are unlikely to provide causal estimates of the impact of market structure on premiums. Differences in HHI across markets—or even changes in HHI within markets—are likely to be driven by many factors that are not exogenous to premiums. These include differences (or changes) in consumer preferences and constraints, product offerings and pricing strategies, and the market conduct of hospitals, physicians, and other health-care providers. 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 (relative to other markets). This pattern does not imply consolidations in such a market would reduce premium growth, ceteris paribus.

In order to address the endogeneity challenge and obtain a credible estimate of the impact of concentration on premium growth, 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 premerger market shares of the two firms varied significantly across specific geographic markets, resulting in very different shocks to post-merger concentration. For example, in our sample the premerger market shares of Aetna and Prudential in Jacksonville, Florida were 19 and 24 percent, respectively, versus just 11 and 1 percent, respectively, in Las Vegas, Nevada. Holding all else equal, this implies an increase in post-merger HHI of 892 points in Jacksonville, but only 21 points in Las Vegas. 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 and controlling as fully as possible for changes in the characteristics of health plans (such as benefit design).

The point estimates indicate that rising concentration in local health insurance markets accounts for a nontrivial share of premium growth in recent years. Specifically, our instrumental variables estimates imply that the mean increase in local market HHI between 1998 and 2006 (inclusive) raised premiums by roughly 7 percent from their 1998 baseline, all else equal. Given private health insurance expenditures of $490 billion in our base year 1998, *if* this result is generalizable, then the "premium on premiums" by 2007 is on the order of $34 billion per year, or about $200 per person with employer-sponsored health insurance.[5]

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 health-care personnel, we exploit the differential impact across geographic markets of the Aetna-Prudential merger to examine whether there is a causal link between concentration and these outcomes. Our analysis suggests that the growth in insurer bargaining power following this merger reduced earnings and employment growth of physicians and raised earnings and employment growth

[5] Source: National Health Expenditure Data provided by the Center for Medicare and Medicaid Services; available online at http://www.cms.hhs.gov/NationalHealthExpendData/. The vast majority of this spending is due to employer-sponsored plans; only 9 percent of the nonelderly privately insured have policies that are not employment based (DeNavas-Walt, Proctor, and Smith 2009). Additionally, this figure understates the size of the private health insurance industry as it excludes expenditures by Medicaid and Medicare managed care plans.

of nurses. This pattern of results is consistent with postmerger substitution of nurses for physicians, and the exercise of monopsony power vis-à-vis physicians.

The paper is organized as follows. Section I describes the data in detail. We examine the association between local market concentration and premium growth in Section II. In Section III we investigate whether a causal relationship exists between these two variables using the variation across geographic markets in the merger-induced increase in insurer concentration. Section IV contains our analyses of the relationship between concentration and health-care employment and earnings. Section V concludes.

I. Data

Our primary source is the Large Employer Health Insurance Dataset (LEHID). LEHID contains information on all of the health plans offered by a large sample of employers between 1998 and 2006, inclusive. It is an unbalanced panel gathered and maintained by a leading benefits consulting firm. The data are proprietary, and employers included in the dataset have some past or present affiliation with the firm. Online Appendix 1, which contains additional details of the data not presented here, illustrates that LEHID plans are on average very similar to the plans offered by a representative sample of large employers nationwide.

The original unit of observation is the health plan–year. A health plan is defined as a unique combination of employer, market, insurance type, insurance carrier, and plan type (e.g., Company X's Chicago-area fully insured Aetna HMO). There are 813 unique employers, 139 geographic markets, two insurance types (self- and fully insured), 357 insurance carriers[6] and four plan types (HMO, POS, PPO, Indemnity) represented in the data.[7] Most employers in LEHID are large, multisite, publicly traded firms, such as those appearing 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 data source using three-digit zip codes. According to the data provider, the 139 markets reflect the geographic boundaries typically used by insurance carriers when quoting prices. Large metropolitan areas are separate markets, and nonmetropolitan areas are lumped together within state boundaries (e.g., "New Mexico—Albuquerque" and "New Mexico—except Albuquerque").[8]

The sample includes both *fully insured* and *self-insured* plans. As these terms suggest, the former is "traditional" insurance in which the insured pays the carrier to bear the risk of realized health-care outlays. Many large employers choose to self-insure, outsourcing benefits management, provider contracting, and/or claims

[6] Many of these carriers are third-party administrators, who "rent" provider networks and process claims for self-insured employers.

[7] HMO and POS plans control utilization through primary care physicians ("gatekeepers"). HMOs cover only in-network providers, while POS and PPO plans provide some coverage for out-of-network providers. Indemnity plans have no gatekeepers or network restrictions.

[8] There is only one market that crosses state boundaries, "Massachusetts—Southern and Rhode Island." A few rural areas of the United States are excluded. A map of the markets is available in Dafny (2010).

administration but paying the realized costs of care. The percent of LEHID enrollees in self-insured plans increased from 55 to 80 percent during the study period.

In addition to the elements that jointly define a plan, our dataset includes 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, as well as premiums for stop-loss insurance, if any). 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. All of these characteristics are important determinants of average expected costs per enrollee in a plan. *Plan design factor* captures the generosity of benefits within a particular carrier–plan type, with an emphasis on the levels of coinsurance, copayments, and deductibles. Both factors are calculated by the source, and the proprietary formulae were not disclosed to us. Higher values of either factor are associated with higher premiums.

The LEHID also records the *number of enrollees* in each plan. This figure includes only employees of the relevant firm; dependents are accounted for by the demographic factor described above. The total number of enrollees in all LEHID plans averages 4.7 million per year. Given an average family size of more than two, this implies that more than ten million US residents are part of the sample in a typical year, representing approximately 7 percent of those with employer-sponsored insurance (ESI) during this period, and a much larger share of those insured through large firms.

We supplement the LEHID data with time-varying measures of local economic conditions (the unemployment rate, as reported by the Bureau of Labor Statistics), a measure of health-care utilization (Medicare costs per capita, as reported by the Centers for Medicare and Medicaid services), and the concentration of the hospital industry (HHI as calculated by the authors using the *Annual Surveys of Hospitals* administered by the American Hospital Association).[9] As the first two measures are reported at the county-year level, and LEHID markets are defined by three-digit zip codes, we make use of a mapping between zip codes and counties and, where necessary, use population data to calculate weighted average values for each LEHID market and year.

We perform most analyses using data aggregated to the employer-market-year level. Table 1 presents descriptive statistics for this unit of observation for 1998, 2002, and 2006, which represent the initial, middle, and final years of the sample respectively. Because our primary outcome is *growth* in health insurance premiums (in order to avoid cross-sectional identification of the coefficients of interest),

[9] To calculate HHI for each geographic market and year, we use data on the number of beds for all general hospitals located in the set of three-digit zip codes that define the market, assigning hospitals with the same "system ID" to a common owner.

Table 1—Descriptive Statistics (*Unit of Observation: Employer-Market-Year*)

	1998	2002	2006
Premium ($)	4,104.47	5,624.70	7,832.46
	(1,047.76)	(1,280.61)	(1,807.98)
Number of enrollees	399.86	370.42	361.47
	(1,465.47)	(1,397.66)	(1,245.86)
Demographic factor	2.35	2.29	1.84
	(0.47)	(0.41)	(0.38)
Plan design	1.05	1.05	0.98
	(0.06)	(0.06)	(0.07)
Plan type			
HMO	29.4%	30.6%	25.4%
Indemnity	22.4%	7.2%	2.8%
POS	28.1%	16.8%	14.1%
PPO	20.0%	45.4%	57.6%
Percent fully insured	33.0%	24.2%	14.4%
Observations	10,033	14,851	11,497

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

aggregating the data to the employer-market-year level enables us to use a much larger proportion of the data. With the health plan–level data, growth in premium is undefined when an employer terminates a particular plan. Analogously, new plans can enter the analysis only after multiple observations are available. Changes to plan offerings are quite common in our data (24 percent of plans in year t whose firm-markets are still present in year $t + 1$ no longer exist). Moreover, changes in market concentration may affect the insurance carriers and plan types chosen by employers, so we do not want a priori to eliminate this substitution from our sample.[10] Given this aggregation, both fully and self-insured plans must be included together in the analysis sample to ensure the set of employees represented over time is stable (but for hiring, attrition, and changes in employees' decisions to take up employer-sponsored insurance).

II. Is Premium Growth Correlated with Local Market Concentration?

In this section, we examine the relationship between the growth in health insurance premiums and 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 the corresponding market HHI. We include market fixed effects in our models, so that we identify the coefficient of interest using changes in *within-market* HHI. The richness

[10] This occurs very frequently in the LEHID. For example, consider employer-market pairs that are present in both 1999 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.

of the data also permits us to control for important time-varying differences (such as the percent of enrollees in HMOs and the magnitude of copayments). Although interesting as a descriptive exercise, this analysis is unlikely to yield unbiased estimates of the causal impact of changes in market structure on premium growth, as changes in market structure are unlikely to be exogenous.

A. *Market Structure of Large Group Insurance Markets, 1998–2006*

During our nine-year study period, the average market-level HHI (estimated using our sample and scaled from 0 to 10,000) increased from 2,286 to 2,984.[11] Using the categorization from the *Horizontal Merger Guidelines* issued in 1997, the fraction of markets falling into the top "highly concentrated" category (HHI $>$ 1,800) rose from 68 to 99 percent. The median four-firm concentration ratio increased from 79 to 90 percent. Thus, our data support the conclusions of well-publicized reports issued by the American Medical Association and the General Accounting Office: local health insurance markets are concentrated and becoming more so over time.[12]

Figure 1 presents histograms of the market-level changes in HHI, separately for 1998–2002, 2002–2006, and 1998–2006. The larger increases tended to occur during the second half of the study period, but sizable 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 or more 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 or more points when HHI already exceeds 1,800 are "presumed … likely to create or enhance market power or facilitate its exercise." 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 can be subdivided into "structural" (related to entry, exit, and consolidation) and "nonstructural" sources. Using data on fully insured HMOs only, Scanlon, Chernew, Swaminathan, and Lee (2006) report that 61 to 65 percent of the variation in HHI between 1998 and 2002 is attributable to structural changes. These changes are also important in our sample: the mean number of carriers per market declined from 18.9 in 1998 to 9.6 in 2006.[13] Of course, neither source of HHI change can be presumed exogenous to other determinants of premium growth. Consumer preferences simultaneously determine market shares and premium growth, and exit and consolidation of carriers may be impacted by expectations of premium growth.

[11] To gauge the impact of this change on concentration, consider the following two examples. A market with five insurers, four of which have a market share of 23.75 percent, would have an HHI of 2,281. A market with four insurers, three of which each have a market share of 31.33 percent, would have an HHI of 2,981.

[12] AMA *ibid*; GAO (2009a).

[13] As the data on HHI suggest, 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.

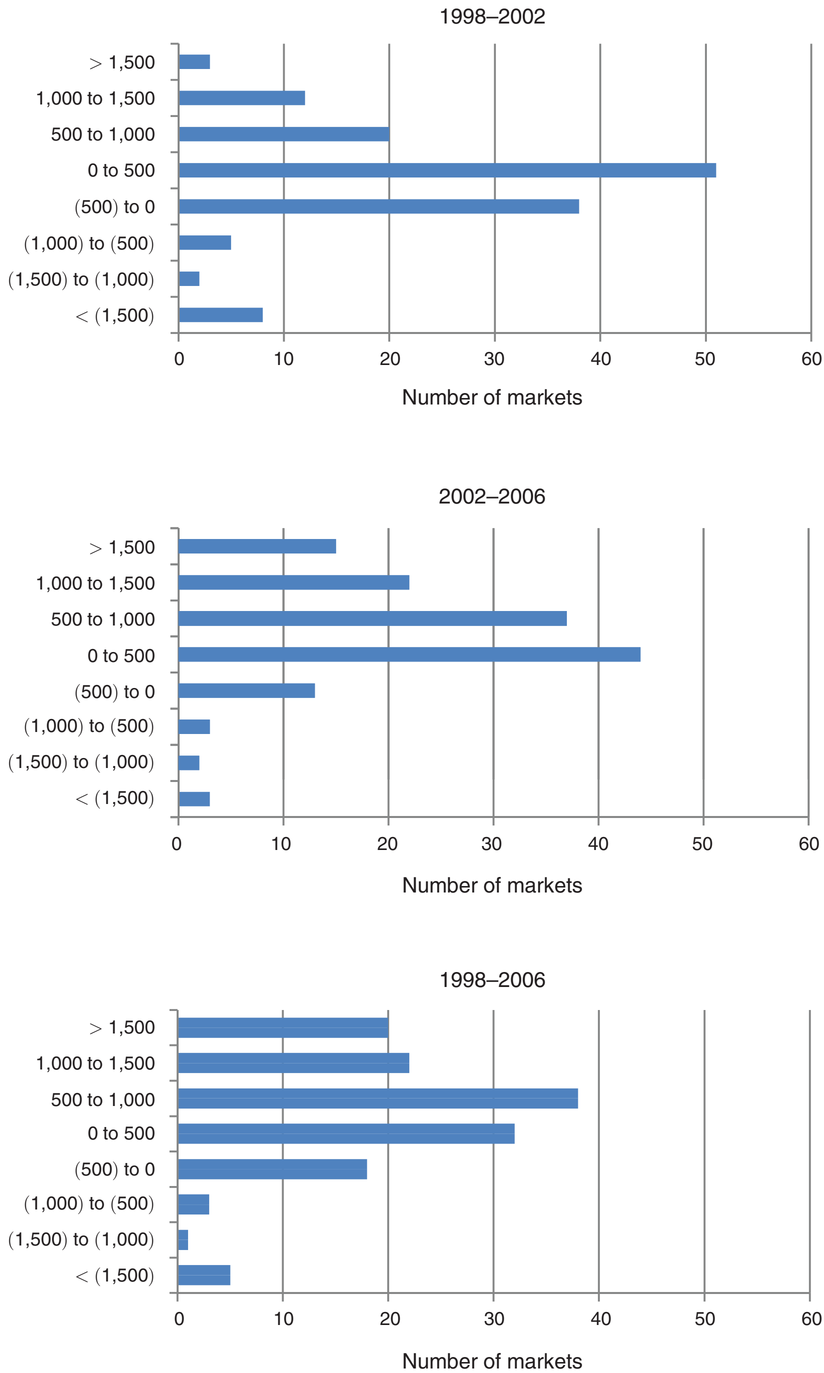

Figure 1. Change in Local Market Herfindahl

Note: HHI is scaled from 0 to 10,000.

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

To explore the relationship between premium growth and market concentration, we begin by estimating equations of the following form:

$$(1)\ \Delta \ln(premium)_{emt} = \alpha + \beta HHI_{mt-1} + \phi X_{mt-1} + \rho \Delta C_{emt} + \boldsymbol{\tau}_{\mathbf{t}} + \boldsymbol{\lambda}_{\mathbf{m}}$$
$$[+ \boldsymbol{\varsigma}_{\mathbf{e}}][+ \omega \Delta plan\ type\ shares_{emt} + \vartheta \Delta plan\ design_{emt}]$$
$$+ \varepsilon_{emt}.$$

In this specification, we model premium growth between year $t - 1$ and year t for a given employer e in market m as a function of lagged market characteristics (including HHI),[14] contemporaneous changes in observable characteristics of the insured population (such as demographics), and year and market fixed effects. Market characteristics are lagged by one year because premiums are set prospectively, i.e., premiums for 2006 are determined in 2005. In addition to HHI, the market-year covariates (denoted by X_{mt-1}) include the unemployment rate (to capture local economic conditions), the log of per-capita Medicare costs (to capture trends in health-care utilization), and the general, acute-care hospital HHI (to capture concentration in the provider market, which could independently lead to premium increases). Note these characteristics are included in level form (rather than first differences) to allow for a delayed response to changes in market structure or in local economic conditions.[15]

In contrast, we anticipate concurrent premium responses to changes in characteristics measured at the employer-market-year level (ΔC_{emt}), specifically demographic factors and the percentage of enrollees in self-insured plans. The year fixed effects capture average national changes in premium growth, and the market fixed effects capture differences in average growth rates across markets. Finally, we also estimate specifications including the terms in brackets: employer fixed effects, changes in the share of enrollees in each plan type, and changes in the average generosity of these plans.[16]

Results are presented in columns 1 through 3 of Table 2. There is no significant association between concentration levels and premium growth, and the estimates change little upon inclusion of additional controls.[17] Of course, causality can be inferred from this model only if within-market variation in insurer concentration is uncorrelated with other unobserved determinants of premiums, and if variation in premium growth does not induce variation in concentration. As previously noted,

[14] From a theoretical standpoint, HHI is a valid measure of competition if firms compete à la Cournot. While the Cournot model does not accurately describe the health insurance market, we follow the lead of most prior studies in the related literature, as well as the Horizontal Merger Guidelines, in adopting the HHI as a measure of competition.

[15] Given the inclusion of market fixed effects in equation (1), the coefficients on market-year covariates (including HHI) are identified by within-market changes in these variables.

[16] Note that employer fixed effects will substantially affect the coefficient on HHI only if employers with high or low growth in premiums are systematically located in markets that have high or low levels of HHI.

[17] The estimates are similarly small in magnitude and statistically insignificant if we use the change in HHI in place of the level of HHI as the key explanatory variable. For the most part, the coefficient estimates on the market-level control variables are statistically insignificant. The coefficient estimates on the employer-market controls are highly significant and generally have the expected signs. For example, a shift from 100 percent enrollment in POS plans (the omitted category) to 100 percent enrollment in HMO plans is associated with a 5 percent decline in premiums.

Table 2—Effect of Consolidation on Premiums (OLS Models)
(*Study Period: 1998–2006*)

	Dependent variable = annual change in ln(premiums)		
	(1)	(2)	(3)
Lagged HHI	0.002	−0.004	−0.005
	(0.010)	(0.009)	(0.009)
Market-year controls			
Lagged ln(Medicare costs per cap)	−0.015	−0.020	−0.019
	(0.016)	(0.016)	(0.016)
Lagged unemployment rate	0.118	0.147	0.158*
	(0.098)	(0.090)	(0.092)
Lagged hospital HHI	0.008	−0.003	0.001
	(0.016)	(0.015)	(0.015)
Employer-market controls			
Δ Demographic factor	0.303***	0.314***	0.311***
	(0.004)	(0.004)	(0.004)
Δ Fraction of self-insured employees	0.028***	0.032***	0.024***
	(0.005)	(0.005)	(0.004)
Δ Plan design			0.349***
			(0.022)
Δ Fraction in indemnity plans			0.085***
			(0.006)
Δ Fraction in HMO plans			−0.052***
			(0.006)
Δ Fraction in PPO plans			0.002
			(0.003)
Employer FE	No	Yes	Yes
Observations	66,906	66,906	66,906

Notes: The unit of observation is the employer-market-year. All specifications include market and year fixed effects. HHI is scaled from 0 to 1. Standard errors are clustered by market.
***Significant at the 1 percent level.
**Significant at the 5 percent level.
*Significant at the 10 percent level.

there are good reasons to doubt the validity of these assumptions. Hence, in the section that follows we pursue an instrumental variables approach.

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

In this section, we estimate the causal effect of changes in market concentration on premium growth by exploiting shocks to local market concentration produced by mergers and acquisitions (M&A).[18] Because M&A activity in local or regional markets may itself be motivated by expected trends in premium growth, we considered only large, nonlocal mergers as candidates for this analysis. We also ruled out mergers with insufficient pre or post periods (e.g., Aetna and NYLCare in 1998, the

[18] Our approach is similar in spirit to that of Hastings and Gilbert (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. They find that nonintegrated rival stations face higher costs, controlling for several time-varying station characteristics.

first year for which we have data), 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 remained: the Aetna-Prudential merger of 1999. Postmerger, the new firm (known as "Aetna") was widely reported to be the nation's largest insurer, covering 21 million individuals.[19] 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 sizable postmerger changes in market concentration. Online Appendix 2 provides additional discussion of the circumstances surrounding the merger. Importantly, there is no ex ante evidence that Aetna targeted Prudential because of expectations about premium growth or changes in insurer concentration in affected markets.

Our analysis is subdivided into four sections. First, we estimate the impact of the merger on market concentration (the "first stage" analysis). In so doing, we document the range of premerger market shares for Aetna and Prudential as well as the degree of premerger overlap. Second, we perform a reduced-form analysis, in which we examine the impact of the merger on premium growth. Third, we combine these analyses to produce our estimate of the causal impact of concentration on premiums. Last, we investigate the plausibility of alternative explanations for our findings. In particular, we estimate specifications to tease out the reaction of Aetna's *rivals*, as these responses are informative vis-à-vis the market dynamics.

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

Immediately prior to the merger in 1999, Aetna and Prudential were the third and fifth largest insurers in our sample 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 premerger shares of each firm. We hypothesize that markets served by both firms experienced increases in market concentration immediately following the merger, and that these increases varied by the premerger shares of the two merging firms. Specifically, for every market we calculate the "simulated change in HHI" ($sim\,\Delta HHI_m$) as the merger-induced *change* in market m's HHI that would have occurred from 1999 to 2000 absent any other changes, i.e.,

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

For example, if Aetna and Prudential had market shares of 10 percent each in 1999, $sim\,\Delta HHI_m$ (scaled by 10,000 as discussed above) would equal 200.

Figure 2 provides detail on the actual distribution of $sim\,\Delta HHI_m$ in the 139 LEHID markets. There is significant variation in this measure, with 46 largely unaffected markets ($sim\,\Delta HHI_m < 10$) and 42 highly affected markets ($sim\,\Delta HHI_m \geq 100$).

[19] Sanders, "Will the Aetna-Prudential Merger Hurt the Patient?" *TIME*, June 22, 1999.

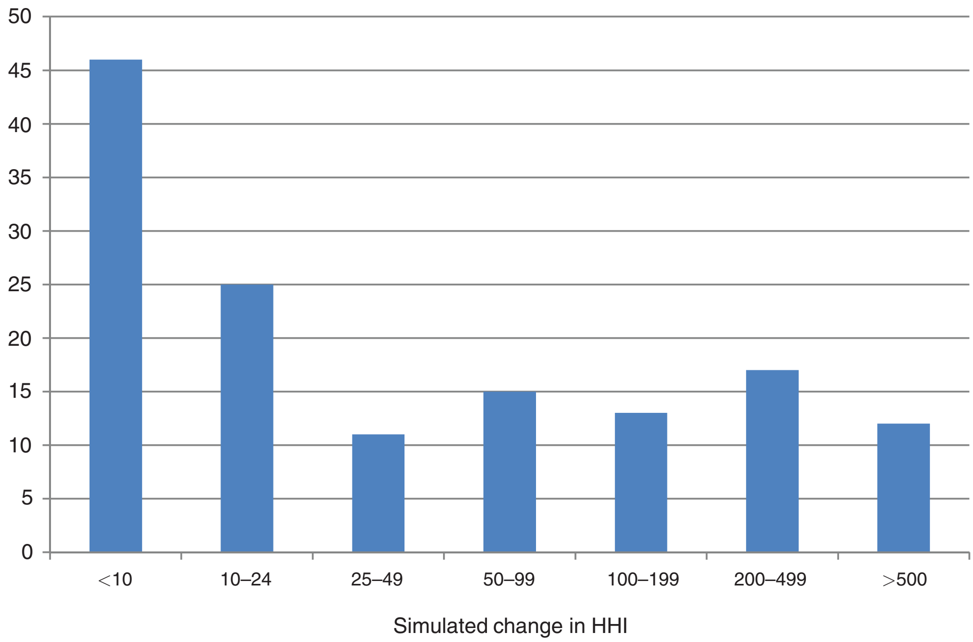

Figure 2. Distribution of Simulated Change in HHI Resulting from Aetna-Prudential Merger

Notes: $N = 139$. *HHI* is scaled from 0 to 10,000.

One state in particular stands out for its high levels of $sim\,\Delta HHI_m$: Texas. Five of the six markets in Texas have $sim\,\Delta HHI_m$ greater than 500. The high degree of overlap in Texas provoked action by the Department of Justice. To address the concerns raised by the Department, Aetna agreed to divest the Texas-based HMO businesses it had acquired from NYLCare in 1998.[20] We therefore examine whether the consent decree in Texas successfully neutralized the effect of the merger in these markets; to the extent it did, markets in Texas can serve as a "placebo" group for the natural experiment we study.

We propose to use $sim\,\Delta HHI_m \times post_t$ as an instrument for HHI in equation (1), where *post* is an indicator variable for the postmerger years in the sample. To evaluate this instrument, we estimate the following equation using market-year data, initially excluding observations from Texas:

$$HHI_{mt} = \alpha + \boldsymbol{\lambda}_{\mathbf{m}} + \boldsymbol{\tau}_{\mathbf{t}} + \beta sim\,\Delta HHI_m \times \boldsymbol{\tau}_{\mathbf{t}} + \varepsilon_{mt}. \tag{3}$$

[20] 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. DOJ stated that "The required divestitures … will preserve competition and protect consumers from higher prices" and "deny Aetna the ability to unduly depress physician reimbursement rates." See http://www.justice.gov/opa/pr/1999/June/263at.htm. Although the allegations pertained to Houston and Dallas, because Aetna divested all NYLCare plans in Texas, the consent decree affected the entire state. Source: "Blue Cross and Blue Shield of Texas to Purchase NYLCare Texas Operations," Aetna press release, 9/14/1999, http://www.aetna.com/news/1999/pr_19990914.htm.

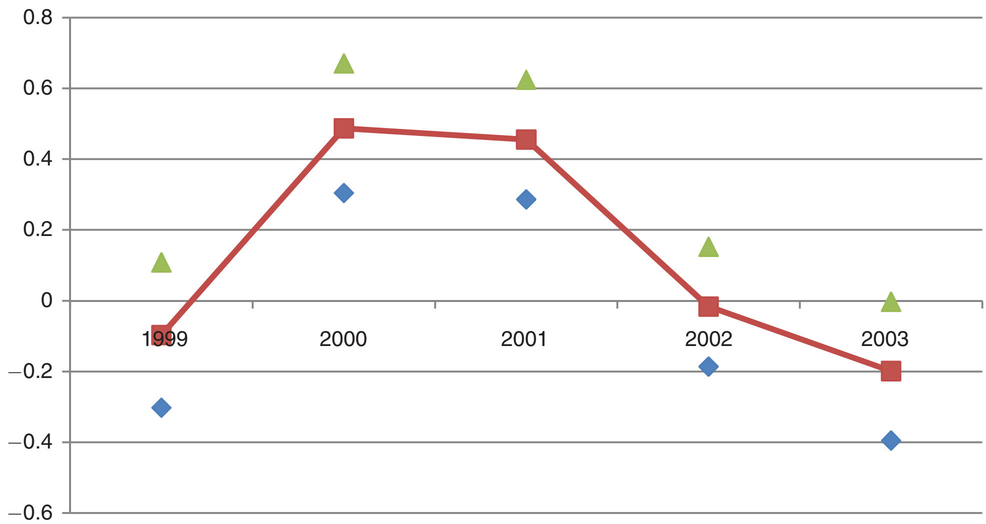

Figure 3. Estimated Coefficients and 95 percent Confidence Intervals from Regression of HHI on Simulated Change in HHI

Note: Coefficients and standard errors are reported in column 1 of Table 3.

The vectors denoted by λ_m and τ_t represent a full set of market and year fixed effects, respectively. By interacting $sim\,\Delta HHI_m$ with separate indicators 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 estimated coefficients will also help to determine the appropriate study period for our analysis. In this and all specifications including $sim\,HHI_m$, we use a scale of 0 to 1 for this measure.

Figure 3 graphs the coefficient estimates on the yearly interactions with $sim\,\Delta HHI_m$, together with the 95 percent confidence intervals. The sample includes data from 1998 to 2003. Estimates are presented in numerical form in column 1 of Table 3. Relative to the omitted interaction term, $sim\,\Delta HHI_m \times (year = 1998)$, only the interactions with indicators for 2000 and 2001 are statistically significant. At −0.10, the coefficient estimate for β in 1999 is small and (insignificantly) negative, whereas estimates for β in 2000 and 2001 are large (0.49 and 0.46, respectively) and significant at the 5 percent level. The timing is consistent with expectations: the merger was effectively cleared in July 1999, when the Department of Justice submitted its Proposed Final Judgment. The coefficients in 2000 and 2001 are significantly smaller than 1, implying that employers to some extent substituted away from Aetna and Prudential in the wake of the merger. In addition, there is likely attenuation bias due to measurement error, as we have only a sample (rather than a census) of insurance contracts.

The coefficient estimates of β in 2002 are 2003 are both noisy and negative indicating that the merger-induced shocks to local concentration dissipated quickly.[21] In

[21] This finding is consistent with reports from industry experts. According to a 2004 *Health Affairs* article by Robinson, "[G]ossip speculates [Aetna] would be lucky to still have 30,000 of the 5 million it acquired from Prudential."

Table 3—Effect of the Aetna-Prudential Merger on Market Concentration
(*Dependent Variable* = *HHI*)

	1998–2003 (1)	1998–2001 (2)	1998–2001 (3)	1998–2001 (4)
$Sim\,\Delta\,HHI$ × (year = 1999)	−0.097 (0.180)			
$Sim\,\Delta\,HHI$ × (year = 2000)	0.487** (0.204)			
$Sim\,\Delta\,HHI$ × (year = 2001)	0.455** (0.194)			
$Sim\,\Delta\,HHI$ × (year = 2002)	−0.017 (0.205)			
$Sim\,\Delta\,HHI$ × (year = 2003)	−0.199 (0.248)			
$Sim\,\Delta\,HHI$ × (year >= 2000)		0.520*** (0.166)	0.512*** (0.165)	0.520*** (0.166)
$Sim\,\Delta\,HHI$ × (year >= 2000) × (Texas == 1)			−0.646*** (0.224)	−1.262*** (0.291)
Texas × (year >= 2000)				0.052 (0.037)
Texas included?	No	No	Yes	Yes
Observations	798	532	556	556
R^2	0.677	0.674	0.677	0.678

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

***Significant at the 1 percent level.
**Significant at the 5 percent level.
*Significant at the 10 percent level.

order to use the merger as an instrument for market concentration, we must therefore focus our analyses on the early years of our sample: 1998–2001 for the first-stage model, and 1998–2002 for the second stage (because HHI impacts premiums with a lag). However, in Section IIIB below, we discuss reduced-form analyses of the longer-term impact of changes in simulated HHI on health insurance premiums by extending the study period out to 2006.

Next, we use data from 1998 through 2001 to estimate a more parsimonious model that replaces the individual year interactions with a single "post" indicator that takes a value of one during 2000 and 2001:

$$(4)\quad HHI_{mt} = \alpha + \boldsymbol{\lambda}_{\mathbf{m}} + \boldsymbol{\tau}_{\mathbf{t}} + \beta_0 sim\,\Delta HHI_m \times post_t + [\beta_1 sim\,\Delta HHI_m \times post_t \times Texas_m] + [\psi post_t \times Texas_m] + \varepsilon_{mt}.$$

After estimating the baseline model (which excludes the terms in brackets), we add the six Texas markets to the sample and include a triple-interaction,

$sim\,\Delta HHI_m \times post_t \times Texas_m$, to explore whether the post-merger impact of $sim\,\Delta HHI$ differs in these markets. We then add the term $post_t \times Texas_m$ to control for average changes in Texas as compared to other states during the post period, although it may be difficult to separately identify the coefficient on the two Texas interactions because there are only six Texas markets and two post years.

The results are displayed in column 2 of Table 3. As anticipated, the coefficient on $sim\,\Delta HHI_m \times post_t$ is statistically significant: 0.52, with a standard error of 0.17. The results in columns 3 and 4 show 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 in both specifications and fully offsets the impact of the merger. In both models, we cannot reject the hypothesis that the sum of the relevant double- and triple-interaction terms equals zero. Observations from Texas are therefore suitable for the placebo test (or falsification exercise) previously noted. If premium growth has a similar relationship with $sim\,\Delta HHI$ in Texas as in other parts of the United States, then changes in insurer concentration may not be driving the observed relationship.

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

To investigate the effect of merger-induced increases in local market concentration on plan premiums, we estimate models of the following form:[22]

$$(5)\quad \Delta \ln(premium)_{emt} = \alpha + \kappa_0 sim\,\Delta HHI_m \times post_t + \phi X_{mt-1} + \rho \Delta C_{emt} + \boldsymbol{\tau_t} + \boldsymbol{\lambda_m}[+\boldsymbol{\varsigma_e}][+\omega\,\Delta plan\ type\ shares_{emt} + \vartheta\,\Delta plan\ design_{emt}] [+\kappa_1 sim\,\Delta HHI_m \times post_t \times Texas_m] [+\gamma post_t \times Texas_m] + \varepsilon_{emt}.$$

In light of the results from the preceding section, we focus on the period between 1998 and 2002 (i.e., annual premium growth from 1998–1999, 1999–2000, 2000–2001, and 2001–2002). Note that in this model $post_t$ takes a value of one for the 2000–2001 and 2001–2002 changes, and is otherwise equal to zero.[23] As in the OLS regressions presented in Section II, we begin with a parsimonious specification that controls for lagged market covariates and changes in employer-market characteristics, as well as fixed differences across years and markets in average premium *growth* (captured respectively by year and market fixed effects, denoted τ_t and λ_m).

The results are reported in column 1 of Table 4. The estimated coefficient on $sim\,\Delta HHI_m \times post_t$ is positive and statistically significant. Given the mean

[22] In a companion set of specifications (results available upon request), we define the outcome variable to be ln(premium) (rather than the change in this measure) and include market time trends. The results are similar to those presented in this section.

[23] Recall the last year of the merger-induced HHI increase was 2001, and premiums for 2002 are set in 2001.

Table 4—Merger Effects on Premiums
(*Study Period: 1998–2002*)

	Dependent variable = annual change in ln(premiums)					
	(1)	(2)	(3)	(4)	(5)	(6)
$Sim\,\Delta\,HHI$ × (year >= 2001)	0.177***	0.202***	0.186***		0.193***	0.188***
	(0.056)	(0.048)	(0.050)		(0.049)	(0.049)
$Sim\,\Delta\,HHI$ × (year = 2000)				0.011		
				(0.061)		
$Sim\,\Delta\,HHI$ × (year = 2001)				0.181**		
				(0.071)		
$Sim\,\Delta\,HHI$ × (year = 2002)				0.200***		
				(0.067)		
$Sim\,\Delta\,HHI$ × (year >= 2001)					−0.238***	−0.056
× (Texas = 1)					(0.069)	(0.191)
(Year >= 2001) × (Texas = 1)						0.016
						(0.017)
Market-year controls						
Lagged ln(Medicare costs per cap)	−0.029	−0.047	−0.039	−0.040	−0.046	−0.048
	(0.034)	(0.036)	(0.036)	(0.036)	(0.033)	(0.033)
Lagged unemployment rate	0.479***	0.579***	0.567***	0.570***	0.575***	0.535***
	(0.174)	(0.161)	(0.155)	(0.155)	(0.152)	(0.152)
Lagged hospital HHI	0.003	−0.004	0.003	0.003	−0.010	−0.008
	(0.022)	(0.021)	(0.020)	(0.020)	(0.020)	(0.020)
Employer-market controls						
Δ Demographic factor	0.304***	0.328***	0.323***	0.323***	0.324***	0.324***
	(0.006)	(0.005)	(0.006)	(0.006)	(0.005)	(0.005)
Δ Fraction of self-insured	0.048***	0.054***	0.019***	0.019***	0.017***	0.017***
employees	(0.007)	(0.007)	(0.006)	(0.006)	(0.006)	(0.006)
Δ Plan design			0.223***	0.223***	0.210***	0.211***
			(0.040)	(0.040)	(0.040)	(0.040)
Δ Fraction in indemnity plans			0.089***	0.089***	0.091***	0.091***
			(0.008)	(0.008)	(0.008)	(0.008)
Δ Fraction in HMO plans			−0.081***	−0.081***	−0.084***	−0.084***
			(0.009)	(0.009)	(0.009)	(0.009)
Δ Fraction in PPO plans			0.000	0.000	−0.001	−0.001
			(0.006)	(0.006)	(0.006)	(0.006)
Employer FE	No	Yes	Yes	Yes	Yes	Yes
Texas observations included?	No	No	No	No	Yes	Yes
Observations	28,645	28,645	28,645	28,645	30,493	30,493

Notes: The unit of observation is the employer-market-year. All specifications include market and year fixed effects. HHI is scaled from 0 to 1. Standard errors are clustered by market.
*** Significant at the 1 percent level.
** Significant at the 5 percent level.
* Significant at the 10 percent level.

$sim\,\Delta HHI_m$ of 0.014 (across all 139 geographic markets), the point estimate of 0.177 implies that, in a typical market, the merger induced an average premium increase of approximately 0.25 percent in both 2001 and 2002, and thus a total increase of approximately 0.50 percent. The point estimate changes little upon inclusion of employer fixed effects (column 2), and as expected the standard errors decrease. Adding controls for changes in the generosity of plans (column 3) also has little impact on the estimate.

Next, we study the pattern of premium growth over time by replacing the term $sim\,\Delta HHI_m \times post_t$ with $sim\,\Delta HHI_m \times \tau_t$ (interactions with individual year dummies, with 1998 as the omitted year). The results, in column 4, provide two key insights. First, there is no evidence of a "pretrend" in premium growth; that is, the estimated reaction to the merger is not due to a premerger trend in markets with large overlapping Aetna and Prudential market shares. Second, the effect of the merger on premium growth is very similar in both "post" years.

This finding strongly suggests that the impact of the merger is appropriately modeled, i.e., that concentration affects the growth rate rather than the level of premiums.[24] If the sample is extended to 2006, we find the coefficients remain of similar magnitude for two more years, and then fall down close to zero.[25] The fact that the coefficient estimates remain positive and do not become negative suggests some amount of hysteresis: consolidation results in a higher rate of premium growth, and even when circumstances change (in this case, the effect of the merger on concentration eventually disappeared) premiums remain elevated.[26]

Columns 5 and 6 of Table 4 present 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, $sim\,\Delta HHI_m \times post_t \times Texas_m$.[27] The estimated coefficient on this term is highly significant and negative (-0.24) and almost perfectly offsets the main effect of $sim\,\Delta HHI_m$ in this specification (0.19). Although the result is not robust to including a separate term for $post_t \times Texas_m$ (column 6), this is not surprising given there are only six markets in Texas and just two post years. On net, the results suggest that the market power effect of the merger in Texas was indeed neutralized by the DOJ's actions.[28]

C. *IV Estimates*

Table 5 presents the first-stage, reduced-form, and second-stage models corresponding to our IV estimate; the reduced-form model is repeated from column 3 of Table 4. At 0.39, the estimated effect of lagged HHI on premium growth is positive, statistically significant, and roughly twice as large as the reduced-form estimate. This is anticipated given the first-stage coefficient of 0.48 reported in column 1.[29]

Because our estimates suggest that changes in HHI affect the growth *rate* (rather than just the level) of premiums, to estimate the average effect of consolidation over the entire study period, we must consider the timing of consolidation between 1998

[24] An alternative explanation is that an increase in concentration does raise the level (rather than the growth rate) of premiums, but it takes multiple years to reach the new level.

[25] To be precise, the coefficients on interactions of the simulated change in HHI with indicators for 2003 and 2004 are 0.293 and 0.203 respectively, and are both significant with $p < 0.01$.

[26] As noted earlier, the results of the first stage necessitate a study period ending in 2002. However, the results just described suggest the estimates will be conservative.

[27] Note a second-order interaction (i.e., $post_t \times Texas_m$) is arguably not necessary in this model as market fixed effects already control for differences in average annual growth rates across markets.

[28] As an additional 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 $sim\,\Delta HHI_m \times post_t \times initial\ HHI_m$ (and variants thereof) were very imprecise.

[29] Note this first-stage coefficient differs slightly from the coefficient obtained using market-year data, as the unit of observation is the employer-market-year.

Table 5—The Impact of HHI on Premiums
(*Study Period: 1998–2002*)

	Dep var = lagged HHI	Dep var = annual change in ln (premium)		
	First-stage estimates	Reduced-form estimates	IV estimates	OLS estimates
Sim Δ *HHI* × (year >= 2001)	0.475***	0.186***		
	(0.014)	(0.050)		
Lagged HHI			0.391***	0.015
			(0.130)	(0.018)
Market-year controls				
Lagged ln(medicare costs per cap)	0.034**	−0.039	−0.052	−0.018
	(0.014)	(0.036)	(0.037)	(0.034)
Lagged unemployment rate	0.204***	0.567***	0.488***	0.474***
	(0.048)	(0.155)	(0.163)	(0.162)
Lagged hospital HHI	−0.060***	0.003	0.026	0.006
	(0.007)	(0.020)	(0.022)	(0.020)
Employer-market controls				
Δ Demographic factor	0.004***	0.323***	0.321***	0.323***
	(0.001)	(0.006)	(0.004)	(0.003)
Δ Fraction of self-insured employees	0.000	0.019***	0.019***	0.019***
	(0.001)	(0.006)	(0.004)	(0.004)
Δ Plan design	0.019*	0.223***	0.216***	0.222***
	(0.010)	(0.040)	(0.028)	(0.027)
Δ Fraction in indemnity plans	0.001	0.089***	0.089***	0.089***
	(0.002)	(0.008)	(0.005)	(0.005)
Δ Fraction in HMO plans	−0.003	−0.081***	−0.080***	−0.081***
	(0.002)	(0.009)	(0.006)	(0.006)
Δ Fraction in PPO plans	0.001	0.000	0.000	0.001
	(0.002)	(0.006)	(0.005)	(0.005)
Observations	28,645	28,645	28,645	28,645

Notes: The unit of observation is the employer-market-year. All specifications include employer, market, and year fixed effects. HHI is scaled from 0 to 1. Standard errors are clustered by market.
***Significant at the 1 percent level.
**Significant at the 5 percent level.
*Significant at the 10 percent level.

and 2006. As previously noted, the average increase in HHI across all markets was 698 points during this period. If this increase were evenly distributed over time, the effect of consolidation on premiums during our study period would be approximately 13 percent. However, consolidations tended to occur later in the study period, yielding a cumulative estimated effect of approximately 7 percent.[30]

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 coefficient from the OLS model (presented in column 4) is near zero (and imprecisely estimated).

[30] Details of our calculation are available in online Appendix 3. If one assumes that an increase in concentration between t and $t + 1$ affects premium growth for only two years (i.e., until $t + 3$, rather than indefinitely), then the implied increase in premiums caused by the increase in HHI between 1998 and 2006 is somewhat lower at 5 percent.

Hausman specification tests reject the null assumption of consistency for this model ($p < 0.01$), underscoring the need for instrumental variables estimation.

Collectively, the results presented in this section show that consolidation does result in a "premium 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. Two key results indicate our conclusions are not driven by unobserved factors correlated with the pre-merger market shares of Aetna and Prudential. First, there is no evidence that concentration *or* premiums in markets with higher $sim\,\Delta HHI$ were trending differently before the merger took effect. Second, we find no 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 $lagged\ HHI_m$. In online Appendix 4, we examine the impact of consolidation on health plan characteristics other than price, such as *plan design* and the share of employees enrolled in HMOs. [31]

D. *Alternative Explanations*

The findings summarized above are consistent with the exercise of market power in the wake of consolidation. However, the pattern of results is also consistent with alternative explanations, in particular a "mistake" in Aetna's postmerger pricing strategy, and/or increases in insurance quality (and therefore price). In this section, we discuss the evidence with regard to these alternative hypotheses.

Our results show that prices increase *on average* in markets with higher $sim\,\Delta HHI_m$. If this price increase is primarily due to actions by Aetna, then Aetna's subsequent loss of market share would suggest the price increase was unsuccessful, i.e., they were not able to exercise market power following the merger. On the other hand, if competitors followed suit by increasing their prices as well, that would suggest that Aetna's action softened competition marketwide, implying the presence (and exercise) of market power.

To investigate whether Aetna's competitors increased their premiums in response to the merger, we estimate a set of specifications analogous to those in Table 4 for the 61 percent of employer-markets that were *not served* by either Aetna or Prudential at the time of the merger in 1999. Our point estimates for the coefficient of particular interest (κ_0 from equation (5)) are similar to the estimates for the full sample, as shown in online Appendix 5. This implies that insurers not directly involved in the merger responded to the merger-induced change in concentration by raising their premiums, which supports the market-power explanation for our findings.

Importantly, when we restrict the sample to employer-markets that *were served* (either partially or fully) by Aetna or Prudential at the time of the merger, our estimates for κ_0 are approximately twice as large. This suggests that the merged entity increased its premiums more than its competitors in markets where Aetna and Prudential had significant overlap, which is consistent with the merged entity

[31] Among other results, we find that employers reduced the generosity of *plan design*. This is consistent with efforts by employers to reduce the burden of higher insurance premiums through so-called "benefit buybacks." We emphasize that our premium results *do* control for changes in plan design. We find a somewhat counterintuitive shift *away* from HMOs; however, we discuss plausible explanations for this pattern in online Appendix 4.

exercising price leadership and its oligopolistic rivals following. Last, it is notable that premiums remained elevated in high-$sim\,\Delta HHI$ markets through at least 2006, notwithstanding Aetna's loss of market share by 2002. This hysteresis in market price is again consistent with a new oligopolistic pricing equilibrium facilitated by Aetna's original exercise of market power.

The second alternative explanation, that Aetna raised quality and competitors followed its lead, is less amenable to exploration using our data. Conceptually, there are at least two reasons to question this hypothesis. First, quality is "lumpy" (e.g., enhancing consumer access to claims) and far more difficult to calibrate across different markets than price. Second, quality changes take time to implement and to communicate to the marketplace, and the impact of the merger on price occurs within the first year. These points notwithstanding, quality remains an important omitted factor in our analysis.

IV. 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 degree of competition in the insurance industry will also potentially affect upstream suppliers, such as health-care providers, pharmaceutical firms, and medical device manufacturers. To the extent that suppliers have few outside options, a lack of vigorous competition among insurers may lead to monopsonistic practices. Capps (2010) reviews the theoretical and practical implications of monopsony in the context of health insurance mergers.[32]

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. In fact, the DOJ's formal complaint regarding the Aetna-Prudential merger alleged that 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."[33]

In this section, we consider the possibility that consolidation facilitates the exercise of monopsony power by estimating the relationship between our instrument for HHI ($sim\,\Delta HHI_m$) and both the employment (or "quantity") and average compensation (or "price") of health-care 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 orthogonal to other determinants of employment and compensation *growth*, our results can be interpreted as causal estimates of the impact of consolidation on these outcomes.

To execute this analysis, we supplemented the LEHID data with the Occupational Employment Statistics (OES) survey on income and employment in

[32] A number of recent studies examine the effect of insurer bargaining power on hospital prices, including Feldman and Wholey (2001), Sorensen (2003), Moriya, Vogt, and Gaynor (2010), and Ho (2009).

[33] See Complaint, *United States vs. Aetna Inc*. (ND TX, 21 June 1999). More recently, the DOJ required a similar divestiture before approving a 2005 merger between United Health Group Inc. and Pacificare Health Systems Inc. Both divestitures were driven by concerns about the effect on physician services in specific markets (see Complaint, *United States vs. UnitedHealth Group Inc. and Pacificare Health Systems Inc.*, Dec 20, 2005).

health care–related occupations. We restrict our attention to the 43 occupation categories that are classified under the Standard Occupational Classification (SOC) system as "Healthcare Practitioner and Technical Occupations." These include 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 pertaining to physicians and the two for nurses.[34] Nurses are by far the largest group, accounting for 56 percent of personnel in our sample; pharmacists are second (4.3 percent), and physicians are a close third (4.2 percent). Additional details, including descriptive statistics for our sample, are available in online Appendix 6.

The unit of observation for the OES data (as well as our analysis) is the occupation-MSA-year and the variables of interest are the mean annual wage and estimated employment. Using a crosswalk that matches LEHID markets to MSAs, we merge this data with our measures of insurer concentration (including our instrument). We estimate parsimonious specifications using the change in log average earnings or employment between 1999 and 2002 as the dependent variable, and $sim\,\Delta HHI$ as the main predictor:

$$\begin{aligned}(6)\quad \Delta \ln y_{os,99-02} = {} & \alpha + \gamma sim\,\Delta HHI_s + \omega Physician_o \times sim\,\Delta HHI_s \\ & + \vartheta Nurse_o \times sim\,\Delta HHI_s \\ & + \varsigma Physician_o + \theta Nurse_o + \upsilon \Delta HospitalHHI_s \\ & + [\Delta \ln y_{os,97-98} + \boldsymbol{\varsigma}_{\mathbf{o}}] + \varepsilon_{os}.\end{aligned}$$

The subscripts o and s 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 $sim\,\Delta HHI_s$. The indicators capture differences in earnings and employment growth for each category (relative to other health-care occupations), 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 in each market. As specification checks, we progressively add each of the terms in brackets. The first term, $\Delta \ln y_{os,97-98}$, represents the change in earnings or employment between 1997 and 1998 and serves as a control for preexisting 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 MSA.

[34] 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 during our study period. The "Nurses" category includes Registered Nurses (RNs) and Licensed Vocational Nurses (LVNs).

Table 6—Effect of the Aetna-Prudential Merger on Health-care Provider Earnings and Employment

	Dep var = Δ log (average income) from 99–02; mean = 0.121			Dep var = Δ log (employment) from 99–02; mean = 0.191		
Simulated Δ *HHI*	0.111	0.078	0.091	−2.372***	−2.723***	−2.437**
	(0.180)	(0.215)	(0.204)	(0.809)	(0.941)	(0.978)
Physician indicator	0.193***	0.184***	NA	0.523***	0.497***	NA
	(0.034)	(0.035)		(0.170)	(0.167)	
Physician × *simulated* Δ *HHI*	−2.007**	−2.180***	−2.195***	−2.507	−2.582	−2.858
	(0.833)	(0.801)	(0.811)	(7.934)	(8.441)	(8.439)
Nurse indicator	−0.013**	−0.015**	NA	−0.154***	−0.160***	NA
	(0.006)	(0.006)		(0.025)	(0.027)	
Nurse × *simulated* Δ *HHI*	0.440**	0.471*	0.457*	1.707**	2.012*	1.738*
	(0.221)	(0.257)	(0.254)	(0.845)	(1.071)	(1.032)
Δ *Hospital HHI*, 1999–2002	0.023	0.021	0.024	−0.024	−0.027	−0.067
	(0.029)	(0.031)	(0.032)	(0.254)	(0.247)	(0.235)
Trend in dep var, 1997–1998	No	Yes	Yes	No	Yes	Yes
Occupation fixed effects	No	No	Yes	No	No	Yes
Observations	2,110	1,631	1,631	2,110	1,631	1,631

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

***Significant at the 1 percent level.
**Significant at the 5 percent level.
*Significant at the 10 percent level.

The results are presented in Table 6. 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 $sim\,\Delta HHI_s$ in columns 1 through 3 is positive but imprecisely estimated, implying no significant impact of the merger on average earnings across all health-care occupations. The coefficient on the physician indicator in columns 1 and 2 demonstrates that physicians experienced an increase of around 21 percent in average nominal earnings between 1999 and 2002 (relative to nonnursing health-care personnel). However, the coefficient estimate on $Physician_o \times sim\,\Delta HHI_s$ is negative and significant in all models, revealing that earnings growth for physicians was lower in markets affected by the merger. Given the average value of 0.014 for $sim\,\Delta HHI_s$, the point estimate implies that the merger restrained growth in physician earnings by approximately 3 percent in a typical market. The coefficient on the nurse indicator reveals that nurses experienced a small decrease in relative 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 premerger overlap (by approximately 0.6 percent in the typical market).

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 health-care occupations, employment of physicians increased, while that of nurses decreased, during the study period. The point estimate on $sim\,\Delta HHI_s$ is negative and significant: in a typical market, the merger led to a drop in health care–related employment of 2.7 percent. The interaction between the physician indicator and $sim\Delta HHI_s$ is negative but noisily estimated, whereas the interaction between the nurse indicator and $sim\Delta HHI_s$ is large, positive, and marginally significant. The relative increase in nurse employment in geographic markets differentially affected by the merger suggests there was some substitution toward nurses in these markets. This explanation is buttressed by the earnings regressions, which found the merger depressed growth in physicians' earnings while modestly boosting nurses' earnings.[35]

To summarize, we find that increases in market concentration predicted to occur in the wake of the Aetna-Prudential merger resulted in pronounced declines in health care–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).[36] The evidence suggests that market power facilitates the substitution of nurses for physicians. The results are also consistent with the exercise of monopsony power by insurers vis-à-vis physicians, as their relative earnings and employment growth declined most in markets with the largest predicted merger impact. Paired with the findings of the previous section, we conclude that in markets where Aetna and Prudential had substantial premerger overlap, insurers were able to exercise market power simultaneously in input *and* output markets postmerger. Thus, the premium increases documented in the previous section likely understate the increase in insurer profits due to consolidation.

V. Discussion and Conclusions

The scope of the private health insurance industry is difficult to overstate. More than 170 million nonelderly Americans are privately insured, and this figure does not include the millions of publicly insured individuals whose coverage is outsourced to private insurers. The recent health insurance reform legislation will further expand the reach of this industry, with the Congressional Budget Office projecting an increase of 16 million in the number with private primary insurance by 2019 (CBO 2010). In addition, the annual growth in employer-sponsored health insurance premiums has exceeded the annual growth in earnings by a factor of seven during the last several years (Romer and Duggan 2010).[37] In this study, we investigate whether and to what extent increasing consolidation in the US health insurance industry is responsible for this rapid growth in premiums.

We arrive at four main conclusions. First, most Americans live in markets served 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

[35] As a robustness check, we estimated all models using 1999–2001 as the study period, as the BLS changed its methodology for constructing mean wages in 2002 (see online Appendix 6). Our findings are qualitatively similar.

[36] We also estimated specifications subdividing the nurse category into two large subgroups (Registered Nurses —RNs and Licensed Vocational Nurses—LVNs). We find that only RNs earned higher relative raises in markets where the merger had most impact. LVNs enjoyed significant relative employment gains, whereas the employment gains for RNs were not statistically significant (although they are of similar magnitude). On the whole, the results are consistent with outward shifts of demand for both nursing types, with a less-elastic short-run supply curve for RNs. The results from these specifications are presented in online Appendix 6. We thank an anonymous referee for this suggestion.

[37] Data from the BLS "Employer Costs for Employee Compensation" survey indicate that workers' real average hourly wage and salary income increased by 0.7 percent annually from 2000 to 2009. During that same period, the growth rate in ESI premiums was substantially higher at 5.1 percent per year (Romer and Duggan 2010).

concentrated" category (per the DOJ'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 plan characteristics.

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 that the average market-level changes in HHI between 1998 and 2006 resulted in a premium increase of approximately 7 percentage points by 2007, ceteris paribus. Given our sample includes both fully and self-insured plans, and insurers have less control over pricing of the latter, it is plausible that consolidation is associated with an even larger impact on fully insured plans, which are dominant in the individual and small group markets.

Fourth, we find evidence that consolidation reduces the employment of health-care workers and may facilitate 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 3 percent and raised nurse earnings by 0.6 percent. The magnitude of this effect was higher (lower) in markets where the premerger shares of the two companies overlapped more (less). Thus, the results imply that insurers exercised monopsonistic power against physicians in some markets during the period 1998–2002.

Our findings indicate that Americans are indeed paying a premium on their health insurance premiums as a result of recent increases in market concentration of the health insurance industry. However, consolidation explains only a fraction of the steep increase in premiums in recent years. While 7 percent is large in absolute terms (it translates into approximately $34 billion in extra annual premiums), and large relative to operating margins of insurers, it is only one-eighth of the increase in average, inflation-adjusted premiums observed in our sample during the same 1998 to 2006 time period.[38]

We caution that our analysis relies on a single merger whose substantial effects on market concentration persisted for just two years. However, it is among the largest mergers to date in the health insurance industry, and one with differential impacts across 139 geographic markets in the United States (implying 139 small experiments). Additional research that utilizes other plausibly exogenous sources of variation in market structure would be valuable to assessing conduct in this important industry. We also emphasize that our sample consists primarily of large, multisite firms, and the results may not be generalizable to all market segments, including the small group and individual markets.[39] Finally, there has also been a great deal of consolidation across (as opposed to within) markets, the effects of which are not reflected in our estimates.

[38] As shown in Table 1, average premiums in our sample increased from $4,104 in 1998 to $7,832 in 2006. Adjusting these both to 2007 dollars yields an increase in average, inflation-adjusted premiums of 54 percent. The $34 billion figure is based on an estimated $490 billion in total private insurance premiums in the United States as of 1998 (CMS 2011). The aggregate effect of consolidation on profits should be larger as the "premium on premiums" does not incorporate reductions in provider payments obtained through the exercise of monopsony power.

[39] High and increasing concentration has also been documented in the individual/small group market (GAO 2009b).

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