Distributional Effect of Insurer Competition on Premiums: Evidence from the Employer-Sponsored Health Insurance Market*

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Abstract

We study how insurer competition affects premiums for fully-insured plans in the employer-sponsored health insurance market. By utilizing quasi-experimental variations induced by mergers among national insurers, we show that increased insurer concentration leads to higher premiums, but there exists substantial heterogeneity. In particular, we find that a 10% increase in insurer concentration increases premiums for mid-sized firms by 1.2% but reduces premiums for large firms by 3.5%. We explore additional sources of heterogeneity to shed light on the underlying mechanisms and conclude by discussing policy implications.

JEL Codes: I11, I13, L11

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1 Introduction

In the United States, approximately 60% of the non-elderly population is covered by employer-sponsored health insurance (hereafter ESI) [Kaiser Family Foundation (2020b)]. However, ESI premiums have more than tripled over the last two decades from $\$2,196$ in 1999 to $\$7,188$ in 2019 for single coverage, and from $\$5,791$ to $\$20,576$ for family coverage [Kaiser Family Foundation (2019)]. Such an increase in premiums represents an average growth rate of more than 6% annually, outpacing wage growth over the same period. Accompanying this premium growth is the increased insurer concentration, partly driven by a recent wave of mergers and acquisitions.[1] According to the American Medical Association, around 50% of the 40 largest Metropolitan Statistical Areas (MSAs) were considered highly concentrated in 2001. By 2018, this number had risen to 96% among all 383 MSAs.[2]

This article examines the role of insurer competition in determining health plan premiums in the fully-insured ESI market. Theoretically, the effect of insurer competition on premiums is ambiguous [Dafny et al. (2012); Ho and Lee (2017)]. On the one hand, the market power effect suggests that a higher level of market concentration enables insurers to exercise their market power and raise premiums in the downstream market. On the other hand, the purchase/countervailing power effect allows insurers to take advantage of their enhanced bargaining leverage in negotiating lower payment rates with healthcare providers in the upstream market,[3] and some of the cost savings could be passed to consumers in the form of lower premiums. As a result, the net effect of insurer competition on premiums remains an empirical question and depends on the relative strengths of these two offsetting effects.

To study the effect of insurer competition, we take advantage of a novel dataset, Form 5500 (Annual Report required by the U.S. Department of Labor), which provides detailed plan-level information on premiums, enrollment, and insurers for the near-universe of ESI offerings.[4] The enrollment information helps us attribute local market shares for the calculation of the Herfindahl-Hirschman Index (HHI here-

[1] Drawing from the Irving Levin Associates Reports, around 400 mergers occurred in the managed care market from 2000 to 2014.

[2] Markets with the Herfindahl-Hirschman Index larger than 1,800 were considered highly concentrated, based on the *Merger Guidelines* issued by the Department of Justice and Federal Trade Commission in 1997 and 2023.

[3] Other offsetting effects could also exist. For example, mergers could lead to cost savings for consolidated parties due to efficiency gains.

[4] To the best of our knowledge, a few other studies have used this data source, including Bova et al. (2019), Dalton and Holland (2019), Craig (2022), and Tong (2024).

after). As this measure is known to be susceptible to endogeneity concerns, we follow Dafny et al. (2012) to construct an instrument (a simulated HHI) by utilizing quasi-experimental variations in the level of local market competition created by mergers between national insurers. As the empirical variation in HHI is essentially driven by changes in competition, not only can our study reveal the impact of insurer mergers retrospectively, but it can also directly inform antitrust assessment of potential mergers and acquisitions in which the use of HHI is economically sensible and well-founded [Miller et al. (2021)].

We first assess the aggregate effect of increased insurer concentration to quantify the extent to which it explains rising premiums during our study period. We then explore heterogeneous effects along various dimensions. We first consider firm size as an important source of heterogeneity. Firm size matters partly because large firms can be less vulnerable to the market power effect, thanks to their ability to use self-insurance as a credible threat when negotiating with insurers. However, the option of self-insurance is less feasible for mid-sized firms due to the associated financial risks and large fixed costs. On the other hand, we suspect that the potential cost savings from the purchase power effect are likely distributed equally across firms because regardless of size, firms typically share provider networks under a given insurer. As a result, mid-sized firms would expect a larger increase in premiums. We also explore additional sources of heterogeneity. Those analyses will not only shed light on the relative magnitudes of the market power and purchase power effects, but also help us gain an understanding of whether the benefit of competition is distributed equally across employers/customers — an aspect of merger evaluation that has not yet been examined in the ESI literature.

Our empirical analyses take advantage of eight largest mergers that occurred during our sample period. We first offer an event study design to directly assess the impact of those mergers. Our results, largely based on a stacked difference-in-differences (DiD) approach, reveal that mergers lead to a 4-6% increase in premiums. Moreover, we find no evidence that the control and treated markets (which vary in their exposure to mergers) have exhibited differential pre-trends in premiums. Despite the fact that we were constrained in the coverage of markets in order to achieve a clean event study design, these results nonetheless help to reinforce our following identification strategy that uses mergers as exogenous shocks to study market outcomes.

We next turn to an instrumental variable (IV) approach, which enables us to exploit all the variations

in market competition induced by mergers and to establish a direct link between insurer market power and premiums. Our results reveal that a 10% increase in insurer HHI leads to a 0.7% increase in premiums. These findings suggest that, on average, the market power (premium-increasing) effect dominates the purchase power (cost-saving) effect. They also suggest that, during our study period, around 7% of the increase in premiums can be explained by increased insurer consolidation. These results are not sensitive to alternative specifications and sources of identification. Our further analyses reveal substantial heterogeneity in the impact of insurer concentration along the distribution of firm size. In particular, we find that insurer concentration leads to a premium increase of 1.2% for mid-sized firms but a decrease of 3.5% for large firms, when using 1,500 enrollees as the cutoff for firm size (averaged within a firm over time).[5] We then address concerns related to multi-location among large firms, as well as our construction of insurer HHI, before we explore other market outcomes, including plan characteristics, total enrollment, and the choice of funding type.

To shed light on the underlying mechanisms, we further examine heterogeneity based on firms' potential to use self-insurance as a credible threat, hospital market competitiveness, and a plan's merging status. We find that more credible access to self-insurance can protect employers from the premium-increasing effect of a merger. Moreover, the purchase power effect is most evident in less concentrated hospital markets, and firms subscribing to merging insurers experience larger premium increases than those who do not. These results help guide us to develop a bilateral bargaining model in explaining our empirical results. We show that under certain circumstances, the model yields predictions consistent with our empirical findings.

The existing evidence on insurance market competition and ESI premiums is fairly limited, due to challenges on data availability and identification [Gaynor et al. (2015)]. One important exception is Dafny et al. (2012), which used proprietary data on predominantly self-insured health plans offered by a sample of 813 super-large firms between 1998 and 2006. Specifically, they examined the 1999 merger of two national insurers (i.e., Aetna and Prudential) and found that an average increase in HHI increased premiums by about 13% during the specified timeframe. Another two related studies include Trish and Herring (2015) and Guardado et al. (2013). Our work contributes to this strand of literature in several

[5] Using different thresholds has delivered similar patterns of results.

ways. First, to the best of our knowledge, our study is the first covering the near-universe of employers that offer ESI. Second, we focus specifically on fully-insured plans, where insurer market power is expected to play a more critical role in determining premiums as compared to self-insured plans. Last, we assess recent trends in insurer competition and ESI premiums as new waves of consolidations have further intensified concentration in the U.S. insurance market over the past two decades.

Our work also contributes more broadly to the literature on vertical bargaining. A growing strand of literature examines whether downstream concentration leads to countervailing buyer power [Galbraith (1952); Chipty and Snyder (1999); Inderst and Shaffer (2008); Grennan (2013); Iozzi and Valletti (2014); Gaudin (2018); Barrette et al. (2020); Chorniy et al. (2020)]. Our study offers suggestive evidence of the countervailing power effect, especially in less concentrated hospital markets where insurers can more credibly threaten to substitute with alternative hospitals, a result consistent with earlier empirical findings, such as Town and Vistnes (2001), Sorensen (2003), and Ellison and Snyder (2010).

More importantly, although the majority of the empirical studies thus far have focused on input prices [e.g., Chipty and Snyder (1999); Grennan (2013); Barrette et al. (2020)], we study the impact of downstream concentration on consumers' final prices (i.e., insurance premiums).[6] Existing work has highlighted the importance of the nature of competition in the downstream market [Crawford and Yurukoglu (2012); Sheu and Taragin (2021)], the pass-through rate of input prices to retail prices [Gaudin (2018)], and the bargaining leverage of downstream firms [Dranove et al. (2019); Ho and Lee (2017)]. Of particular interest, Sheu and Taragin (2021) find that the countervailing effect could lead to lower final prices in the downstream market; our study lends empirical support to these findings. In addition, we find that there exists substantial heterogeneity in the effect of downstream concentration, dependent on the characteristics of final consumers (i.e., employers in our setting).

[6] It is interesting to note that insurers play an intermediary role. To some extent, insurer concentration in the downstream market has a similar effect to hospital mergers in the upstream market. This similarity allows us to draw insights from related studies on hospital mergers in a bargaining framework, such as Ho (2009), Town and Vistnes (2001), Capps et al. (2003), and Gowrisankaran et al. (2015).

2 Data

Our main data source is Form 5500, or the Annual Returns of Employee Benefit Plan, which is required annual reporting under Title I and Title IV of ERISA (Employee Retirement Income Security Act) and the Internal Revenue Code. Form 5500 collects information concerning the operation and funding of employee benefit plans. It consists of a main form and a number of schedules and attachments, which offer information such as the identity of the employer, the funding and benefit arrangements, and the corresponding premium costs, plan enrollment, and the name of the insurer.

Our data cleaning process consists of several steps. First, we narrow our sample to single-employer health benefit plans by removing about 4% of multi-employer plans. This step ensures a more straightforward relationship between firms and insurers, since it is not possible to track which employers bargain as a group in multi-employer plans. Second, we categorize plans based on their funding type (fully- or self-insured) following the Annual Report to Congress on Self-Insured Group Health Plans, prepared by the U.S. Department of Labor. This step is critical because these two types of plans are fundamentally different. In a fully-insured plan, the employer pays a pre-determined premium to an insurer, beyond which the employer bears no additional risks. In this case, an important component of premiums is the insurer's profit margin, which is tied closely to the insurer's market power. By contrast, self-insured employers retain the risk of paying medical claims themselves. They usually set aside a specific amount each month to pay for administrative fees, stop-loss insurance, and expected claims; this amount reflects the plan's expected costs. Ultimately, because we are interested in how premiums respond to insurer market power and our premium data for self-insured plans are quite sparse, our analysis focuses on fully-insured plans offered by fully-insured employers. Third, to back out the insurer ownership structure, we identify an insurer's parent company, using the NAIC Listing of Companies.[7] Lastly, to better measure market competition, following Dafny (2010), we restrict our sample to firms located in MSAs with at least 20 fully-insured employers. These MSAs represent 34% of markets, 91% of firms, and 89% of enrollees in the raw data. Eventually, we are left with a total of 38,356 distinct firms in 128 markets over our study period.

[7] Note that insurer identity in Form 5500 is reported at the EIN level. The parent company data are available for years 2005 to 2018. For the years before 2005, we used insurer group identifiers from 2005. We did not count Blue Cross Blue Shield as one insurer group because it claims to be an association of independently operated companies.

Our main data are supplemented with data from other sources. First, to obtain information on insurer mergers and acquisitions (M&A), we relied on the Health Care Acquisition Reports from the Irving Levin database, in addition to hand-collected M&A news for the years 2000 – 2017. We also merged data on market-level controls, such as the county unemployment rate from the Bureau of Labor Statistics and the county poverty rate and median household income from the Census Bureau. Finally, we acquired data on MSA-level hospital HHI from the American Hospital Association for the years 2000 – 2015.

The unit of our main analysis is a plan-year, where "plan" is defined as a unique combination of firm, insurer, and plan type (HMO, PPO, Indemnity, or Other). For each specific plan, a firm files a Schedule A to disclose its plan information. Table 1 displays summary statistics wherein all of the value variables are inflation-adjusted to year 2010. Premium per enrollee averages $\$5,303$, with a standard deviation of $\$2,891$. Total enrollment, which includes dependents and spouses for family plans, is highly right-skewed and has a mean of 756.[8] We construct a family ratio as the ratio of total enrollment over employee enrollment, in order to reflect the composition of single versus family plans within each firm. The two most popular plan types are HMO (38%) and PPO (23%).[9] On average, 42% of plans cover prescription drugs, 13% dental benefits, and 11% vision benefits.[10]

[TABLE 1 ABOUT HERE]

We also offer those summary statistics separately for mid-sized and large firms, using 1,500 enrollees as the cutoff for firm size (averaged within a firm over our study period). On average, premium per enrollee is slightly lower for a large firm. Although most of the plan characteristics are similar across these two samples, large firms are more likely to offer an HMO plan and have a lower probability of attaching drug coverage to a medical plan (which is consistent with large firms being more likely to contract a prescription drug plan separately).

We calculate HHI at the MSA-year level as a measure of insurer concentration on account of its ease

[8] We explain in Appendix A4 how we deal with the measurement error in total enrollees for a given plan in the data. Regarding whether our calculation of premiums is reliable, we also show in Figure A1 of the Appendix that our data are comparable to other data sources that have been used in the existing literature.

[9] HMO, PPO, Indemnity, and Other are not mutually exclusive because some firms aggregate their plans to file one Schedule A.

[10] Prescription drug coverage might seem low; this is largely due to the fact that many firms choose to cover prescription drugs through a separate plan, rather than combining the coverage with a medical plan (the latter of which is the focus of our study). For example, according to a Kaiser Family Foundation survey, about 30% of workers had carved-out drug benefits.

to construct and its central role in merger analysis [Nocke and Whinston (2020); Miller et al. (2021)]. More specifically, $HHI = \sum_i s_{it}^2 \times 10,000$, where s_{it} is the market share (by enrollment) for insurer i at a given MSA market. We are reassured that our constructed insurer HHI is comparable to the literature [e.g., Dafny et al. (2012); Trish and Herring (2015)].[11] To offer some preliminary evidence from the data, Figure 1 plots insurer HHI and premiums (inflation-adjusted and weighted by enrollment) at the MSA level for 2000 and 2018. For an easy comparison over time, we categorize markets into quintiles using cutoffs derived from the distribution in the baseline year (i.e., 2000), with a darker color suggesting a higher quintile. As is evident in this figure, both insurer HHI and premiums have witnessed a large increase during our study period. The question is whether this observed increase in premiums is driven by the concurrent increase in insurer market power.

[FIGURE 1 ABOUT HERE]

3 Empirical Strategy

In this section, we discuss our strategies in identifying a causal effect of insurer competition on premiums. We first present evidence that mergers among national insurers have generated important variations in market concentration and that these variations allow us to isolate the effect of increased insurer market power due to changes in competition. We then discuss an event study approach to assess the impact of mergers directly. Lastly, given that the majority of treated markets in our sample have experienced multiple mergers and are therefore excluded from our event studies, we turn to an instrumental variable (IV) strategy, which enables us to exploit all the variations induced by mergers and to establish a direct link between insurer market power and premiums.

[11] For example, the average insurer HHI in Dafny et al. (2012) was 3,000 in 2006, as compared to 3,163 in our data. The mean insurer HHI in Trish and Herring (2015) from 2006 to 2011 was 3,107, whereas ours is 3,207.

3.1 Mergers as the Source of Variation

We study the top eight mergers that occurred during our sample period.[12] These mergers had a significant market overlap between the acquirer and target and varied in their pre-merger market presence, which generates important variations across mergers and across markets within a merger. Furthermore, as these mergers were also negotiated nationally, they were arguably exogenous to local market conditions. In particular, during our study period, several major mergers, including UnitedHealth's acquisition of the PacificCare Health System and Aetna's acquisition of Coventry Health Care, were primarily driven by the acquirer's expansion into non-ESI related businesses, such as government-held Medicare and Medicaid.

To capture merger-induced variations in insurer concentration across markets, we construct delta HHI ($DHHI = 2s_i s_j$), with s_i and s_j referring to market shares for the acquirer and target one year prior to a merger. Figure 2 plots a map for cumulative DHHI. Note that a large number of MSAs were exposed to multiple mergers, so this figure reflects the cumulative DHHI induced by all eight mergers. Different colors represent quintiles of the DHHI distribution. Among a total of 128 MSA markets covered in this study, more than 30 (20) markets had a cumulative DHHI larger than 100 (200). A small set of markets (26) was never exposed to any of those major mergers during the sample period.

[FIGURE 2 ABOUT HERE]

To offer additional information regarding the mergers, Panel (a) of Figure 3 plots the histogram of DHHI among treated markets, representing a total of 225 MSA-year observations for 102 distinct MSAs affected by those mergers.[13] More than 130 of the MSA-years in our sample experienced a merger that led to a DHHI larger than 10; among those, 60 MSA-years had a DHHI larger than 50. Panel (b) of Figure 3 shows a scatter plot of DHHI and post-merger HHI (defined as the insurer HHI measured at the first year post-merger). Among markets that had a DHHI larger than 200, most had a post-merger HHI of less than 3,000. Among markets with a post-merger HHI larger than 4,000, most had a DHHI smaller than 100. Overall, those markets exhibit a significant amount of variation in their exposure to mergers,

[12] In Appendix A2, we list these mergers and offer more detailed discussions of each. We also acknowledge the caveat that mergers that are most likely to have led to anti-competitive effects are those most likely to garner the attention of antitrust enforcement and get blocked.

[13] In this figure, each observation is an MSA-year, so if a market was affected by multiple mergers across different years, that MSA would show up multiple times in the figure. Some markets were affected by multiple mergers in a single year; in this case, DHHI is calculated by combining all the simultaneous increases in HHI.

and apparently, mergers were more likely to go through when the resulting post-merger HHI was lower than 4,000.

[FIGURE 3 ABOUT HERE]

3.2 Event Study Design

We now assess the impact of mergers directly within a stacked difference-in-differences (DiD) framework following Cengiz et al. (2019) and Deshpande and Li (2019). Specifically, we group the eight mergers into three cohorts (2005/2006, 2008, and 2013), depending on the timing of their completion within a calendar year, as detailed in Appendix Table A1. We then construct cohort-specific datasets [denoted as c in Equation (1)] by including treated and control markets. Note that we combine the 2005 and 2006 mergers together, as this allows us to include markets that were exposed to both mergers in our analysis and therefore enhance statistical power.[14] We last stack the three cohort-specific datasets and use the following event-study specification to estimate the treatment effects of mergers within a 7-year event window (k=-3 to k=3):

$$ln(Premium_{pfmtc}) = \sum_{k=-3}^{3} \delta_k D^k_{mtc} + X_{pfmtc} + u_{fc} + \tau_{tc} + \xi_{pfmtc} \tag{1}$$

where $ln(Premium_{pfmtc})$ is the natural logarithm of premiums per enrollee for plan p at firm f in market m at time t for cohort c. We are mainly interested in the coefficients for dummies indicating each year of an event window for the treated group (D^k_{mtc}) (k=0 denotes the first year post-merger). We also control for cohort-specific firm (u_{fc}) and time fixed effects (τ_{tc}), as well as other controls (X_{pfmtc}), including lagged county-level economic conditions (unemployment rate, poverty rate, and median household income), lagged firm-level controls (firm size and family ratio), and contemporaneous plan-level controls (plan type and other provisions, such as dental, vision, and drug coverage). Standard errors are clustered at the firm level.

Several steps have been taken in creating the cohort-specific datasets. For a given cohort, we first specify a time window in order to select the treated and control markets. This window starts 6 years prior

[14] In the 2005/2006 cohort, we uniformly treat 2005 as the event year. This cohort has a total of 25 treated markets. Among those, 7 markets were exposed to mergers exclusively in 2005, 10 in 2006, and 8 in both 2005 and 2006.

to the event year and lasts for a total period of 10 years.[15] For example, for mergers that occurred in 2008, the window ranges from 2002 to 2011. We then consider a market to be treated if it is meaningfully affected (i.e., merger-induced DHHI larger than 10) only by the event of interest during this pre-specified window. A control market is defined as being unaffected by any mergers during the same window. Note that our definition of the treated and control markets requires that these markets were not contaminated by other major mergers during the corresponding window. For example, neither the treated nor control markets for the 2008 cohort were affected by the 2005/2006 mergers. These steps help to deal with the complication that a large number of markets were exposed to multiple mergers and allow us to only include "clean" treated and control markets in each cohort-specific dataset.[16]

We go through an additional screening step to eliminate the concern that some markets (both treated and control) were simultaneously contaminated by small mergers. Note that these small mergers are not specifically used in our identification, as they are likely endogenous to local market conditions. Although such mergers are considered small in terms of the number of affected geographic markets, some actually could result in a significant increase in local market concentration. We therefore drop those markets from our event study, wherever any of the small mergers generated a DHHI larger than the cutoff of 10 during the defined window.[17]

Eventually, our event study design results in a clean set of treated (38) and control (63) markets. For the 2005/2006 and 2008 cohorts, their control markets consist of those never-treated by any mergers or later-treated by the 2013 mergers. For the 2013 cohort, as 2005/2006 is outside its 10-year window, its treated markets include both those first-treated in 2013 and 2 additional markets that were initially treated in 2005/2006. Its control markets include those never-treated as well as 13 markets that were initially treated in 2005/2006. Dropping those 15 earlier-treated markets from the 2013 cohort does not change our results.

[15] Note that we take a conservative approach by using a 10-year window in defining the treated and control groups, which ensures that our dynamic estimates are not contaminated by events excluded from the specification [Sun and Abraham (2021)]. We also note that our event study results, if based on a 7-year window in defining the treated and control groups (which starts 3 years prior to the merger year), are largely consistent.

[16] For example, 59 distinct markets were meaningfully affected by 2005/2006 mergers. However, because some of those markets were also exposed to mergers in 2008, only 25 markets were exclusively affected by the 2005/2006 mergers during the window between 1999 and 2008 and therefore were included in our event study.

[17] In our stacked DiD study, this step reduces the total number of treated markets by 6 and the number of control markets by 10. Results without this screening step are largely consistent with our main findings.

3.3 An IV Approach

Note that our event-study design is constrained to a small set of markets that were not exposed to multiple mergers for a specific window. To exploit all merger-induced variations, we next take an IV approach. We specifically examine how changes in insurer HHI induced by mergers affect premiums. Focusing on insurer HHI allows us to capture differential changes in insurer competition across mergers, as well as across markets within a merger. It also has the advantage of revealing an intuitive relationship between market concentration and prices.[18] Furthermore, our findings can directly inform antitrust enforcement, where HHI analysis is considered to be economically well-founded and play a central role in merger assessment [Nocke and Whinston (2020); Miller et al. (2021)].

We proceed by examining the aggregate effect of insurer concentration following Equation (2):

$$ln(Premium_{pfmt}) = \beta_1 ln(HHI_{mt-1}) + X_{pfmt} + u_f + \tau_t + \varepsilon_{pfmt} \tag{2}$$

where $ln(HHI_{mt-1})$ is the natural logarithm of the lagged local market concentration. We use a lagged HHI because premiums are usually negotiated one year in advance. X_{pfmt} are the same set of controls as we specified in Equation (1). We also control for firm (u_f) and year (τ_t) fixed effects. The inclusion of firm fixed effects alleviates concerns regarding time-invariant unobserved differences in the risk profiles of enrollees and plan characteristics that might be correlated with local market conditions. Standard errors are clustered at the market-year level.

Equation (2) presents a challenge for identification because market structure and premiums are simultaneously determined. For example, a market experiencing a positive demand shock might experience an increase in premiums and attract more insurers to enter. Ignoring such endogeneity would lead us to underestimate the effect of insurer market power on premiums. A related challenge arises when measuring market concentration using insurer HHI because observed variation in HHI could be driven by fundamentally different economic primitives. As a result, multiple paths can explain a given correlation between concentration and market outcomes [Berry et al. (2019)].

[18] We acknowledge that there is no corresponding theoretical model that suggests a direct effect of HHI on prices. However, to the extent that some of our controls can be used to measure cost differences (such as firm fixed effects, plan type, and benefit coverage), one can interpret our model as establishing a causal relationship between markup and HHI in the Cournot setting [Cowling and Waterson (1976)].

To address these concerns, we adopt the strategy used in Dafny et al. (2012). Specifically, we utilize mergers among insurers that had significant market overlap pre-merger as the underlying drivers for changes in HHI. To proceed, we calculate a simulated HHI ($SimHHI_{mt}$) by utilizing exogenous shocks induced by our top eight insurer consolidations. More specifically, for an insurer i in market m, we start with its first-year market share s_{imt}. For year $t+1$, we check to determine whether there were mergers affecting market m. If yes, for the merging insurers, we calculate s_{imt+1} by summing up s_{imt} for both the acquirer and the target. For non-merging insurers or for markets not affected by a merger, $s_{imt+1} = s_{imt}$. We then calculate the simulated HHI ($SimHHI_{mt+1}$) accordingly. This process is repeated until we reach the end of our study period. Note that the simulated HHI will vary only due to mergers and acquisitions. For a market that has never been exposed to any mergers during our study period, the simulated HHI will stay constant at its baseline year (i.e., 2000) level.

4 Main Findings

In this section, we first present evidence from our event study approach. We then delve into the results from our IV approach, first presenting the aggregate effect of insurer concentration and then exploring the heterogeneity and robustness of our main findings. We last examine other market outcomes.

4.1 Event Study Results

Panel (a) of Figure 4 shows the stacked DiD event study results. We find that mergers lead to about a 4-6% increase in premiums, with an immediate 5% increase in the first year post-merger. There is no evidence that the treated and control markets have exhibited differential pre-trends in premiums. Panel (b) presents the event study results following Callaway and Sant'Anna (2021).[19] These results exhibit no discernible pre-trends and the estimated price effect is also largely consistent with the stacked DiD

[19] Note that in Callaway and Sant'Anna (2021), a market can only be treated once. We therefore define the year of treatment as the first year in the data that a market was meaningfully affected by one of the eight major mergers (i.e., the merger-induced DHHI is greater than the cutoff of 10). We also make sure that the treated markets were not contaminated by smaller mergers before their respective year of treatment, and removed control markets that were affected by those mergers too. Due to the closeness in timing for the 2005/2006 and 2008 mergers, 34 markets that were treated by the 2005/2006 mergers were later treated by the 2008 mergers. Results stay consistent if we drop those observations after they were later treated by the 2008 mergers.

results.

[FIGURE 4 ABOUT HERE]

We also present the robustness of our stacked DiD results. First, we show that our results remain similar if dropping all the time-varying covariates [Panel (a) of Figure 5]. Second, our results are robust if weighted by the number of firms in an MSA averaged over our sample period [Panel (b)].[20] Third, though we think the 2013 merger is reasonably distant from the 2005/2006 ones, we also show that our results are consistent if dropping the 15 earlier-treated markets (by 2005/2006 mergers) from the 2013 cohort [Panel (c)]. Lastly, as some markets are included in multiple cohort-specific datasets, we conduct an analysis by randomly assigning these markets to only one cohort-specific dataset; these results [Panel (d)] are also largely consistent with our main findings. In Appendix Figure A2 and Figure A3, we present other robustness results wherein we experiment with alternative cutoffs in defining treated and control markets, as well as event study results by firm size wherein we find that the increase in premiums is largely driven by mid-sized firms in response to increased insurer concentration.

[FIGURE 5 ABOUT HERE]

Our event study results reveal that mergers lead to higher premiums. Also, they help to validate our identification strategy of using mergers as exogenous shocks to study market outcomes. Note that our IV of simulated HHI can be considered a Bartik instrument, in which the DHHI measures differential exposures to a common shock (here, a merger). To further strengthen the credibility of our IV approach, we follow a strategy proposed by Goldsmith-Pinkham et al. (2020) and offer an equivalent test of parallel trends across markets with different DHHIs. Specifically, we run the parallel trend test by interacting the DHHI with event time fixed effects following our event study specification. We report the coefficients for the interaction terms in Figure 6. These results suggest no evidence that markets with different DHHIs exhibit differential pre-trends in premiums, a finding that offers additional support to our IV strategy. We also conduct an additional test of the parallel trends by estimating market-specific pre-trend coefficients using the event study framework. These results are not included due to space limitations, but again we find no clear relationship between those pre-trends and DHHI.

[20] Results are also similar if weighted by the number of employee enrollment for an MSA averaged over our sample period.

[FIGURE 6 ABOUT HERE]

4.2 Aggregate Effect of Insurer Concentration

Our IV strategy based on Equation (2) takes advantage of the largest eight mergers among insurers that occurred during our study period. We first offer empirical evidence on how these mergers have generated important variations in insurer concentration due to differences in their pre-merger market presence. Figure 7 depicts a binscatter plot of the local insurance market HHI and the simulated HHI for each market-year observation in the data. As evident in this figure, our simulated HHI captures a large amount of the variation in insurer HHI that is observed in the data. Indeed, the solid line represents the best linear fit with an estimated slope of about 0.57. We also present two sets of first-stage regression results in Table 2. Column (1) is analyzed at the MSA-year level (without the inclusion of X_{pfmt}), and Column (2) is analyzed at the plan-year level, which corresponds to the first stage for Equation (2). The estimated coefficient of $ln(SimHHI)$ is positive and statistically significant for both specifications, with the large t-statistics suggesting a strong first stage. Interestingly, the point estimate is smaller at the plan level, suggesting a weaker association between the observed HHI and the simulated HHI among larger MSAs. This result is likely if larger MSAs tend to experience more variability in local market structure.

[FIGURE 7 ABOUT HERE]

[TABLE 2 ABOUT HERE]

We now proceed to discuss the aggregate effect of insurer concentration. Column (1) of Table 3 presents the OLS results. It suggests that a 10% increase in insurer HHI is associated with a 0.3% decline in premiums.[21] However, as we discussed earlier, the OLS results are likely to be contaminated by confounding factors. Columns (2) and (3) of Table 3 present the reduced-form and IV results, respectively. Different from the OLS results, these results suggest that a higher level of insurer concentration induced by insurer mergers actually leads to higher premiums. The comparison of the IV and OLS estimates suggests a downward bias due to the endogeneity of insurer concentration.

[21] Our OLS results might seem surprising given that a positive correlation between concentration and premiums has been documented in various settings. We note, however, that we do see a positive and statistically significant coefficient for HHI when conducting a pooled OLS regression without controlling for firm and time fixed effects.

[TABLE 3 ABOUT HERE]

More specifically, we find that premiums increase by 0.7% for a 10% increase in insurer HHI. In our data, inflation-adjusted premiums have increased from $\$4,116$ in 2001 to $\$5,781$ in 2018, representing a 0.3 increase in the logarithm of premiums. During the same period, the insurer HHI increased from $2,529$ to $3,412$ on average. Our point estimates suggest that the rising concentration in the local insurance market explains about 7% of the premium increases between 2001 and 2018.[22]

We next explore several alternative specifications and sources of identification in Figure 8. Recall that our preferred specification includes time-varying covariates at the plan, firm, and market level. We first show that our results are not sensitive to the choice of those covariates. We also consider two finer sources for identification, in which we focus on 1) four mergers that were driven by expansions in non-ESI markets and therefore arguably more exogenous to the ESI market, and 2) the largest four mergers in terms of the number of affected markets. Lastly, we study whether our results remain robust to the inclusion of hospital market HHI. We estimate the effects separately across all combinations of specifications and identifications. Note that firm and year fixed effects are always included. These results offer some reassurance that our main findings are indeed robust to a range of specification-identification combinations.

[FIGURE 8 ABOUT HERE]

4.3 Heterogeneous Effect of Insurer Concentration

The above analyses suggest that the market power effect dominates the purchase/countervailing power effect, leading to an overall increase in premiums. Given our focus on the bargaining between insurers and employers, we explicitly explore potential heterogeneity by differentials in firm size. Regarding the purchase power effect, we suspect that the benefit of lower provider prices is distributed equally across firms because firms, regardless of their size, usually share a similar provider network under a given insurer. That said, we also note that large firms likely possess greater bargaining power, which may facilitate a higher pass-through rate of any lower costs and enable them to capture more cost savings.

[22]Note that the logarithm for HHI increased by 0.3 (ln(3,412)-ln(2,529)). Multiplied by the point estimate of 0.07, the model predicts an increase of 0.02 in the logarithm of premiums, which represents 7% of the total increase of 0.3.

Regarding the market power effect, we suspect that a large firm could also be less amenable to insurer consolidations. One reason for this conjecture is that large firms could more easily turn to self-insurance, an option that is less feasible for mid-sized firms, due to the exposure to risks and the large, associated fixed costs. Their size also could entail a larger reduction in insurers' total profits should the negotiation break down, a fact that further enhances the bargaining leverage for large firms. What is more, their size possibly gives them greater opportunities to assume the cost of acquiring information and gathering resources (e.g., hiring a team of skilled negotiators), which offers them more bargaining power and puts them in a better negotiating position [Lewis and Pflum (2017)]. Their bargaining power could also benefit from their lower level of risk aversion as compared to mid-sized firms. All of these differences could put large firms at a relative advantage in the case of insurer consolidations as compared to mid-sized firms.

To test the existence of such heterogeneity, we first offer an analysis that extends the baseline model by introducing an interaction term of insurer concentration and firm size, measured as a firm's total enrollment averaged over our sample period. The estimated coefficients are reported in Column (1) of Table 4, which suggests that the effect of insurer concentration is indeed decreasing in firm size. To explain this relationship better, we plot the estimated effect in the upper graph of Figure 9 where the X-axis measures firm size. This figure suggests that the estimated effect of insurer concentration could be as large as 4% given a 10% increase in insurer HHI. However, the magnitude of the effect decreases and, after a certain threshold of firm size, becomes negative. In addition, we show the distribution of enrollment by firm size in the lower graph of Figure 9. Firms with a size between 200 and 400 had the largest fraction of enrollment. Overall, large firms (with an average total enrollment larger than 1,500) covered around 35% of the enrollment.

[TABLE 4 ABOUT HERE]

[FIGURE 9 ABOUT HERE]

To offer a more straightforward demonstration of such heterogeneity, we now divide our sample into two parts by firm size and study the effect of insurer concentration separately for each sample. The results are reported in Column (2) of Table 4 for firms with a respective total enrollment below the threshold of $1,500$ and in Column (3) for firms above this threshold. Consistent with the results in Column (1),

we find that mid-sized firms expect a 1.2% increase in premiums for a 10% increase in insurer HHI. In contrast, large firms actually experience a 3.5% decrease in premiums given the same amount of increase in HHI. Given that our choice of threshold can be seen as somewhat arbitrary, we also report the results in Columns (4) and (5) for a cutoff of $1,000$ and in Columns (6) and (7) for a cutoff of $2,000$.

We discuss two additional pieces of evidence to support our results regarding the heterogeneous effect. First, we explore such heterogeneity separately for mid-sized and large firms. We adopt the same model that includes the interaction term between insurer concentration and firm size, and we report the results in Column (8) of Table 4 for mid-sized firms and in Column (9) for large firms. These results suggest a coherent story that firm size plays an important role in explaining the effect of insurer concentration, even within mid-sized firms. Secondly, we focus on the sample of large firms and examine whether our results might be driven by some outliers. To proceed with this inquiry, we create a subsample by randomly dropping 10% of the large firms and then run our analysis using this subsample. We repeat this process 500 times and plot a histogram of the estimated main coefficient of interest in Figure 10. These results suggest that our main results hold for varying samples; that is, outliers are unlikely driving our findings.

[FIGURE 10 ABOUT HERE]

Overall, we find consistent results that the effect of insurer concentration is decreasing in firm size. These results suggest that a larger fraction of premium growth could be explained by insurer concentration for firms at the lower end of the size distribution; by contrast, however, large firms actually see a decrease in premiums when the insurance market becomes more concentrated. This decrease in premiums offers suggestive evidence of the countervailing power, whereby insurers take advantage of their enhanced bargaining leverage when negotiating with healthcare providers. That said, we cannot rule out other efficiencies associated with a merger. For example, the consolidations of medical claims across merged insurers could allow them to price risk more accurately — an efficiency that may be more readily realized for large employers, who typically generate a higher volume of claims.

4.4 Robustness Analyses

To assess the robustness of our main results, we first specifically address several concerns related to the multi-location of large firms. We then show that our main results remain robust when we construct insurer HHI by considering both fully- and self-insured plans.

Multi-Location of Large Firms Large firms might operate across multiple geographic markets, and plans offered across their varied locations could differ. Due to data limitations, we do not observe individual plans for each operating site. Instead, Form 5500 reports an aggregate plan, which can be considered an aggregation of all plans offered by a given insurer within a given firm. Our data also only allow us to observe a firm's headquarters location.

Consider a large firm that operates in its headquarters market, as well as in other geographic markets. The premium that we observe in the data captures the aggregate premium weighted by enrollment across all the operating markets. In a scenario where the headquarters market represents the majority of the employee population, our main analysis would still be valid. In other scenarios, it could be challenging to define the insurance market boundary for a multi-site firm. Our main analysis assigns each plan to its headquarters location, and we use MSAs to define a market following Dafny et al. (2012). One might argue that the headquarters MSA is not necessarily the market boundary for a large employer. However, it is worthwhile to note that this would go against our finding of any significant effect of local market concentration. Moreover, we note that our previous analyses, particularly the results in Table 4, have demonstrated substantial heterogeneity in the effect of insurer concentration along the dimension of firm size, even within mid-sized firms. This heterogeneity suggests that our findings are not driven exclusively by large and multi-location firms.

In addition, we offer several robustness tests. We first examine whether our results hold if we remove the sample of multi-site publicly listed firms. To do this, we acquired information from Compustat and Bureau Van Dijk,[23] which allows us to determine whether a publicly listed firm has ever operated across multiple MSAs. Then we removed those multi-site firms from our analysis. These results are reported

[23]Compustat has information on publicly listed firms, including their names, EIN, and stock tickers. Bureau Van Dijk covers these firms' subsidiary locations and stock shares held by the parent firm. We kept all subsidiaries whose parent stock holdings were no less than 50%. Then, we merged Compustat with Bureau Van Dijk using cusip ID, ticker symbol, and firm names. Finally, we merged this data with our Form 5500 using EIN and year. Of all the Compustat firms, 34% were merged to Form 5500. In our analysis sample, around 86% of the publicly listed firms were located in multiple MSAs.

in Column (1) of Table 5. A much larger fraction of observations was removed from the sample of large firms, which is consistent with the fact that they are more likely to operate in multiple locations. These results lend support to our main results, although we acknowledge that they do not capture multi-locations for privately owned firms.[24]

[TABLE 5 ABOUT HERE]

Multi-location also leads to a complication related to the calculation of insurer concentration for a local market. In our main analysis, we assign a firm's total enrollment to its headquarters location. However, there remains a concern of misallocating enrollment and thereby miscalculating insurer concentration. We therefore offer an alternative measure of insurer HHI that excludes firms with an enrollment of more than $1,500$.[25] These results, reported in Column (2) of Table 5, remain consistent with our main findings.

We also report results whereby we replicate our analysis using "state" as the market boundary. We consider MSAs to be the preferred measure of a market throughout our study, but defining markets by state might be useful, possibly for large firms, if we are concerned that those firms might operate across multiple MSAs. These results are reported in Column (3) of Table 5. The estimated effects of insurer concentration remain positive for mid-sized firms and turn negative for large firms, although both effects become smaller in absolute magnitude.

HHI That Incorporates Self-Insured Plans We study premiums in the fully-insured ESI market, and we construct HHI based on fully-insured plans only. Although this construction largely reflects data availability, we present two pieces of evidence in our data to show that it is reasonable to assume that the market of fully-insured plans is distinct from that of self-insured plans, for the majority of the firms. First, we find in our data that most firms show a high level of persistence in their choice of funding type (i.e., self- or full-insurance). For example, about 74% of firms stick to one funding type between 2001 and 2018. Second, although there is evidence that some large firms change their funding type in response to changes in insurer concentration (detailed results are presented in Appendix A5), the magnitude of the

[24] We run into a power issue if we only focus on single-location publicly listed firms. For example, the number of observations is smaller than 300 for large firms.

[25] The choice of $1,500$ seems arbitrary, but the results remain robust with alternative cutoffs, such as $1,000$ and $2,000$.

effect is relatively small. We therefore believe, at least for the majority of the firms, that the fully-insured market can be considered a separate market.

However, there remains a concern that insurer concentration in self-insured plans could also play a role, especially for large firms that are more likely to consider self-insurance as a relevant option. To address this concern, we test whether our results hold if we measure insurer HHI for both fully- and self-insured plans. Though not a perfect way of measuring this, we utilize Schedule A filings among self-insured employers and construct an MSA-level HHI based on total market shares, which include both fully- and self-insured plans.[26] We observe a correlation of 0.82 between the fully-insured HHI (our main measure of HHI) and the overall HHI that also incorporates self-insured plans. The results, reported in Column (4) of Table 5, are largely consistent with our main findings.

4.5 Other Outcomes

We now explore how insurer competition affects other market outcomes that are welfare-relevant and can help to strengthen our causal interpretation of our main findings. First, we study whether firms adjust their plan characteristics in response to changes in the competitiveness of the insurance market. Second, we examine the quantity effect by examining changes in total enrollment for a given firm. Lastly, we study how insurer competition affects a firm's incentive to switch between offering fully- and self-insured plans. We highlight some main findings here but leave other discussions in Appendix A5 where we present those results.

Plan Characteristics Firms might possibly change the level of generosity of their offered plans when facing the pressure of higher premiums. Our data capture several key features of plan designs, including plan types (HMO or PPO) and the offering of dental, vision, or drug coverage. Overall, there is some evidence that plan characteristics respond to changes in insurer concentration. More interestingly, we find that premium increase has incentivized mid-sized employers to substitute to less generous HMO plans, possibly to offset increased prices. On the contrary, premium reduction has slackened the constraint for

[26] The instruction of Form 5500 suggests that employers should file Schedule A if they have insurance contracts. Although self-insured employers are exempt from filing Schedule A (because self-insurance is not considered as having an insurance contract), a large fraction of them (i.e., 62% in terms of self-insured employers, or 70% in terms of self-insured population) still do this filing. Their choice to file, in these cases, largely reflects that 1) self-insured employers use TPAs (third-party administrators, and therefore they still sign contracts with insurers through TPAs, and/or 2) most self-insured employers purchase stop-loss contracts (so their filing of Schedule A reflects the stop-loss contracts).

large employers to offer more generous PPO plans.

Note that our main results regarding premiums remain consistent, whether or not we explicitly include those plan characteristics as additional controls in our main specification. We also acknowledge that those observable measures of plan characteristics are not comprehensive. However, to the extent that unobserved plan characteristics might respond in a fashion similar to the observed ones, our study of how these latter characteristics respond to insurer concentration offers some reassurance that our main results are not likely attributable to their unobserved counterparts.

Total Enrollment We next examine how a firm's total enrollment responds to increased insurer concentration. We conduct this analysis at the firm-year level to better capture changes in total enrollment, rather than employees switching across plans within a firm. Our results suggest some decrease in total enrollment among mid-sized firms, yet some increase among large firms. The estimated effect is nevertheless small and lacks statistical precision. These quantity results are consistent with the heterogeneous premium effects identified in our main findings. Our results also support findings from the previous literature that the local price elasticity of take-up, conditional on being offered ESI plans, is very small [Cutler (2003); Gruber and Washington (2005); Royalty and Hagens (2005)].

Funding Type Although our main sample consists of fully-insured employers only, firms generally have the freedom to switch between offering a fully- and self-insured plan over time. The literature has documented various factors that come into play in firms' decisions over funding type, including firm size [Park (2000); Eibner et al. (2011)], firms' risk aversion preferences [Jensen et al. (1995)], differences in investment opportunities and external financing costs [Dalton and Holland (2019)], and state regulatory and tax environments [Jensen et al. (1995); Robinson (2022)].

Of particular interest to us is whether changes in insurer concentration offer incentives for firms to switch their funding type. We now expand our sample to the universe of employers in the data. Overall, our results suggest that the funding type of the majority of firms is not responsive to changes in insurer concentration. Although we do find some evidence that large firms tend to increase their likelihood of offering fully-insured plans, the estimated effect is small. These results also suggest that changes in sample composition are not likely to pose an appreciable threat to our identification in our main analysis. In this regard, we also note that when we conduct our main analyses focusing on firms that have been

consistently fully-insured over our sample period, all the results hold.

5 Underlying Mechanisms

We find, on aggregate, that premiums increase due to lessened insurer competition and that this effect is diminishing in firm size. To shed light on the underlying mechanisms, we further examine heterogeneity based on firms' ability to use self-insurance as a credible threat, hospital market competitiveness, and whether a plan's insurance carrier has been directly involved in a merger.

Self-Insurance as a Credible Threat Employers might differ in their ability to wield self-insurance as a credible threat in their bargaining with insurers on fully-insured plans. Generally speaking, mid-sized firms might lack the resources to manage a self-insured plan in the first place. They might also consider self-insurance less appealing due to the difficulty of assuming financial risks and dealing with the unstable cash flow caused by volatility in claims. As a result, the option of self-insurance might be a less credible threat for mid-sized firms than for their larger counterparts, and this prospect might help explain why the results show that mid-sized firms are hurt more in the wake of lessened insurer competition. Indeed, the probability of choosing self-insurance is much lower among mid-sized firms in our data.

To test this hypothesis, we construct a predicted self-insured ratio to assess a firm's credibility in using self-insurance as a threat in their negotiations with insurers. More specifically, for all firms regardless of their funding type, we first calculate a self-insured ratio, which is defined as the number of years a firm has used self-insurance divided by the total number of years that we have observed for that firm. To address the concern that the calculated self-insured ratio could be an artifact of differences in insurer market structure and premiums, we construct a predicted self-insured ratio based on a linear regression model by controlling for variables that are arguably exogenous, such as firm size and industry.[27] Note that because large firms generally have a higher self-insured ratio, we regress the self-insured ratio separately for mid-sized and large firms.

[27] This regression is run at the firm level, as the outcome is the calculated self-insured ratio, which is fixed within a firm. The control variables include firm size, industry fixed effects (i.e., 9 categories based on SIC industry code), MSA fixed effects, and other market controls (i.e., unemployment rate, poverty rate, and median income, all averaged over time). Detailed results are reported in Appendix Table A4.

We examine how premiums respond to insurer market concentration, based on whether a firm has credible access to self-insurance. These results are reported in Panel A of Table 6. We first discuss mid-sized firms, which we divide into two groups using the 70th percentile of the distribution of their predicted self-insured ratio as the cutoff. We find that insurer concentration has a much larger price-increasing effect for mid-sized firms with a predicted self-insured ratio lower than the 70th percentile, as suggested in Column (2) of Table 6. In contrast, the effect of insurer concentration is much smaller for those with a predicted self-insured ratio above the 70th percentile [Column (5)].

[TABLE 6 ABOUT HERE]

Note that compared to the sample of mid-sized firms, the majority of large firms have arguably better access to self-insurance and can more credibly claim that they will use it. Thus, we now turn to large firms and study whether their ability to opt for self-insurance makes a difference. Similar to the analyses among mid-sized firms, we are interested in exploring whether premiums would respond differentially based on the relative ease of access that large firms have to self-insurance. We classify large firms with more credible access to self-insurance as those observing a predicted self-insured ratio greater than the 30th percentile of the distribution for those firms, and we find the negative effect of insurer concentration to be large and statistically significant [Column (6)]. For large firms below this cutoff, however, the estimated effects are much smaller in absolute value and lose statistical significance [Column (3)].

Although we divide our sample based on firms' relative access to self-insurance, it is worth noting that the distribution of firm size remains similar among these sub-samples [e.g., Column (2) v.s. Column (5)]. This suggests that our results are not likely driven by differences in firm size. When we pool mid-sized and large firms together, we find that the aggregate price effect remains positive and large only for firms with less credible access to self-insurance [Column (1)]. In addition, when we extend our analyses to different cutoffs in dividing the sample, we find largely consistent patterns of results.

In sum, our results imply that a firm's ability to opt for self-insurance could be an important channel through which a higher level of insurer concentration can have a differential effect on premiums. Being able to threaten more credibly that it will use self-insurance helps to shield a firm from increased premiums. However, the majority of mid-sized firms lack this credibility; accordingly, premiums increase among them overall. On the contrary, most large firms seem to benefit from more credible access to

self-insurance, and this advantage may help to explain the decrease in premiums that they experience when the insurer market becomes more concentrated.

Hospital Market Competitiveness Hospital prices are an important determinant of premiums. Ho and Lee (2017) offer a rigorous study of the joint determination of premiums and hospital prices. They explicitly allow insurers and hospitals to bargain over hospital prices in a similar fashion as employers and insurers bargain over premiums, and they find that removing an insurer from consumers' choice set will produce two offsetting effects on hospital prices: a positive effect due to hospitals extracting rents from premium increases, and a negative effect due to hospitals losing bargaining leverage. Therefore, the question of how hospital prices and thereafter premiums would be affected by lessened market competition remains ambiguous. Ideally, we would like to empirically examine hospital prices in the same fashion as Cooper et al. (2019), but we are very much constrained in our access to reliable data on hospital rates. In light of this limitation, we utilize a subsample of "experience-rated" health plans from Form 5500, whose premiums are considered more closely tied to medical claims;[28] we have found (see Table A5) that both premiums and claims declined in response to increased insurer concentration among these plans. We believe that these results offer evidence suggestive of lowered provider prices and thus of countervailing power in the wake of increased insurer concentration.

As an alternative test, we also explore whether downstream premiums respond to insurer concentration differentially by the heterogeneity in hospital market competitiveness. One would expect, in a concentrated hospital market, that the positive effect (due to hospitals extracting rents from premium increases) would be the largest and the negative effect (due to hospitals losing bargaining leverage) would be the smallest, thereby leading to higher provider prices and therefore premiums. To proceed with this test, we rank markets using their baseline year (i.e., 2000) hospital HHI and divide the sample in order to reach a relatively balanced sample across these two sets of markets. The results are reported in Panel B of Table 6. First, we find that premiums experience a bigger increase in more concentrated hospital markets [Column (4) v.s. Column (1)]. Second, we find, regardless of the competitiveness of the hospital

[28] Note that all large-group plans are experience-rated. Here, "experience-rated" plans refer to plans that reported premiums, premium reserves, claims, and claim reserves in Schedule A under the category of "experience-rated" contracts (Line 9a-9e). Compared to other fully-insured plans, "experience-rated" plans are eligible for a return of premiums at the end of a policy year if the plans' claims experiences are less than expected. Given this, to some extent, they are more like self-insured plans.

market, that the estimated price effects for mid-sized firms remain positive and largely similar to our main results. However, the price-decreasing effect for large firms is only evident in less concentrated hospital markets [Column (3)]. In contrast, in more concentrated hospital markets, we find the estimated effect for large firms to be large and positive, although not statistically significant.

These results offer additional evidence in support of the countervailing power effect of increased insurer concentration. In particular, in less concentrated hospital markets, increased insurer concentration is more likely to allow insurers to reach lower hospital prices. Some of the cost savings are then passed down to large firms in the form of lowered premiums. These results are consistent with earlier findings from Town and Vistnes (2001), Sorensen (2003), and Ellison and Snyder (2010).

Merging Status We now explore how premiums respond differentially by a plan's merging status. We take an ad hoc approach in defining merging status, following Bhattacharya et al. (2023). Specifically, a plan is considered "merging" if its insurer has ever been involved in one of the eight mergers of our focus. Merging status, therefore, is plan- and time-specific, based on our definition. We then proceed by running our regressions separately by a plan's merging status and examine how merger-induced changes in local market structure affect these plans differently.

We present these results in Panel C of Table 6. We find that merging plans have witnessed a larger increase in premiums as compared to non-merging plans. Such a difference in the magnitude of effects is largely driven by mid-sized firms. Both merging and non-merging plans respond in a similar fashion to increased insurer concentration, although the magnitude of this effect is more pronounced among merging plans. The fact that we find consistent patterns of results among merging and non-merging plans offers suggestive evidence that our main results are not likely driven by any direct effect of mergers on product/plan attributes. These results also suggest that input costs (hospital prices) are lowered for non-merging insurers as well, possibly due to these insurers exercising their enhanced countervailing power; however, we cannot rule out competitive spillover as an alternative explanation.

Summary The above analyses suggest important heterogeneity in the effect of insurer concentration on premiums. We offer a summary of these findings and discuss how they help guide us to develop a model presented in Appendix A7. First, we highlight the importance of credible access to self-insurance in explaining how premiums respond to insurer concentration; specifically, insurer consolidations would

hurt more among firms without credible access to self-insurance, thus offering a channel for differential price effects based on firm size. Second, we find that increased insurer concentration leads to lower premiums for large firms in less concentrated hospital markets, which suggests that decreased hospital prices (due to the purchase power effect) are another important channel through which insurer concentration leads to lower premiums. Lastly, we find consistent patterns of results among merging and non-merging plans, which allows us to abstract away from modeling any direct effect of mergers (e.g., on plan attributes) and to focus on the channel of enhanced market power among insurers.

In Appendix A7, we introduce a model of bilateral bargaining between firms and insurers, whereby we explicitly allow insurer competition to affect premiums by limiting the outside option and changing hospital prices. We show that, under certain circumstances, this model yields predictions that are consistent with our main empirical findings.

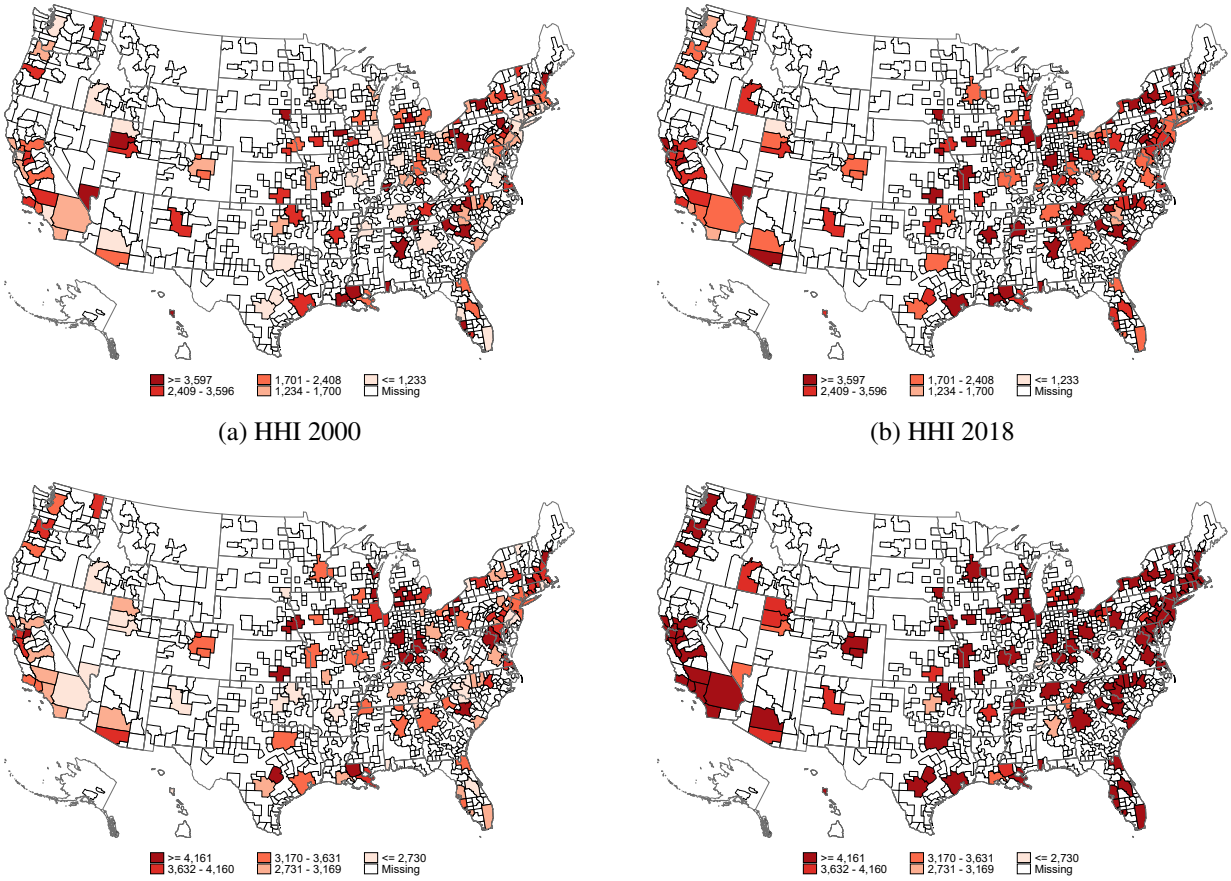
6 Conclusion

In this article, we investigate the causal relationship between insurer competition and market outcomes by analyzing a dataset that covers the near-universe of ESI plan offerings. Our identification takes advantage of a list of large mergers among national insurers. We offer two sets of results. First, we find that mergers lead to higher premiums and that a 10% increase in insurer concentration causes premiums to rise by 0.7% for fully-insured plans, suggesting that around 7% of the premium increase during our study period can be explained by increased insurer concentration. Second and more interestingly, we find important heterogeneity in how premiums respond to insurer concentration. In particular, the effect of concentration on premiums decreases with firm size and, for a subset of large firms, premiums actually decrease in the wake of increased insurer concentration. These results suggest that the purchase/countervailing power effect dominates the market power effect for those large firms. Although these results further suggest that insurer consolidation could lead to lowered hospital prices, the consequent implication for consumer welfare is largely ambiguous, due to concerns that lowered hospital reimbursement rates might negatively impact the quality of patient care and/or access to care because of supply-side responses. Our additional analyses have revealed other heterogeneity as well. Specifically, we find that firms lacking credible access to self-insurance, located in more concentrated hospital

markets, or subscribing to merging insurers are most vulnerable to insurer consolidations.

Our results offer important policy implications. First, we find that increased insurer concentration puts firms at a disadvantage when negotiating premiums with insurers. In particular, we find that mergers raise prices in proportion to the merger-induced increase in HHI. Our results therefore offer empirical support to the most recent 2023 Merger Guidelines, in which market concentration and the change in concentration due to a merger are often useful indicators of how a merger could potentially lessen competition.

Second, regarding antitrust enforcement and merger assessment, our results suggest that insurer consolidations could produce heterogeneous effects among final consumers. However, antitrust authorities have typically ignored the opportunity to explore heterogeneity. One possible reason for this lack of attention is the common practice of defining markets narrowly, based on a specific combination of product and customer characteristics for antitrust purposes. We therefore contribute by highlighting the importance of exploring heterogeneous effects. Related, antitrust authorities usually seek voluntary compliance before actually bringing a case to court. Although legal standards typically would not credit merger efficiencies realized in one market in order to offset consumer harms found in another, antitrust authorities can decide to weigh these impacts across markets when conducting their internal assessment on whether to litigate a merger. Furthermore, understanding those heterogeneous effects would help antitrust authorities to target the most vulnerable populations when proposing remedies. In this way, the remedies could be most properly designed to help minimize the anti-competitive aspects of a proposed merger.



(a) HHI 2000

(b) HHI 2018

(c) Premiums 2000

(d) Premiums 2018

Figure 1: Evolution of Insurer HHI and Premiums

Note: This figure shows the evolution of HHI and premiums from 2000 to 2018. The market is defined at the MSA level. Premiums are inflation-adjusted and weighted by enrollment within each MSA-year. Colors represent quintiles based on the distribution of HHI and premiums in the baseline year (2000), with darker shades indicating higher values.

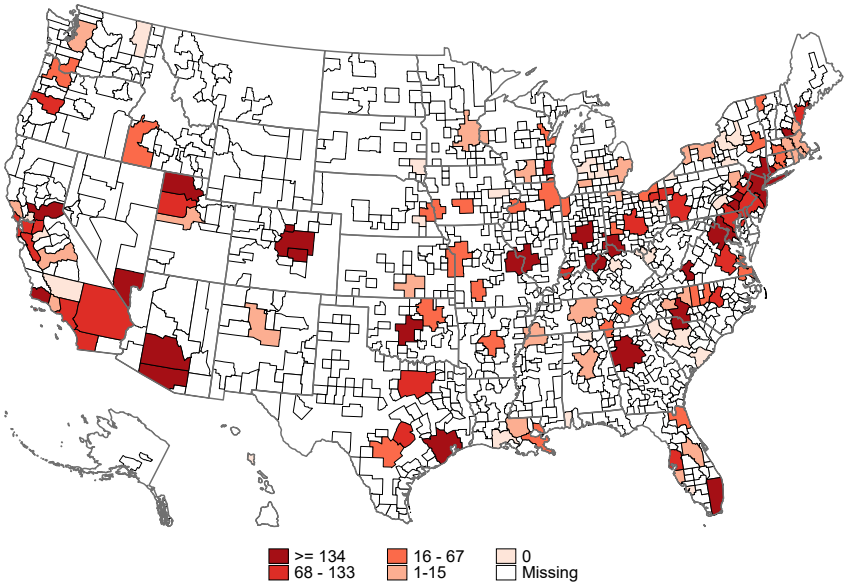
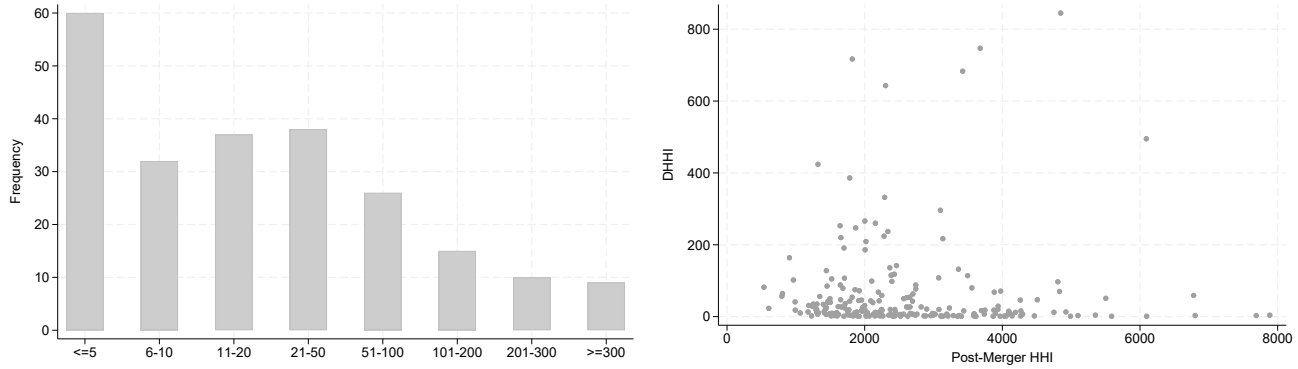


Figure 2: Cumulative Simulated Change in HHI

Note: This figure shows the cumulative simulated change in HHI (DHHI) induced by the top eight mergers used in our main analysis. Colors represent quintiles of the DHHI distribution. $DHHI = 2s_i s_j$, where s_i and s_j refer to market shares for the acquirer and target one year prior to a merger.

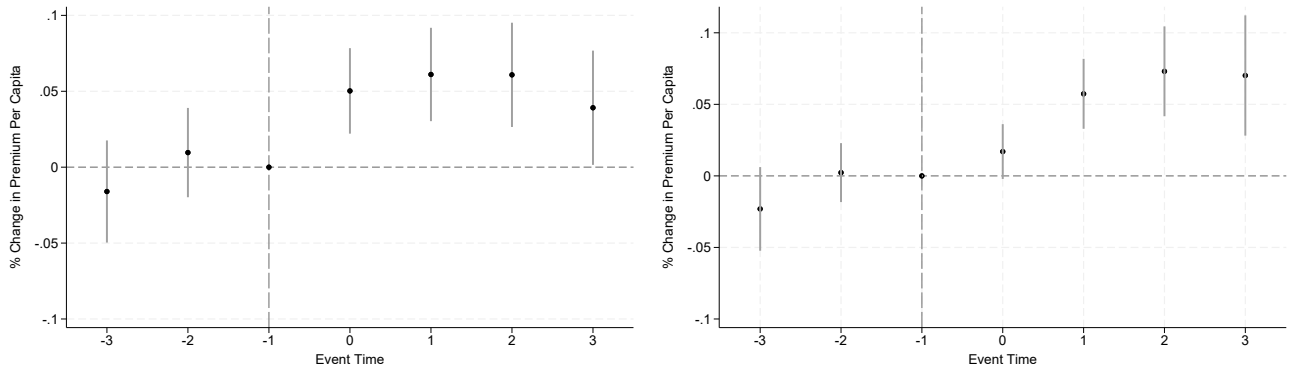


(a) Histogram of DHHI

(b) Scatter Plot of DHHI on Post-Merger HHI

Figure 3: Distribution of DHHI among Treated Markets

Note: Each observation is an MSA-year. Panel (a) plots the distribution of merger-induced DHHI among treated markets by the top eight mergers used in our main analysis. Panel (b) presents the scatter plot of DHHI and post-merger HHI among those treated markets. $DHHI = 2s_i s_j$, where s_i and s_j refer to market shares for the acquirer and target one year prior to a merger. Post-merger HHI is the HHI calculated in the first year after a merger took place.

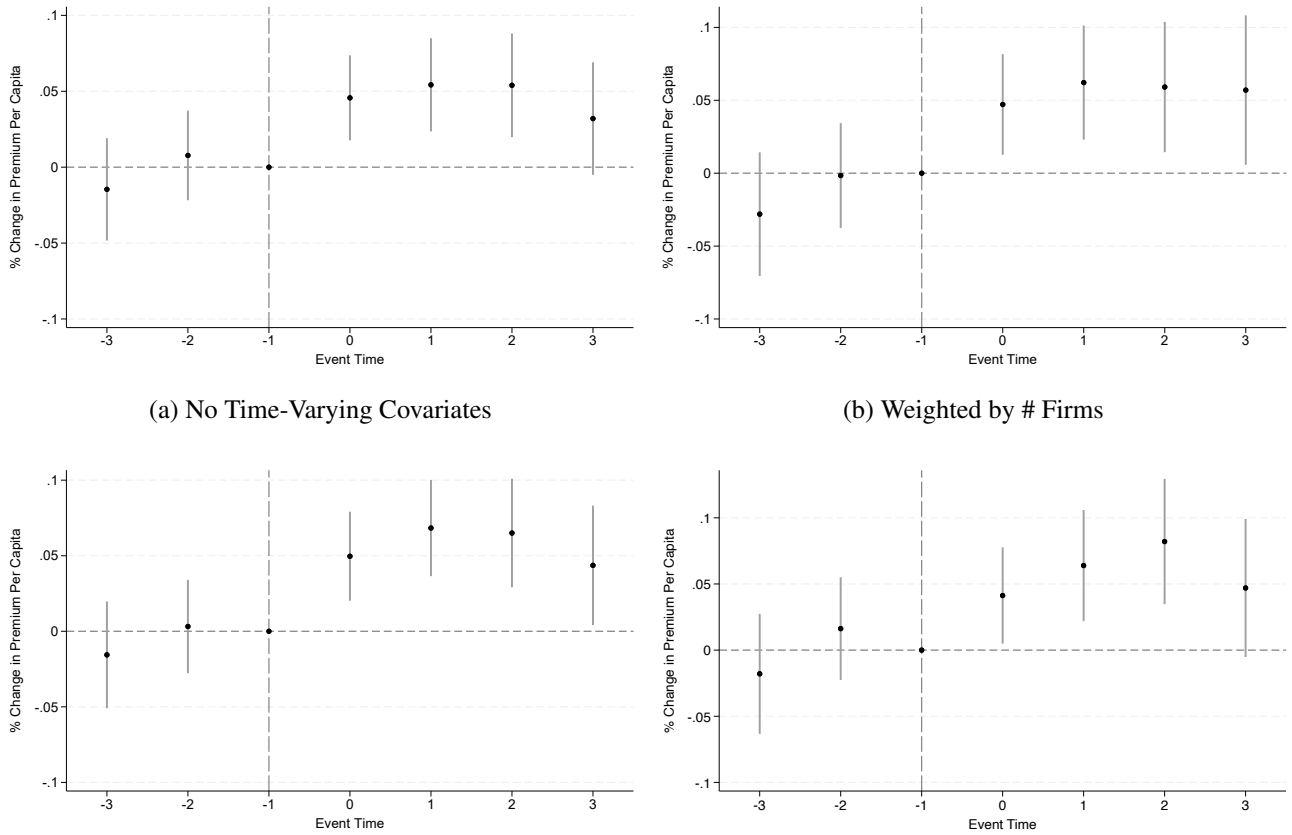


(a) Stacked DiD

(b) CS DiD

Figure 4: Event Study on the Effect of Mergers on Premiums

Note: This graph presents event study estimates of the impact of insurer mergers on the logarithm of per capita premiums, using the year prior to the merger as the baseline period. In Panel (a), we use a stacked DiD approach based on Equation (1), controlling for firm-by-cohort and year-by-cohort fixed effects, as well as contemporaneous plan-level controls and lagged market- and firm-level controls. Panel (b) presents the event study results following Callaway and Sant'Anna (2021) with both later-treated and never-treated as the control group. The coefficients are plotted along with a 95% confidence interval.



(a) No Time-Varying Covariates

(b) Weighted by # Firms

(c) Drop Early Treated

(d) Random Assignment of MSAs

Figure 5: Robustness of Stacked DiD Results

Note: These graphs show the robustness of our stacked DiD event study results. Panel (a) removes time-varying covariates on plan, firm, and market observables; Panel (b) is weighted by the average number of employers in an MSA throughout its sample period; Panel (c) drops markets that were treated earlier by the 2005/2006 mergers from the 2013 cohort; Panel (d) randomly assigns markets (that appeared multiple times across cohorts) to be either a treated or a control market for a single cohort only. The coefficients are plotted along with a 95% confidence interval.

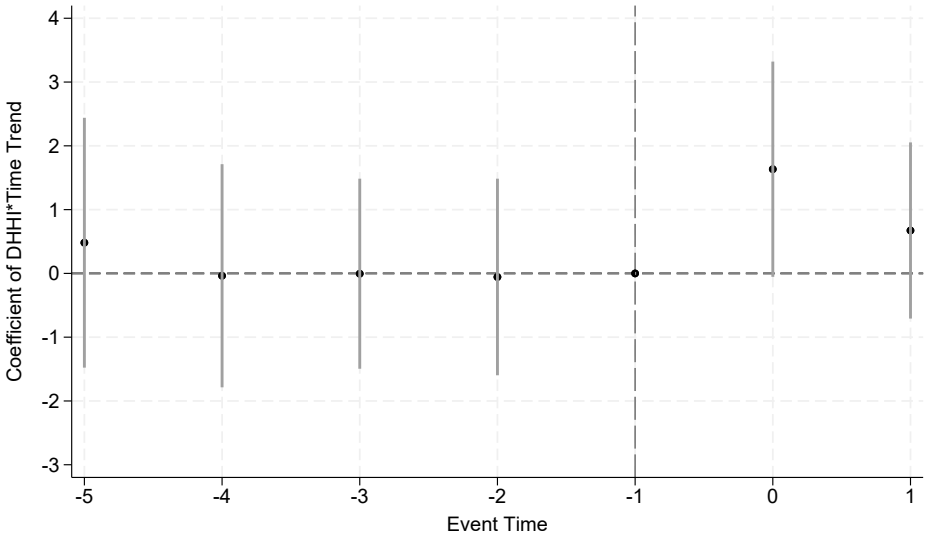


Figure 6: Test of Parallel Trends

Note: This graph presents results from a stacked DiD regression that interacts DHHI with dummies for event years. $DHHI = 2s_i s_j$, where s_i and s_j refer to market shares for the acquirer and target one year prior to a merger. The coefficients of the interaction terms are plotted along with a 95% confidence interval.

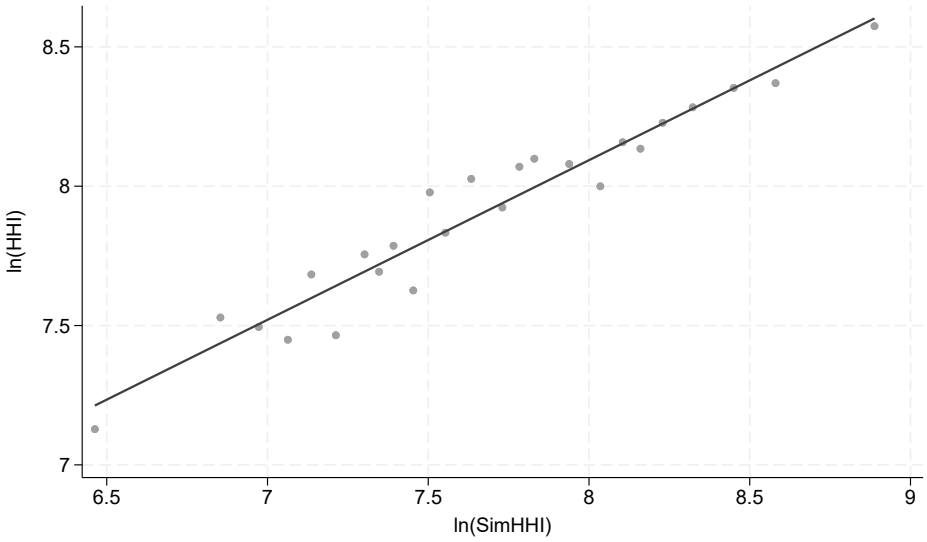


Figure 7: Correlation Between ln(HHI) and ln(SimHHI)

Note: This binscatter plot shows the correlation between ln(HHI) and ln(SimHHI). SimHHI is calculated by only exploiting changes in market shares between acquirers and targets from the top eight mergers of interest. The solid line represents the best linear fit.

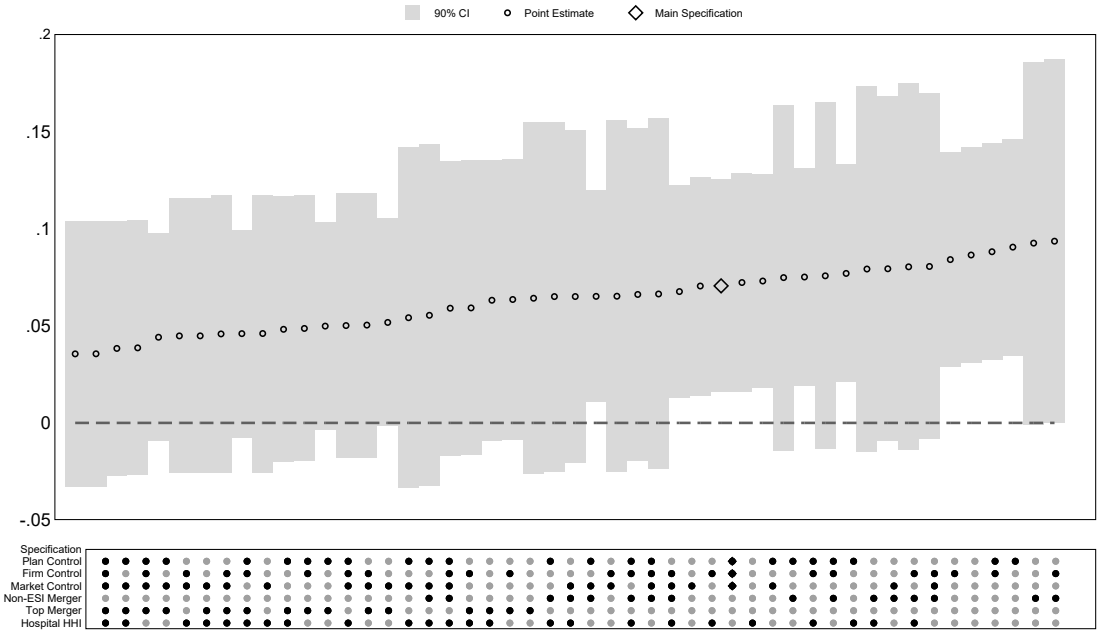


Figure 8: Specification Curve of Aggregate IV Estimates

Note: This figure presents a set of IV estimates for the aggregate effect of insurer concentration on premiums, along with 90% confidence intervals. The upper panel displays point estimates from different model specifications (circles), and the lower panel uses black dots to indicate which covariates are included in each corresponding regression. The diamond markers denote the main specification. Firm and year fixed effects are always included. Standard errors are clustered at the market-year level.

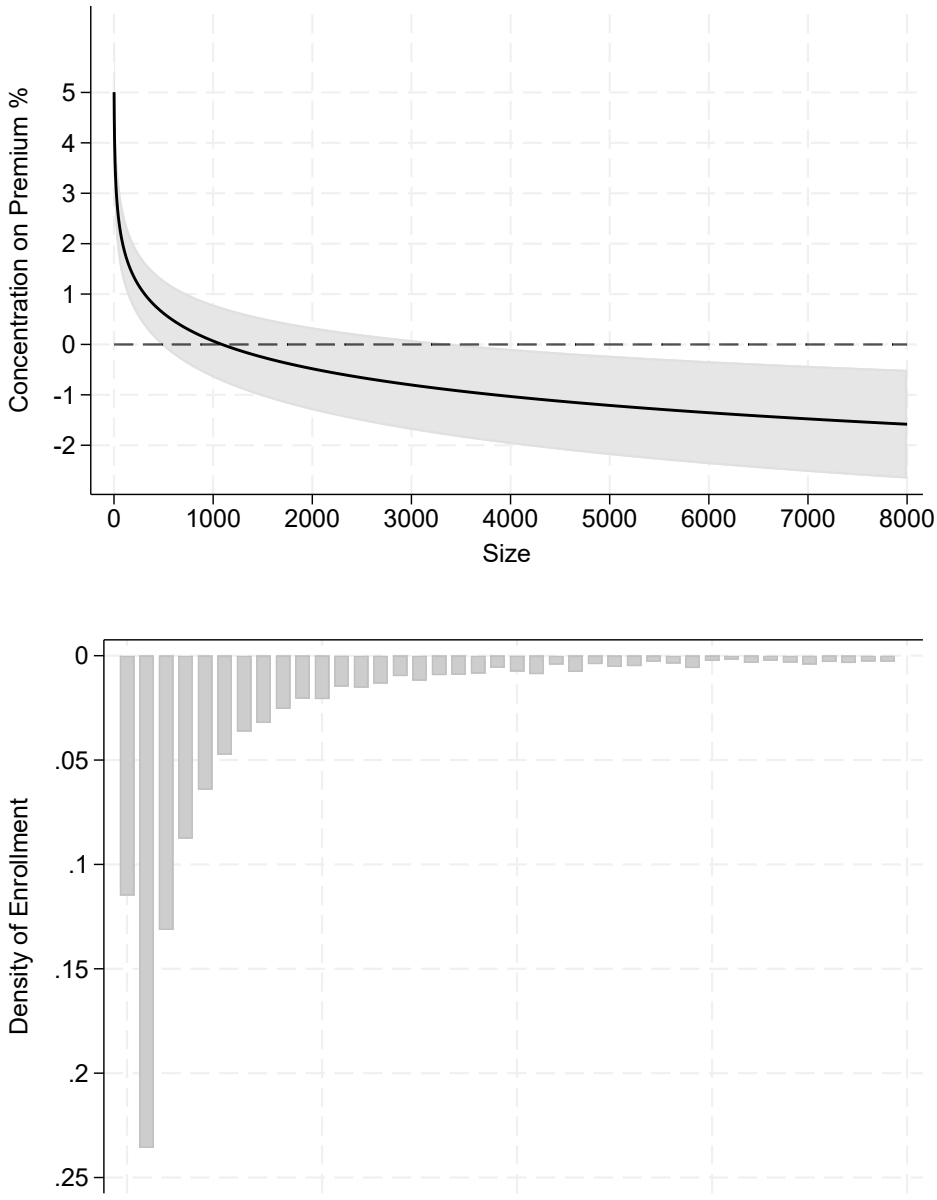


Figure 9: Heterogeneous Effect by Firm Size

Note: This graph shows the heterogeneous effect of insurer market concentration on premiums, alongside the distribution of enrollment by firm size. The upper panel plots the estimated effects for a 10% increase in insurer HHI on premiums by firm size based on Column (1) of Table 4. Firm size is defined as total enrollment averaged within a firm over time. The shaded gray area represents 95% confidence intervals calculated using the Delta Method. The lower panel plots the density distribution of firm enrollment across firms of different sizes (with a bin of 200). Firm size is truncated at 8,000 in this figure.

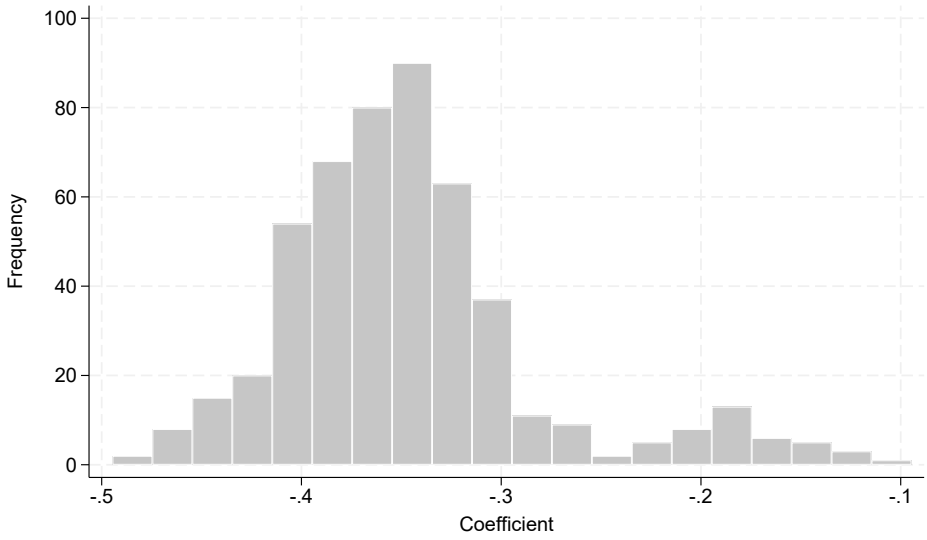


Figure 10: Results from Random Sampling of Large Firms

Note: This graph shows the robustness of the estimated effect of insurer concentration on premiums for large firms (reported in Column (3) of Table 4). In each draw, we randomly drop 10% of the large firms (by firm) from the main sample and replicate the regression. We repeat this process for 500 times and plot the distribution of the estimated coefficients.

Table 1: Summary Statistics

	All Firms		Mid-Sized		Large	
	Mean	SD	Mean	SD	Mean	SD
Plan						
Premium	5,303	2,891	5,337	2,880	4,906	2,993
Total Enrollment	756	3,884	377	422	5,271	13,051
Family Ratio	1.2	2.9	1.2	3	1.1	1.9
HMO (%)	38	49	37	48	52	50
PPO (%)	23	42	24	42	15	35
Indemnity (%)	3	16	3	16	3	17
Other (%)	7	25	7	25	7	25
Drug (%)	42	49	42	49	35	47
Dental (%)	13	33	13	33	13	33
Vision (%)	11	32	11	32	9	29
Market						
Insurer HHI	2,315	1,208	2,331	1,208	2,130	1,188
Hospital HHI	1,395	1,050	1,413	1,059	1,187	919
Unemployment Rate	6.1	2.3	6.1	2.3	5.9	2.2
Poverty Rate	13	4.7	13	4.7	13	4.8
Median Income	59,639	14,339	59,604	14,311	60,051	14,667

Note: This table reports descriptive statistics (mean and standard deviation) of key variables at the plan or market level, stratified by firm size. A plan is defined as a unique combination of firm, insurer, and plan type (e.g., HMO). The full sample contains 328,774 observations, with 303,282 from mid-sized firms (defined as having fewer than 1,500 enrollees averaged over the sample period) and 25,492 from large firms (with 1,500 or more enrollees). Family ratio is defined as the ratio of total enrollment over employee enrollment. HHIs are measured at the MSA-year level. Unemployment rate, poverty rate, and median household income are all measured at the county-year level.

Table 2: First Stage: Mergers and Market Concentration

	(1) MSA	(2) Plan
ln(SimHHI)	1.339***	0.889***
	(0.392)	(0.0888)
MSA FE	yes	no
Firm FE	no	yes
Year FE	yes	yes
Controls	no	yes
N	2,304	328,774
Adj R2	0.726	0.861

Note: This table presents the first-stage estimates of our IV specification in Equation (2). The dependent variable is $ln(HHI)$. Column (1) runs the analysis at the MSA level and Column (2) at the plan level. *** Significant at the 1% level. ** Significant at the 5% level. * Significant at the 10% level. Standard errors are shown in parentheses and clustered at the MSA-year level.

Table 3: Aggregate Effect of Insurer Concentration on Premiums

	(1) OLS	(2) Reduced	(3) IV
ln(HHI)	-0.0280***		0.0705**
	(0.00849)		(0.0333)
ln(SimHHI)		0.0626**	
		(0.0300)	
Firm FE	yes	yes	yes
Year FE	yes	yes	yes
Controls	yes	yes	yes
N	328,774	328,774	328,774

Note: Column (1) shows the OLS results for Equation (2). Column (2) shows the reduced-form results, and Column (3) shows the IV results. *** Significant at the 1% level. ** Significant at the 5% level. * Significant at the 10% level. Standard errors are shown in parentheses and clustered at the MSA-year level.

Table 4: Heterogeneous Effect of Insurer Concentration on Premiums

Sample:	(1) All	(2) Mid-Sized	(3) Large	(4) $\leq$1000	(5) $>$1000	(6) $\leq$2000	(7) $>$2000	(8) Mid-Sized	(9) Large
ln(HHI)	0.556***	0.118***	-0.350***	0.117***	-0.150*	0.116***	-0.412***	0.465***	1.734***
	(0.0793)	(0.0328)	(0.119)	(0.0326)	(0.0806)	(0.0333)	(0.115)	(0.0883)	(0.566)
ln(HHI) $\times$ ln(Size)	-0.0795***							-0.0589***	-0.257***
	(0.0130)							(0.0146)	(0.0692)
Firm FE	yes	yes	yes	yes	yes	yes	yes	yes	yes
Year FE	yes	yes	yes	yes	yes	yes	yes	yes	yes
Controls	yes	yes	yes	yes	yes	yes	yes	yes	yes
N	328,774	303,282	25,492	287,584	41,190	311,307	17,467	303,282	25,492

Note: This table presents heterogeneous effects of insurer concentration on premiums by firm size. The model largely follows the main specification of Equation (2) and is estimated using 2SLS. Mid-sized firms are defined as those with fewer than 1,500 enrollees averaged over the sample period, and large firms have enrollment above this threshold. Column (1) introduces an interaction term to allow the effect of insurer concentration to depend on firm size; Columns (2) and (3) separately study the effect of insurer concentration for the sample of mid-sized and large firms; Columns (4) and (5) use the threshold of 1,000; Columns (6) and (7) use the threshold of 2,000; Columns (8) and (9) replicate Column (1) by firm size. *** Significant at the 1% level. ** Significant at the 5% level. * Significant at the 10% level. Standard errors are shown in parentheses and clustered at the MSA-year level.

Table 5: Robustness of Main Results

Panel A: All Firms				
	(1)	(2)	(3)	(4)
	Drop Multi-Location	HHI Drop Large Firms	State as Market	HHI Include Self-Insured
ln(HHI)	0.0861***	0.124***	0.0164	0.0884***
	(0.0327)	(0.0449)	(0.0388)	(0.0336)
N	308,073	328,774	328,774	328,774
Panel B: Mid-Sized Firms				
	(1)	(2)	(3)	(4)
	Drop Multi-Location	HHI Drop Large Firms	State as Market	HHI Include Self-Insured
ln(HHI)	0.117***	0.164***	0.0419	0.127***
	(0.0327)	(0.0459)	(0.0397)	(0.0339)
N	289,448	303,282	303,282	303,282
Panel C: Large Firms				
	(1)	(2)	(3)	(4)
	Drop Multi-Location	HHI Drop Large Firms	State as Market	HHI Include Self-Insured
ln(HHI)	-0.314***	-0.284**	-0.127*	-0.259**
	(0.121)	(0.137)	(0.0763)	(0.110)
N	18,625	25,492	25,492	25492

Note: This table presents robustness checks for our aggregate and heterogeneous (by firm size) estimates of the effect of insurer market concentration on premiums following the main specification of Equation (2) and estimated using 2SLS. Mid-sized firms are defined as those with fewer than 1,500 enrollees averaged over the sample period, and large firms have enrollment above this threshold. Column (1) drops publicly listed firms that were ever located in multiple MSAs; Column (2) defines HHI by using firms with total enrollment below 1,500; Column (3) replicates the main results by using HHI calculated at the state-year level; Column (4) defines HHI by including self-insured plans. *** Significant at the 1% level. ** Significant at the 5% level. * Significant at the 10% level. Standard errors are shown in parentheses and clustered at the market-year level.

Table 6: Underlying Mechanisms

Panel A: Access to Self-Insurance						
	Less Credible Access			More Credible Access		
	(1)	(2)	(3)	(4)	(5)	(6)
	All	Mid-Sized	Large	All	Mid-Sized	Large
ln(HHI)	0.0975*	0.107**	-0.119	-0.0246	0.0413	-0.465***
	(0.0532)	(0.0533)	(0.184)	(0.0363)	(0.0359)	(0.141)
N	221,456	214,722	6,734	107318	88,560	18,758
Panel B: Hospital Market Competitiveness						
	Less Concentrated			More Concentrated		
	(1)	(2)	(3)	(4)	(5)	(6)
	All	Mid-Sized	Large	All	Mid-Sized	Large
ln(HHI)	0.0461	0.127**	-0.479***	0.118***	0.115***	0.156
	(0.0491)	(0.0493)	(0.170)	(0.0368)	(0.0378)	(0.127)
N	168,634	153,417	15,217	160140	149,865	10,275
Panel C: Merging Status						
	Merging			Non-Merging		
	(1)	(2)	(3)	(4)	(5)	(6)
	All	Mid-Sized	Large	All	Mid-Sized	Large
ln(HHI)	0.142**	0.202***	-0.446**	0.0265	0.0785**	-0.390***
	(0.0609)	(0.0632)	(0.216)	(0.0364)	(0.0348)	(0.128)
N	143,880	133,800	10,080	179278	164,092	15,186

Note: This table reports estimates of the effect of insurer market concentration on premiums by potential underlying mechanisms. In Panel A, we present results based on firms' access to self-insurance, which is calculated by predicting a self-insured ratio based on a linear regression of the actual self-insured ratio on firm size, industry fixed effects, MSA fixed effects, and market controls (i.e., unemployment rate, poverty rate, and median household income). Columns (2) and (5) divide mid-sized firms using a cutoff of the 70th percentile of the distribution of the predicted self-insured ratio among those firms; Columns (3) and (6) divide large firms using a cutoff of the 30th percentile of the distribution of the predicted self-insured ratio among large firms. Panel B tests how hospital market competitiveness affects the way premiums respond to insurer concentration. We divide markets based on hospital market HHI calculated in the year 2000, with the purpose of generating a roughly balanced sample between these two types of markets. Panel C shows the results by a plan's merging status, in which a plan is considered merging if its insurer has ever been involved in one of the eight mergers of our study's focus. Mid-sized firms are defined as those with fewer than $1,500$ enrollees averaged over the sample period, and large firms have enrollment above this threshold. All regressions follow the main specification of Equation (2) and estimated using 2SLS. *** Significant at the 1% level. ** Significant at the 5% level. * Significant at the 10% level. Standard errors are shown in parentheses and clustered at the market-year level.

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For Online Publication: Appendix

A1 Premium Data in Form 5500

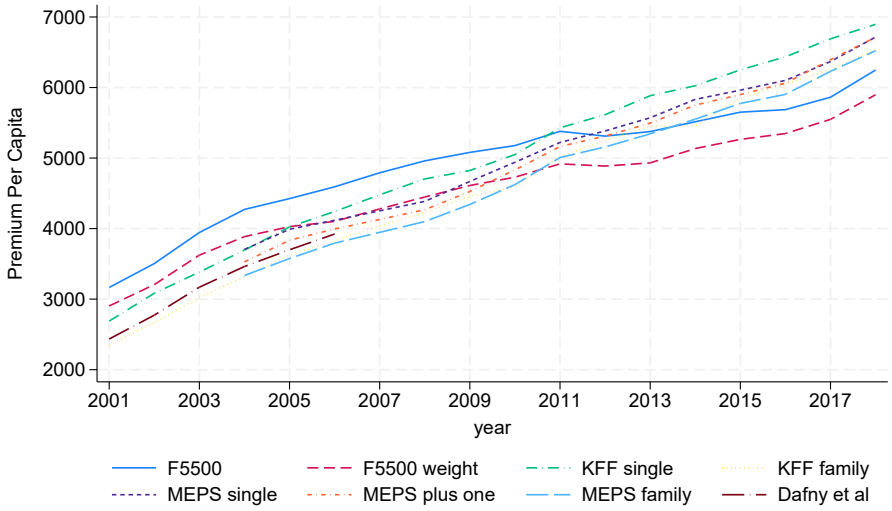


Figure A1: Premium Data Compared to Other Sources

Note: This figure compares our per-enrollee premium data to other data sources. We utilize two measures from Form 5500: a simple average across plans and an average weighted by enrollment, both within the 10th and 90th percentiles. "KFF" denotes the data from the Employer Health Benefits Survey conducted by the Kaiser Family Foundation. "MEPS" is the data from the Medical Expenditure Panel Survey Insurance/Employer Component. For comparison, we divide family plan premiums by 3 and single-plus-one plans by 2. "Dafny et al." refers to the data from the Large Employer Health Insurance Dataset from Dafny et al. (2012); because Dafny et al. (2012) used per-employee premiums, we divide it by the demographic factor of 2. This figure shows that our premium measures generally follow the trends from these different sources.

A2 Merger List

In our main analyses, we utilize the 8 mergers listed in Table A1. Our results are robust when using the largest 4 mergers (with at least 40 overlapping markets) and the largest 14 mergers (with at least 10 overlapping markets). Below, we offer a brief discussion on each of the listed mergers. As a quick summary, several mergers were driven by expansion in non-ESI markets (i.e., Cases 2, 3, 4, and 7), and some mergers were considered quite unexpected (e.g. Case 5).

Table A1: Merger List

Case	Target	Acquirer	Year	Event	Overlap	Value ($ in millions)
1	WellPoint Health Networks	Anthem	2004	2005	65	16,500
2	PacifiCare Health Systems	UnitedHealth	2005	2006	65	9,200
3	Coventry Health Care	Aetna	2013	2013	55	7,300
4	Great-West Healthcare	CIGNA	2008	2008	41	1,500
5	Oxford Health Plans	UnitedHealth	2004	2005	35	4,900
6	John Deere Health Care	UnitedHealth	2006	2006	26	500
7	Sierra Health Services	UnitedHealth	2008	2008	25	2,600
8	WellChoice	WellPoint	2005	2006	24	6,500

Note: This table lists the major merger events used in our main analysis. *Value* denotes the acquisition value of a merger. *Overlap* is the number of overlapping markets (among the 128 MSAs studied) between the acquirer and target in the year prior to the merger. *Year* indicates the year a merger was completed. *Event* is based on when during a year a merger was completed. Namely, if a merger was completed in the first half of a year, we assume HHI is affected in the same year, so the event year is the same as the merger year. However, when a merger was completed during the second half of a year, we assume HHI is affected the following year, and the event year is assigned to t+1.

1. The main goal of this merger was to reduce healthcare costs through economies of scale and through lowered reimbursement rates with providers. The merged entity was (re)named WellPoint Inc.

2. This merger had a significant impact on the Medicare market rather than on the fully-insured commercial market that is the focus of our study. It was expected to create the country's largest manager of Medicare plans at that time.

3. Aetna was mainly focused on commercial health insurance. The main purpose behind this merger was to allow Aetna to expand into government-held Medicare and Medicaid businesses.

4. The main goal of this merger was to expand CIGNA's business to be able to serve small groups' (2-99) health insurance needs.

5. Oxford was actively looking for an acquirer in order to save its declining business. The agreement between UnitedHealth and Oxford came almost immediately after Oxford's talks with WellChoice broke down. As such, it was unexpected for the marketplace.

6. This merger was intended to expand UnitedHealth's business in the Midwest market, where it had a limited presence. From the target's perspective, John Deere Health Care wanted to achieve economies of scale through a reduction in administrative costs and by negotiating rates with providers.

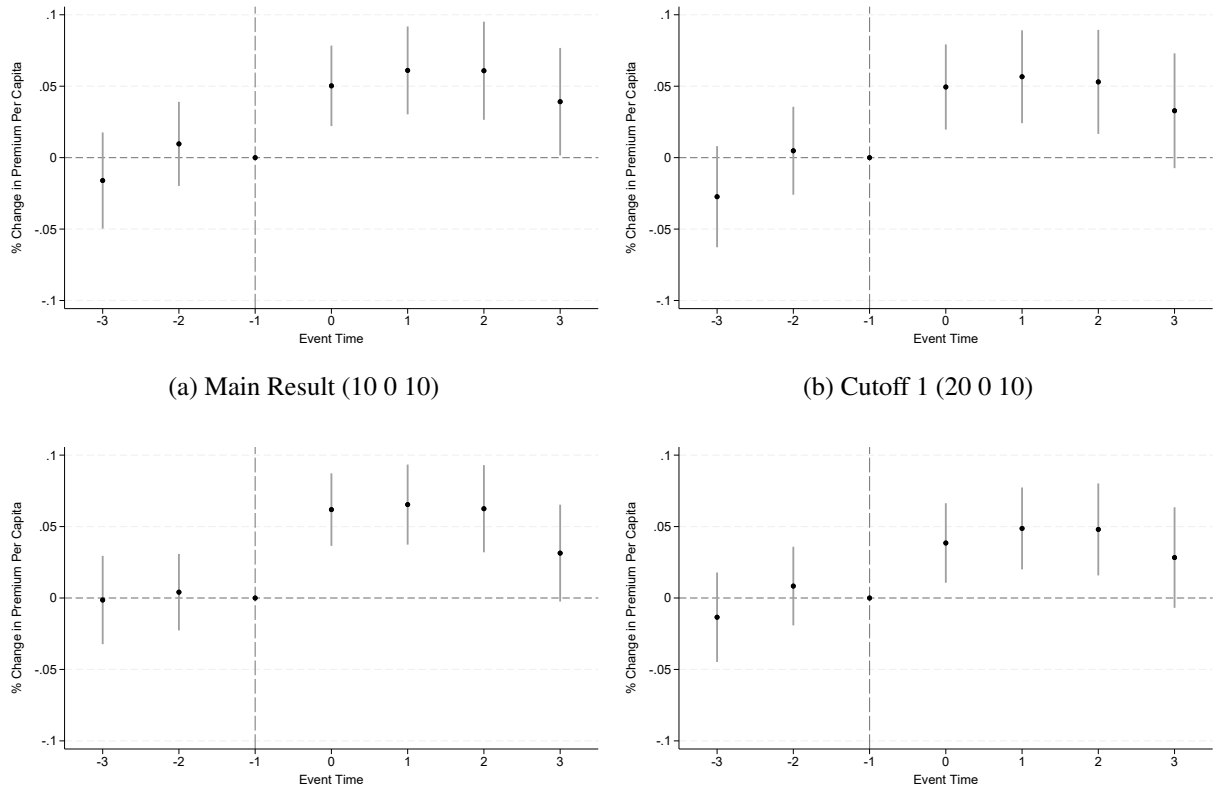
7. The impact of this merger mainly took place in the Medicare Advantage market. The Department of Justice asked UnitedHealth to divest its Medicare Advantage business in Las Vegas, Nevada.

8. WellPoint wanted to improve its leadership in businesses with large national employers, such as Fortune 500 companies. The merger also aimed to bring about synergies, such as reducing administrative costs while improving customer satisfaction.

A3 Event Study

A3.1 Applying Different Cutoffs

Here, we show how our event study results remain robust when applying different cutoffs in defining the treated and control groups. Specifically, three different cutoffs have been used in our main analyses when we compile a cohort-specific dataset. The first cutoff defines a treated market as those experiencing a DHHI of greater than 10 induced by the merger of interest. The second cutoff helps us to drop (both treated and control) markets that were simultaneously affected by other major mergers during a 10-year window of a given merger (i.e., a cutoff of 0 for a DHHI induced by other major mergers). The last cutoff helps to drop (both treated and control) markets that were contaminated by small mergers (i.e., a cutoff of 10 for a DHHI induced by small mergers).



(a) Main Result (10 0 10)

(b) Cutoff 1 (20 0 10)

(c) Cutoff 2 (10 5 10)

(d) Cutoff 3 (10 0 None)

Figure A2: Stacked DiD Event Study: Different Cutoffs

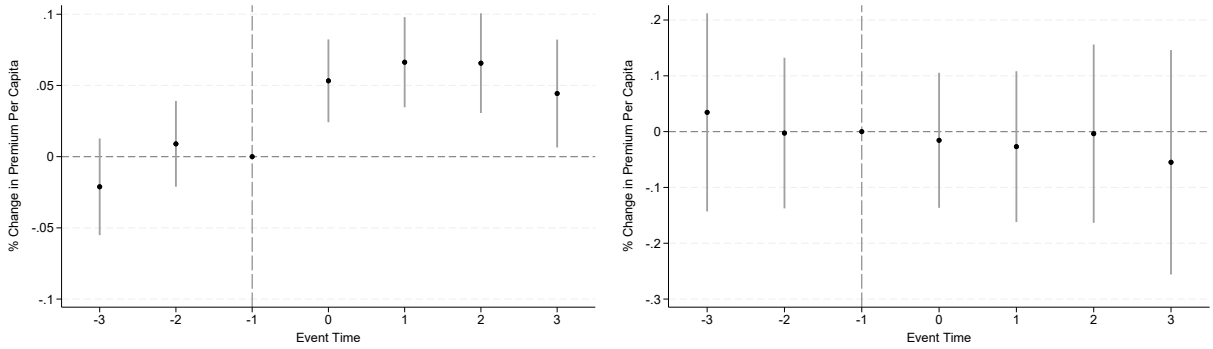
Note: These graphs show the robustness of our event study results with respect to different cutoffs in selecting treated and control groups. We use three different thresholds in defining treated and control groups in our main analyses. For example, our main results refer to the case where we select treated markets to be MSAs with $DHHI > 10$ for the focal event, and both treated and control markets with $DHHI = 0$ for the other major mergers during a 10-year window, while ensuring that neither the treated nor control groups were exposed to any smaller mergers with $DHHI \leq 10$. In Panels (b), (c), and (d), we relax the first, second, and third cutoffs, respectively.

We present three sets of results, with each dealing with one specific cutoff. Panel (a) replicates our main results. Panel (b) considers a scenario wherein the first cutoff is increased to 20. Panel (c) relaxes the restriction by setting the second cutoff to 5, which allows us to add some additional treated and

control markets to our analysis. Panel (d) disregards the last cutoff so that our sample choice is no longer constrained by small mergers. Overall, we find that all three sets of results remain largely consistent with our main findings.

A3.2 Event Study by Firm Size

We also offer an event study to examine heterogeneity by firm size. In the left panel, we focus on mid-sized firms (total enrollment averaged over time within a firm to be smaller than 1,500), and the results are similar to our main findings. That is, we find that mid-sized firms see an increase of about 5% in premiums. Interestingly, however, when focusing on large firms, we find no evidence that mergers lead to any significant changes in premiums.



(a) Mid-Sized Firms

(b) Large Firms

Figure A3: Stacked DiD Event Study by Firm Size

Note: This graph plots event study estimates separately for mid-sized and large firms. Panel (a) shows the event study coefficients for mid-sized firms, and Panel (b) for large firms. The analysis follows the main model specification in Equation (1). Firm size is defined based on total enrollment averaged over time within a firm, with a cutoff of 1,500 enrollees.

A4 Results Concerning Measurement Error in Enrollment

In this section, we address concerns related to noisy reporting of total enrollment in Schedule A. We offer two sets of results to show that our main results remain consistent.

Evidence from Employer-Level Analyses We first report results obtained when we use the employee count (from the main form of Form 5500) in calculating per-employee premium and in defining firm heterogeneity (i.e., firm size). More specifically, we sum up all premiums across Schedule As within a firm–year and divide that by the number of employees reported on the main form. This approach allows us to get around the issue of misreporting in the enrollee count for individual Schedule As. The results, reported in Panel A of Table A2, are largely consistent with our main findings. Note that these analyses are conducted at the firm–year level instead of plan–year level.

Table A2: Results Addressing Measurement Error in Enrollment

Panel A: Analyses at Employer-Level			
	(1)	(2)	(3)
	All	Mid-Sized	Large
ln(HHI)	0.0596*	0.0677*	-0.246*
	(0.0346)	(0.0356)	(0.132)
N	212,542	204,736	7,806
Panel B: Analyses Using Family Ratio			
	(1)	(2)	(3)
	All	Mid-Sized	Large
ln(HHI)	0.0693*	0.102***	-0.388**
	(0.0366)	(0.0368)	(0.154)
N	262,845	246,524	16,321

Note: This table presents results that address measurement error in enrollment. Panel A uses the number of employee participants reported on the main form to measure per-employee premiums and employer size. We aggregate premiums and plan-level controls to the employer-year level. Panel B restricts the sample to plans with an adjusted family ratio of at least 1. Results are first presented with the aggregate effect, and then divided into mid-sized and large firms, where mid-sized firms are defined as those with fewer than 1,500 enrollees averaged over the sample period, and large firms have enrollment above this threshold.. All regressions follow the main specification of Equation (2) and estimated using 2SLS. *** Significant at the 1% level. ** Significant at the 5% level. * Significant at the 10% level. Standard errors are shown in parentheses and clustered at the market-year level.

Evidence from Using the Family Ratio We use a strategy adapted from the one proposed in Craig (2022). Specifically, we first figure out the average family ratio (defined as the total number of health insurance enrollees divided by the number of health insurance employee participants) in the ESI population using MEPS. We then use this average family ratio from MEPS to guide us in choosing our sample accordingly, so that the average family ratio in our sample would indeed represent the broader population.

According to MEPS, over the years 2001–2018 and among private-sector establishments, the average family ratio in the ESI population is 1.805 if assuming 3 enrollees for family plans, and 2.115 if assuming 4 enrollees for family plans. We believe that a cutoff of 2 for the average family ratio would be appropriate, given that an average family with kids has 1.9 children (based on statistics from the U.S.

Census Bureau).

To be able to back out the family ratio from our data, we must first establish a better idea of the extent to which the employee participant count that is reported on the main form might be exaggerating the number of employee health plan participants. Based on estimates from the U.S. Bureau of Labor Statistics, conditional on having access to both life and medical insurance, the take-up rate for life insurance is 98% and the take-up rate for medical insurance is 68%.[29] As a result, if an employer offers both life and medical insurance, the employee participant ratio for life insurance over medical insurance is estimated to be 1.44 (98/68). Assuming that the total number of employee participants reported on the main form represents the universe of those who take up life insurance (given that nearly 100% do so), the number of employee participants for health insurance is the total employee participant count divided by 1.44. For employers that do not offer life insurance coverage (about 25% of the sample), we take a more conservative approach by assuming that the employee participant count reported represents the real employee participant count for health plans.

In the end, by setting a minimum cutoff of 1 for the family ratio, we end up with a sample of plans with family ratios ranging from 1.37 (25th percentile) to 2.52 (75th percentile), with a mean of 2.08. We notice that this sample distribution is rather close to what we have calculated from the broader population, based on the MEPS data.

The regression results are reported in Panel B of Table A2, where we drop about 20% of the sample from our main analysis. These results are also largely similar to our main findings. Not reported here, our results are largely consistent if using alternative cutoffs (e.g., 1.5 and 2) for the family ratio.

A5 Results Studying Other Outcomes

We report detailed results regarding how insurer competition affects other market outcomes in Table A3.

First, we study how plan characteristics respond to changes in insurer concentration. The results are reported in Columns (1)-(5) of Table A3. Regarding the coverage of dental/vision/drug benefits, we first note that those measures indicate whether a health plan has carved each coverage in. Employers can choose carved-out coverage for dental/vision/drug benefits (i.e., covered separately from a health plan). In this case, those benefits can still be offered - they are just not included within the coverage of health plans that we study in our analyses. We find the results to be largely consistent among mid-sized and large firms, although the magnitude of the estimated effects tends to be larger for large firms.

We next explore a firm's total enrollment and these results are reported in Column (6) of Table A3. As we noted in the main text, these quantity results are consistent with the heterogeneous premium effects identified in our main findings.

We last study whether firms switch their funding type. Following the Annual Report to Congress on Self-Insured Group Health Plans, we assign an employer to one of the three funding types: self-insurance (about 37% of firm-year observations), full-insurance (about 55% of firm-year observations), and mixed-insurance (about 8% of firm-year observations), which is defined as a mix of self- and full-insurance.[30] We use a linear probability model; when we study the probability of self-insurance, the outcome variable equals 1 for being self-insured and 0 for being mixed- or fully-insured. We follow the same strategy when studying the probability of full-insurance. These two sets of results, reported in

[29] Life insurance take-up can be found here, and medical insurance take-up can be found here.

[30] This distribution is largely consistent with the KFF Benefit Survey data [Kaiser Family Foundation 2020a]. If weighted by enrollment, about 24% of employees were fully-insured, 43% self-insured, and 33% mixed-insured. The estimate from the KFF survey had a slightly higher share of about 50-60% self-insured employees, though KFF does not code mixed-insured plans separately.

Columns (7) and (8), tell coherent stories. We find that an increase in insurer HHI tends to encourage firms to switch from self-insurance to full-insurance or mixed-insurance and that these results are largely driven by large firms.[31] Specifically, we find that a 10% increase in insurer HHI decreases the probability of self-insurance among large firms by 1.3%, as compared to a mean self-insurance rate of 51.9% among those firms.

Table A3: Other Outcomes

Panel A: All Firms								
	(1) HMO	(2) PPO	(3) Drug	(4) Dental	(5) Vision	(6) Quantity	(7) Self	(8) Full
ln(HHI)	0.0607**	-0.0275	-0.0536**	0.0736***	0.0183	-0.0179	-0.0438**	0.0458***
	(0.0271)	(0.0193)	(0.0264)	(0.0188)	(0.0128)	(0.0512)	(0.0181)	(0.0158)
N	328,774	328,774	328,774	328,774	328,774	227,870	574,397	574,397
Panel B: Mid-Sized Firms								
	(1) HMO	(2) PPO	(3) Drug	(4) Dental	(5) Vision	(6) Quantity	(7) Self	(8) Full
ln(HHI)	0.0939***	-0.0399*	-0.0507*	0.0665***	0.0117	-0.0636	-0.0199	0.0245
	(0.0299)	(0.0215)	(0.0274)	(0.0192)	(0.0133)	(0.0452)	(0.0186)	(0.0176)
N	303,282	303,282	303,282	303,282	303,282	217,851	473,800	473,800
Panel C: Large Firms								
	(1) HMO	(2) PPO	(3) Drug	(4) Dental	(5) Vision	(6) Quantity	(7) Self	(8) Full
ln(HHI)	-0.229***	0.0910**	-0.0887	0.136***	0.0737**	0.194	-0.131***	0.117***
	(0.0678)	(0.0425)	(0.0651)	(0.0493)	(0.0335)	(0.301)	(0.0323)	(0.0272)
N	25,492	25,492	25,492	25,492	25,492	10,019	100,597	100,597

Note: This table examines whether insurer market concentration affects other outcomes. Columns (1) - (5) test how insurer competition affects plan characteristics, measured by whether a plan is HMO/PPO and whether it covers drug/dental/vision benefits. Column (6) tests how insurer competition affects (the logarithm of) total enrollment, aggregated at the firm-year level. Columns (7) and (8) take a linear probability model approach and test how insurer competition affects the probability of self- or full-insurance at the extensive margin. This sample covers both self- and fully-insured firms in the data. This model follows the main specification in Equation (2), but excludes plan characteristics as controls in all columns except Column (6). Mid-sized firms are defined as those with fewer than 1,500 enrollees averaged over the sample period, and large firms have enrollment above this threshold. All regressions follow the main specification of Equation (2) and estimated using 2SLS. *** Significant at the 1% level. ** Significant at the 5% level. * Significant at the 10% level. Standard errors are shown in parentheses and clustered at the market-year level.

[31] Note that here we measure firm size using employee participants (averaged within a firm over our study period) since we lack a measure of total enrollment for a large fraction of self-insured employers. We use a cutoff of 1,500 to define large v.s. mid-sized firms, and these results are consistent with alternative cutoffs, such as 1,000 or 2,000.

A6 Results Explaining the Self-Insured Ratio

We present results from our regression on the self-insured ratio separately by firm size. We run the analysis on a cross section of firms in Form 5500, with the dependent variable being the number of years of self-insurance (N_{self}) over the total number of years that we observe a firm in our data ($N_{self} + N_{full}$). We find that size positively predicts the use of self-insurance for both mid-sized and large firms, whereas economic downturns seem to have the opposite effect. Our results also hold when we use alternative specifications, including when we weigh each employer by the number of years it showed up in the data ($N_{self} + N_{full}$), or when we restrict to employers who appeared multiple times in the data ($N_{self} + N_{full} \geq 2$ or ≥ 3).

Table A4: Self-Insured Ratio Regression

Panel A: Mid-Sized Firms				
	(1)	(2)	(3)	(4)
Size [in millions]	11.48***	12.50***	12.55***	12.56***
	(0.607)	(0.383)	(0.418)	(0.446)
Unemployment Rate [0-1]	-2.061***	-2.447***	-2.539***	-2.699***
	(0.0999)	(0.119)	(0.136)	(0.157)
Poverty Rate [0-1]	-2.121***	-1.626***	-1.571***	-1.512***
	(0.0638)	(0.0653)	(0.0732)	(0.0809)
Median Income [in millions]	-8.209***	-6.935***	-6.785***	-6.689***
	(0.248)	(0.257)	(0.288)	(0.319)
MSA FE	yes	yes	yes	yes
Industry FE	yes	yes	yes	yes
N	119,493	119,493	95,541	78,923
Panel B: Large Firms				
	(1)	(2)	(3)	(4)
Size [in millions]	0.679***	0.467***	0.459***	0.451***
	(0.0916)	(0.0626)	(0.0678)	(0.0713)
Unemployment Rate [0-1]	1.003***	0.666	0.497	0.235
	(0.362)	(0.440)	(0.496)	(0.556)
Poverty Rate [0-1]	-0.0509	-0.566**	-0.591**	-0.632**
	(0.231)	(0.225)	(0.247)	(0.265)
Median Income [in millions]	0.573	-1.551*	-1.727*	-2.106**
	(0.838)	(0.837)	(0.918)	(0.988)
MSA FE	yes	yes	yes	yes
Industry FE	yes	yes	yes	yes
N	8,188	8,188	6,762	5,896

Note: This table presents linear regression results for the self-insured ratio, which is calculated as a firm's number of years of self-insurance over its total number of years in the sample. We include firm size, industry fixed effects, MSA fixed effects, and market controls (i.e., unemployment rate, poverty rate, and median household income) as explanatory variables. Column (1) is unweighted. Column (2) is weighted by the number of years that a firm appears in the sample. Column (3) restricts itself to firms that showed up at least twice in the data, and Column (4) at least three times in the data. Mid-sized firms are defined as those with fewer than $1,500$ enrollees averaged over the sample period, and large firms have enrollment above this threshold. *** Significant at the 1% level. ** Significant at the 5% level. * Significant at the 10% level. Standard errors are shown in parentheses.

A7 Model

In light of our empirical findings, we develop a stylized model whose objective is to highlight two key factors in explaining our reduced-form results: 1) access to self-insurance as an outside option, and 2) hospital prices. Our model extends the Nash-in-Nash with the threat of replacement (NNTR) concept developed in Ho and Lee (2019) to solve the equilibrium price (premium) in the ESI market. The NNTR concept is built upon the Nash-in-Nash solution as proposed by Collard-Wexler et al. (2019), and has the advantage of allowing competing insurers to influence negotiated prices. This feature is particularly useful for understanding the impact of insurer consolidation. We show that under certain circumstances, this model yields predictions that are consistent with our empirical findings.

A7.1 Setup

We are primarily concerned with the determination of equilibrium transfer (premiums) between insurers and employers in the fully-insured ESI market. We consider that an employer is committed to contracting with one single insurer. To the employer, the valuation of an existing plan is W^G, and the employer pays the premium ϕ per enrollee. Insurers compete to attract employers to enroll in their plans. Without loss of generality, we assume that the best outside option in the fully-insured market is given by a competing insurer F, which offers a utility of W^{OF}. There is also an option of self-insurance, denoted by W^{OS}, which is only available to large employers.

Applying the NNTR solution developed by Ho and Lee (2019) to our setting, the equilibrium premium for the incumbent plan is as follows:

$$\phi = min\{\phi^N, \phi^{OO}\} \tag{3}$$

where ϕ^N is the Nash-in-Nash bargaining outcome with the incumbent insurer and ϕ^{OO} is the outside option price, which presents the lowest price that the employer could pay to the incumbent insurer so that the employer would be indifferent between staying with its incumbent insurer or replacing it with a competing one. Different from the Nash-in-Nash solution, in the NNTR solution, the employer can threaten to replace the incumbent insurer with a competing insurer that is currently excluded. The existence of a credible threat allows the employer to obtain additional leverage in its negotiation with the insurer, which, in turn, imposes a cap on the Nash-in-Nash price.

We now explain how ϕ^N and ϕ^{OO} are determined. The Nash-in-Nash price ϕ^N solves:

$$\phi^N = \underset{\phi}{\text{argmax}} (\phi - p)^{\tau} \times (W^G - \phi)^{1-\tau}$$

where p denotes provider price and $\pi_M = \phi - p$ is the insurer profit.[32] $\pi_E = W^G - \phi$ is the employer surplus, with τ denoting the relative bargaining power of the insurer. Solving the first-order conditions, we get:

$$\phi^N = \tau W^G + (1-\tau)p \tag{4}$$

The Nash-in-Nash equilibrium price is a weighted average of the employer's valuation and the cost to the insurer, with the weight given by the relative bargaining power. When τ equals 1, the insurer has all of the bargaining power and the negotiated price equals the employer's valuation for the incumbent

[32] p can be easily extended to include other insurer costs but we focus on provider costs for simplicity.

plan. When τ equals 0, the negotiated price equals the hospital cost, as the insurer will not accept a price lower than its marginal cost.

The outside option price ϕ^{OO} solves:

$$\pi_E^G(\phi^{OO}) = \max[\pi_E^{OF}, \pi_E^{OS}]$$

where $\pi_E^G = W^G - \phi^{OO}$, $\pi_E^{OF} = W^{OF} - \phi_{res}^{OF}$, and $\pi_E^{OS} = W^{OS} - \phi_{res}^{OS}$. ϕ_{res}^{OF} and ϕ_{res}^{OS} are reservation prices for the outside fully-insured and self-insured plans respectively. The reservation prices are defined as the minimum prices that the competing insurers would accept to be included in the employer's choice set. In our setting, the *OF* and *OS* reservation prices are just p because an insurer would not accept a price lower than its cost.[33] As a result, we get

$$\phi^{OO} = W^G - \max[W^{OF}, W^{OS}] + p \tag{5}$$

Plugging Equations (4) and (5) into Equation (3), the equilibrium price can be solved as:

$$\phi = min\{\tau W^G + (1-\tau)p, W^G - \max[W^{OF}, W^{OS}] + p\}$$

Note that the above equilibrium price applies to large employers with the available outside option of self-insurance. For mid-sized employers, self-insurance might not be a relevant option, but one can set W^{OS} to zero.

A7.2 How Mergers Affect Equilibrium Prices

First, consolidation could lead to changes in provider prices p. Existing empirical studies, including Moriya et al. (2010), McKellar et al. (2014), Roberts et al. (2017), Scheffler and Arnold (2017), and Cooper et al. (2019), have found evidence that hospital/physician prices are negatively associated with insurer market power. We discuss below that lowered hospital price is a necessary condition in explaining the price patterns observed in our empirical setting. Second, consolidation could lead to a change in bargaining leverage. Particularly, mergers could hurt an employer's outside option and therefore eliminate or weaken the threat of replacement. Consider a simple scenario where a merger leads to the removal of the best outside option. The resulting price might go up because the employer can no longer threaten its current insurer as credibly as with the previous outside quote.

A7.3 Explaining the Heterogeneous Effect of a Merger

To illustrate how insurer consolidation affects the equilibrium price differently for mid-sized and large firms, we first use a stylized example. We assume $\tau = 0.5$, $W^G = 10$, $W^{OF} = 8$, $W^{OS} = 7.5$, and $p = 4$. Before the merger, the NNTR price is $\phi = min\{7, 6\} = 6$, where the outside option poses a credible threat and sets a cap on the equilibrium price. After the merger, p decreases from 4 to 3 because insurers can engage in more favorable bargaining with hospitals. However, the outside option of full insurance is eliminated due to the merger for mid-sized firms. As a result, employers pay the Nash-in-Nash price $\phi = 6.5$, which splits the gains-from-trade equally between employers and insurers. In contrast, large

[33] It might not be obvious that an employer can credibly threaten to pay the outside option its reservation price. Following the argument of Ho and Lee (2019), we consider that employers can commit to negotiating with a given insurer but they can go back and forth between insurers when negotiating; this allows them to play insurers against each other and, hence, credibly use paying the reservation price as a threat. Please refer to Ho and Lee (2019) regarding the existence and uniqueness of the NNTR solution.

employers keep the option of self-insurance, which poses a credible threat and, therefore, leads to a lower price. In this case, we have $\phi = min\{6.5, 5.5\} = 5.5$. In this example, mid-sized employers experience premium increases due to the loss of the threat of replacement, whereas large employers generate premium savings because the purchase power effect (through reduced hospital prices) dominates the market power effect (through the reduction in outside options) when the outside option of self-insurance is still available.

To generalize, suppose that the cost changes due to a merger are Δp ($\Delta p < 0$ for a reduction in provider prices). We consider that the premiums before and after a merger are

$$\phi = \min\left\{\phi^N = \tau W^G + (1-\tau)p, \quad \phi^{OO} = W^G - \max\left[W^{OF}, W^{OS}\right] + p\right\}$$

and

$$\tilde{\phi} = \min\left\{\widetilde{\phi^N} = \phi_L^N + (1-\tau)\Delta p, \right.$$
$$\left. \widetilde{\phi^{OO}} = \phi^{OO} + \max\left[W^{OF}, W^{OS}\right] - \max\left[W^{OF'}, W^{OS}\right] + \Delta p\right\}$$

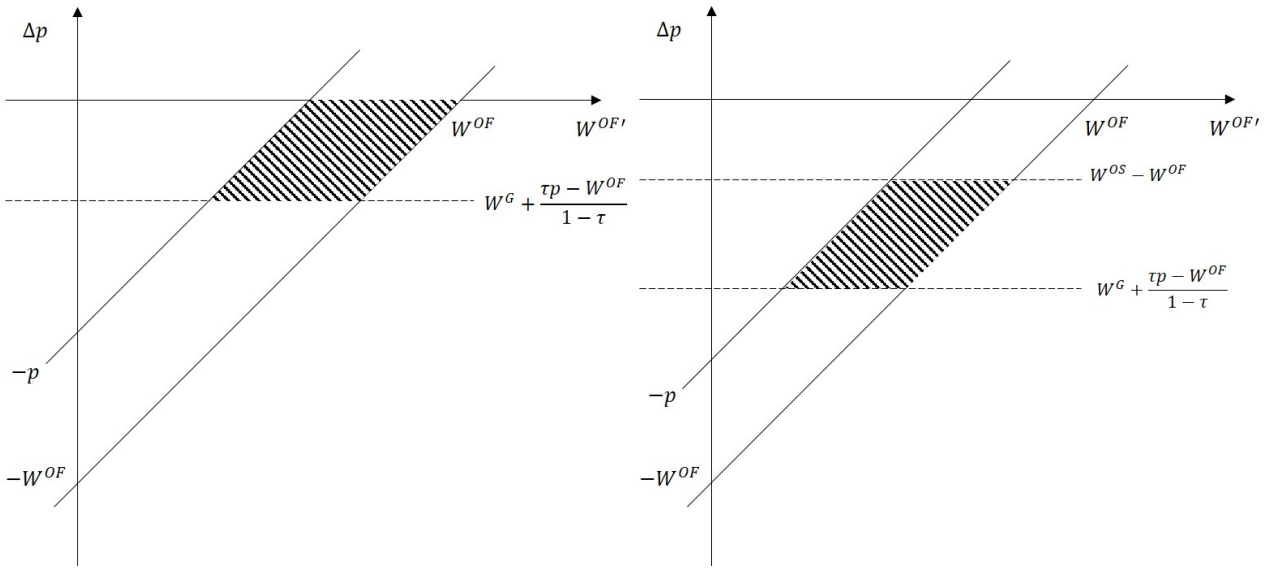
Note that for the Nash-in-Nash bargain, insurers and employers share the surplus of lowered hospital prices, with the weight given by their relative bargaining power. For the outside-option price, employers can take full advantage of the lowered hospital prices, although these, however, are offset by the reduction in outside options. Before the merger, the best alternative is W^{OF} for fully-insured plans. A merger leads to the elimination of W^{OF}, and instead $W^{OF'}$ becomes the best alternative post-merger. Employers differ in their capabilities to use self-insurance as a credible threat.

To explain lowered premiums for large employers, we first note that hospital prices must decrease post-merger (i.e., $\Delta p < 0$). The proof is straightforward. If hospital prices increase, we would expect an increase in premiums post-merger because both the Nash-in-Nash price ϕ^N and the best outside option price ϕ^{OO} would go up as a result of increased hospital prices. This contradicts our empirical finding that premiums actually decrease for large firms.

In what follows, we discuss the necessary conditions for the model to generate the price patterns that we observe in the data. We discuss two scenarios. The first scenario occurs when self-insurance is preferred (i.e., $\max\left[W^{OF}, W^{OS}\right] = W^{OS}$). In this case, a merger does not affect the outside option. Panel (a) in Figure A4 depicts the relationship between Δp and $W^{OF'}$ while holding the other parameters fixed. The highlighted area defines the conditions under which our model would predict a heterogeneous price effect. The second scenario occurs when self-insurance is dominated, i.e., $\max\left[W^{OF}, W^{OS}\right] = W^{OF}$. In this case, the pre-merger price stays the same for mid-sized and large employers due to the same best outside option of full-insurance. Panel (b) in Figure A4 summarizes this scenario and highlights the area under which these conditions are satisfied. Similar to the first scenario, the intuition is that cost reduction Δp needs to be small enough (or big enough in absolute value) to lead to a price reduction for large firms. However, at the same time, it should not be too small to lead to a lowered premium for mid-sized firms.

Extension We briefly consider two extensions. First, we have so far focused on a case when mid-sized and large firms have the same bargaining power parameter. Allowing large firms to have more bargaining power actually facilitates the price-decreasing effect because these firms can enjoy a larger fraction of savings due to lowered hospital prices. As a result, differences in the bargaining power serve to relax restrictions on the model. Second, it is possible that the self-insured option would change as

a result of mergers. We expect costs for self-insured plans to decrease post-merger, which makes the outside option of self-insurance even more attractive to large firms and therefore leads to even lower premiums for those firms in the fully-insured market.



(a) $W^{OF} < W^{OS} < W^G$ (b) $W^{OS} < W^{OF} < W^G$

Figure A4: Circumstances Predicting a Heterogeneous Effect

Note: The first scenario is when the option of self-insurance is preferred to other fully-insured alternatives. The second scenario is when self-insurance is dominated by the best alternative among fully-insured plans.

A7.4 Evidence from the Data

We present empirical evidence in support of the two main assumptions utilized in the model. The first assumption is that hospital/provider prices decrease as a result of higher insurer concentration. This assumption is consistent with some existing findings as discussed in Section A7.2. In addition to drawing insights from previous research, we offer direct evidence from our setting by leveraging the richness of our data. In particular, we observe a sample of "experience-rated" plans. Different from the plans that constitute our main analytic sample, premiums of those "experience-rated" plans more closely reflect enrollees' underlying health care costs (i.e., incurred claims). Had insurer concentration led to lowered provider prices, we would expect incurred medical costs, as measured using insurance claims, to decrease. Premiums might also decrease due to lowered costs of care and a direct linkage between claims and premiums for those plans. To proceed, we utilize a sample of about 22,000 distinct "experience-rated" plans from our data. This analysis follows our main specification and the results are reported in Table A5 Columns (1)-(3) for premiums and Columns (4)-(6) for claims. We find that a higher level of insurer concentration leads to lowered premiums and claims, and these results hold true regardless of firm size. Our finding that premiums decrease for mid-sized firms is particularly interesting; we interpret these results as compelling evidence in support of lowered hospital prices in the wake of a more concentrated insurance market. Nevertheless, we note that these results should be interpreted with caution because there remains a question as to whether the reporting of such a distinct designation for "experience-rated" plans is reliable in Form 5500.

Table A5: Evidence from Experience-Rated Plans

	Premium			Claim		
	(1)	(2)	(3)	(4)	(5)	(6)
	All	Mid-Sized	Large	All	Mid-Sized	Large
ln(HHI)	-0.164***	-0.0901**	-0.308**	-0.163***	-0.0679	-0.388***
	(0.0496)	(0.0450)	(0.121)	(0.0542)	(0.0523)	(0.133)
N	60,031	49,505	10,526	60,031	49,505	10,526

Note: This table presents evidence on the effect of insurer concentration on hospital/provider prices using a subsample of "experience-rated" plans. The first three columns study premiums and the next three study claims per enrollee, both measured in natural logarithm. Mid-sized firms are defined as those with fewer than 1,500 enrollees averaged over the sample period, and large firms have enrollment above this threshold. All regressions follow the main specification of Equation (2) and estimated using 2SLS. *** Significant at the 1% level. ** Significant at the 5% level. * Significant at the 10% level. Standard errors are shown in parentheses and clustered at the market-year level.

The second assumption of our model is that large firms have better access to self-insurance. This is because mid-sized firms might lack the resources for managing a self-insured plan. They might also consider self-insurance unappealing due to the difficulty of carrying the financial risks and dealing with the unstable cash flow caused by volatility in claims. As a result, self-insurance is more prevalent among large firms and we expect the likelihood of full-insurance to decrease with firm size. This is indeed what we observe in our data. For example, among firms with a total number of employee participants less than 1,000, the likelihood of full-insurance is 56% v.s. 28% for firms larger than 1,000. To get a better sense of the relationship between firm size and choosing full-insurance, in Figure A5, we categorize firms' employee participants based on bins of 200, and we present the average rate of full-insurance within each bin (conditional on year fixed effects). We also plot a quadratic fit. Figure A5 clearly suggests that large firms are more likely to opt for self-insurance.

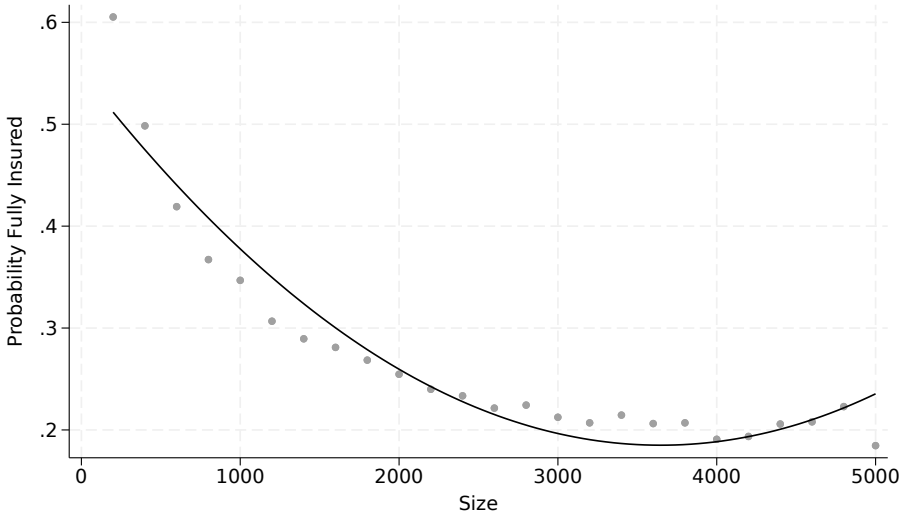


Figure A5: Probability of Full-Insurance by Firm Size

Note: This figure shows the relationship between firm size and the probability of full insurance, using a bin size of 200 and controlling for year fixed effects. The solid line plots a quadratic fit of the scatter plots. Here firm size is defined based on employee enrollment (instead of total enrollment used to define firm size in our main analysis) due to the lack of information on total enrollment (including family members) for self-insured firms.

A7.5 Caveats

We briefly discuss several caveats of our model. It is important to note that there are other plausible models/mechanisms that could rationalize the observed patterns in the data. We are, nevertheless, not well equipped to examine those alternatives. For example, our results might also be explained by a Bertrand model that allows for price discrimination by firm size due to differences in employers' elasticity of demand; however, whether demand is indeed more elastic for large employers desires empirical support. Another alternative is to incorporate employers' (insurers') underinvestment in health insurance provision due to labor (insurance) market turnover and frictions [Fang and Gavazza (2011); Aizawa and Fang (2020)]. Underinvestment could be less severe for large employers in a highly concentrated insurance market, as they tend to enjoy lower labor (and insurance) turnover and can better capture the full surplus from investing in an employee's health. Such dynamics, from the perspective of an employer and an insurer, could lead to better health and lower costs for large employers in more concentrated markets. Alas, we lack empirical support regarding labor turnover, and considering such dynamics is beyond the scope of this article. We also acknowledge that our model does not incorporate other important features of the ESI market, such as search frictions [Cebul et al. (2011)] and invest-harvest pricing dynamics [Craig (2022)].

A8 Data and Data Cleaning

The Employee Retirement Income Security Act of 1974 (ERISA) requires any administrator or sponsor of an employee benefit plan to report on the plan details unless exempt from filing. Plans with fewer than 100 participants are generally exempt, except when they operate a trust. Government and church plans are considered non-ERISA plans and therefore not included. If a plan is offered through an external insurance contract, a Schedule A must be attached, which offers details on the contract. If any assets of a plan are held in a trust, a Schedule H (for plans with 100 or more participants) or Schedule I (for plans with fewer than 100 participants) must be filed.

A8.1 Form 5500

Table A6: Cleaning Procedures for Sample_Form_5500

	Step	Observation	Firm	Employee
Start	1	4,955,282	160,850	120.2
Keep Health	2	3,899,840	131,540	74.0
Drop Incomplete Firm	3	3,899,723	131,531	74.0
Drop Incomplete Plan	4	3,871,022	131,408	74.0
Drop Duplicates	5	3,021,533	131,408	73.8
Non-Missing Participants	6	3,006,443	128,133	75.8
U.S.	7	2,999,812	127,668	75.7
Define Insurance Type	8	2,999,812	127,668	75.7
Define Insurer Group	9	2,965,744	124,188	75.5
County	10	2,963,155	123,421	75.5
Single Employer	11	2,839,683	118,534	67.5

Note: This table lists the data cleaning steps for Sample_Form_5500. *Observation* denotes the number of plan-level observations in the data; *Firm* refers to the number of firms identified using EIN; *Employee* is the total number of enrolled employees (excluding family members) measured for year 2017 (in millions). In Step 6 (Non-Missing Participants), we imputed observations with missing participants based on total active participants, detailed Schedule A enrollment information, or the median employee enrollment of the same plan over the years.

To explain the structure of our main dataset, consider an employer that offers health plans to its employees. We differentiate employers by their reported EIN number.[34] Schedules that are of particular interest to us are A, H, and I. As detailed in Table A6, we took several steps to clean up the data. We started by merging Form 5500 with Schedules A, H, and I, and Form 5500-SF.[35] Schedule A and Form 5500 contained variables needed for our main analysis, as they focus on fully-insured plans. Schedules H and I and Form 5500-SF helped us define other plan types, such as self-insured plans. We focused on plans with health benefits.[36] We then removed any plans that were missing the firm EIN or location information or that had an incomplete plan number. We also removed duplicate observations by only keeping the latest filing for each combination of benefits.[37] As firm size is key to our analysis, we excluded firms

[34] In rare cases (less than 0.37%), some EINs reported multiple observations for each of their operating states. In those cases, we defined an employer as the combination of EIN+state.

[35] Benefits other than welfare plans are pensions. Welfare plans utilize codes 501 and higher for entry 1b ("Three-digit plan number") on Form 5500. For the 2018 data, we used the version published as of January 7, 2020.

[36] Plans with "4A" in entry 8b ("Welfare feature codes") on Form 5500.

[37] Defined as the combination of firm + plan digit + zip code + insurer + type of benefit + year.

with zero or missing participants after imputing the firm size from total active participants (0.2%), the median employee enrollment of the same plan over the sample years (2.4%), or detailed Schedule A enrollment information (0.5%). We further restricted our sample to the 50 states of the United States plus Washington D.C.

Next, we defined plans as fully-insured, mixed-insured, or self-insured following the Annual Report to Congress on Self-Insured Group Health Plans prepared by the Department of Labor.[38] We also matched our sample using carrier EIN, NAIC code, and insurer name to the Listing of Companies from the National Association of Insurance Commissioners to reflect the ownership structure among the insurers.[39] Additionally, we removed any observations that could not be merged to a county because of erroneous address information. With the county FIPS code, we further merged to MSAs with the crosswalk provided by the Bureau of Economic Analysis. Finally, we focused on single-employer plans for a more straightforward relationship between firms and insurers. This treatment is due to the fact that we were not able to track which employers bargained as a group with insurers. It is important to note that single-employer plans accounted for 96% of the plans in our data.

All of the above steps led us to Sample_Form_5500.

A8.2 Fully-Insured Sample

Table A7: Cleaning Procedures for Sample_Fully_Insured

	Step	Observation	Firm	Employee
Start	1	2,839,683	118,534	67.5
Fully- & Mixed-Insured	2	2,118,461	86,414	36.7
Health	3	834,341	81,675	31.8
Construct Enrollment	4	834,341	81,675	31.8
Construct Premium	5	834,341	81,675	31.8
Winsorize Premium	6	723,990	79,538	30.2
Keep Large	7	703,437	72,659	29.2
Firm Panel	8	664,782	60,139	28.0

Note: This table lists the data cleaning steps for Sample_Fully_Insured. *Observation* denotes the number of plan-level observations in the data; *Firm* refers to the number of firms identified using EIN; *Employee* is the total number of enrolled employees (excluding family members) measured for year 2017 (in millions).

Once we got Sample_Form_5500, we kept fully- and mixed-insured plans that indicated health benefits on Schedule A.[40] We defined a plan as a combination of firm + plan digit + insurer + benefit.[41] We also corrected some reporting errors for enrollment.[42]

[38] "Mixed-insured" is defined as plans that indicate both fully- and self-insured components on Form 5500.

[39] The NAIC insurer group information is only available for years 2005 to 2018. For years before 2005, we used 2005 NAIC information.

[40] Meaning that they checked the box for 8a ("Health") on Schedule A. We also considered plans that failed to check "Health", but indicated "HMO, PPO, or Indemnity" without checking supplemental benefits as health plans.

[41] As listed in the checkboxes for Line 8 of Schedule A. For example, if a filer checks HMO, Dental, or Stop-Loss, then it is defined as a benefit. HMO+Vision is then another benefit type.

[42] For each plan, we marked an enrollment outlier if it was outside of the [25, 75] percentile of the enrollment's general distribution and it: (1) deviated from the plan's median enrollment (over the years) by 8 mean absolute deviation (MAD) values, or (2) had missing values, or (3) was over 5 times than the number of employee beneficiaries. Then we imputed the outliers from the plan's enrollment in the previous and following years; these imputed data accounted for 2% of our sample.

To determine premiums, two entries from Schedule A were used. For "experience-rated" contracts, we extracted data from Line 9a on Schedule A. For non-experience-rated contracts, we extracted data from Line 10. Note that "experience-rated" contracts here refer to plans that will get a refund of premiums if their incurred claims are lower than expected. Although these two lines provided premium information for most of the observations, we utilized the Form 5500 Group Health Plans Research File to correct premiums from other entries.[43] After cleaning up the premium variable, we defined premium per capita as total premium divided by total enrollment. We winsorized the premium per capita variable by removing any observations that were out of the $[10, 99]$ percentile of its distribution.[44]

Note that we only keep large plans that had (at least once during the sample period) over 100 participants (i.e., we removed all plans that were voluntarily filing Form 5500). Lastly, as most of our analysis focused on within-firm variation, we restricted our sample to firms that showed up for at least two consecutive years in our data.

A8.3 Analysis Sample

Table A8: Cleaning Procedures for Sample_Plan

	Step	Observation	Firm	Employee
Start	1	664,782	60,139	28.0
After 2001	2	625,705	59,422	28.0
Aggregate Plan	3	602,589	59,424	28.0
Large MSA	4	533,537	50,701	26.1
Fully-Insured	5	443,665	49,529	14.0
Non-Experience-Rated	6	372,112	45,474	12.4
Lag Year	7	328,774	38,356	11.7

Note: This table lists the data cleaning steps for Sample_Plan (our analysis sample). *Observation* denotes the number of plan-level observations in the data; *Firm* refers to the number of firms identified using EIN; *Employee* is the total number of enrolled employees (excluding family members) measured for year 2017 (in millions).

Noticing that the sample size for year 2000 is small, we extracted data starting in 2001. To create our final analytic sample at the firm-plan level, we aggregated the previously defined detailed "plans" into a unique combination of [firm + insurer + plan type], in which plan type included HMO, PPO, Indemnity, and Others. We called this combination "aggregated plans". Correspondingly, drug, dental, and vision coverage are defined, respectively, as the share of enrollees for each particular benefit over the total number of enrollees for each aggregated plan. Premium for the aggregated plan was then calculated as the weighted average of these detailed plans by enrollment.

In order to have an accurate measure of insurer market competition, we restricted our sample to MSAs with 20 or more firms. Although our previous data cleaning steps utilized Schedule A plans from both fully- and mixed-insured firms, we focused in our main analysis purely on fully-insured firms. In addition, among the fully-insured plans, around 16% were "experience-rated", meaning that their

[43] We then marked an outlier if it deviated from the plan's median (over the study years) by a value of 8 MADs and was outside of the [25, 75] percentile of the premium's general distribution in our data. Next, we imputed outliers from the plan's premium in the previous and following years; these imputed data accounted for 2% of our sample.

[44] We selected this criterion out of heuristics so that premiums for an enrollee ranged from around \$552 to \$18,642.

premiums were more likely to be determined by their own incurred claims. We then removed these plans from our main analysis but used them separately for a supplemental analysis. Finally, because our empirical model specified that insurer HHI affects premiums in the following year, we removed firms with missing $t-1$ information. We defined our sample up to this step as Sample_Plan, and it contained $328,774$ observations covering $38,356$ distinct firm EINs.