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RESEARCH ARTICLE

Hospital and Health Insurance Markets Concentration and Inpatient Hospital Transaction Prices in the U.S. Health Care Market

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Objective. To examine the effects of hospital and insurer markets concentration on transaction prices for inpatient hospital services.

Data Sources. Measures of hospital and insurer markets concentration derived from American Hospital Association and HealthLeaders-InterStudy data are linked to 2005–2008 inpatient administrative data from Truven Health MarketScan Databases.

Study Design. Uses a reduced-form price equation, controlling for cost and demand shifters and accounting for possible endogeneity of market concentration using instrumental variables (IV) technique.

Principal Findings. The findings suggest that greater hospital concentration raises prices, whereas greater insurer concentration depresses prices. A hypothetical merger between two of five equally sized hospitals is estimated to increase hospital prices by about 9 percent ($p < .001$). A similar merger of insurers would depress prices by about 15.3 percent ($p < .001$). Over the 2003–2008 periods, the estimates imply that hospital consolidation likely raised prices by about 2.6 percent, while insurer consolidation depressed prices by about 10.8 percent. Additional analysis using longer panel data and applying hospital fixed effects confirms the impact of hospital concentration on prices.

Conclusion. The findings provide support for strong antitrust enforcement to curb rising hospital service prices and health care costs.

Key Words. Competition, health care market, transaction prices, bilateral effects, instrumental variables

The rising cost of health care is a major concern in the United States. Health care costs as a percentage of GDP has grown from 13.8 percent in 2000 to 17.4 percent in 2013 (CMS 2013). A central contributor to this cost growth is the consolidation of insurance and provider markets. Indeed, the U.S. health care market has experienced substantial consolidation in recent years, and many researchers have cited the resulting increase in provider and insurer markets concentration as a key contributor to the growth in provider prices, insurance premiums, and overall health care costs (e.g., Cuellar and Gertler 2003; Dafny 2009; Capps 2010).

The impact of consolidation on hospital markets is a particular area of concern, as these markets have undergone substantial consolidation. From 1990 to 2003, the average metropolitan statistical area (MSA) resident saw a decline in the number of competing local hospital systems from six to four (Vogt and Town 2006). Furthermore, consolidation has continued from 2003 onwards, although at a slower pace. Moreover, the passage of the Affordable Care Act has reignited concerns over consolidation in provider markets (Berenson, Ginsburg, and Kemper 2010; Robinson 2011). Insurer concentration has also increased in recent years. Data from Gaynor and Town (2012) showed that, between 1998 and 2009, the average market-level Herfindahl-Hirschman Index (HHI) for the large group insurance market rose from 2,172 to 2,956, equivalent to a reduction in the number of equal-sized insurers from roughly 5 to less than 4.

While some researchers and health insurance groups are concerned that rising hospital concentration is putting an upward pressure on hospitals prices, provider groups are also worried that rising insurer concentration is depressing hospital prices. Hitherto, the dearth of data on actual payments to hospitals and comprehensive and reliable data on health insurance markets has hindered research into the bilateral price effects of hospital and insurer concentrations in the U.S. health care market (Dafny et al. 2011). As such, much of the prior studies focused on the unilateral price effect of hospital concentration, ignoring the fact that it is the balance of market power between providers and insurers in negotiations determines hospital prices. Studies of hospital competition are also limited along other dimensions. Specifically, these studies are often missing data on actual hospital prices or focus only on the effects of concentration in particular geographic markets or medical conditions (see Gaynor and Vogt 2000; and Gaynor and Town 2012 for detail reviews of the literature).

This study examines the bilateral effects of hospital and insurer markets concentration on negotiated prices for inpatient hospital services in a cross section of hospitals and the unilateral price effect of hospital market concentration in a panel framework. The study uses a large U.S. data set of the

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commercially insured from 2005 to 2008 and an instrumental variables strategy to identify the price effect of market concentration.

The study improves on prior bilateral studies examining the price effects of hospital and insurer markets concentration using U.S. data in an important way. It uses a large and comprehensive data set with actual transaction prices from a more recent period (i.e., 2005–2008) and simultaneously addresses potential endogeneity and measurement concerns associated with measures of market concentration in reduced-form empirical studies using an IV strategy.

The IV results indicate that higher hospital (insurer) concentration significantly raises (lowers) inpatient transaction prices. The estimates suggest that mean increases in hospital concentration between 2003 and 2008 likely raised hospital prices by about 2.6 percent by 2008 (about $4.9 billion in annual hospital expenditures incurred by private payers from their 2003 baseline amount of around $192 billion), ceteris paribus. Simultaneously, mean increases in insurer concentration between 2003 and 2008 possibly depressed hospital prices by about up to 10.8 percent (about $20.7 billion in reduced spending on inpatient hospital services). In addition, I find evidence supporting the argument that the relevant geographic market for hospitals may be far smaller than what the literature, the courts and policy circles generally consider: It may be as low as a 10-minute drive time surrounding the centroid of a census tract.

A limitation of prior bilateral studies is that they are cross-sectional. However, this study also estimates a hospital fixed effects panel regression model as an alternative to the cross-sectional estimates. The panel analysis, which relies on the within-hospital changes in prices and concentration over time, improves identification of the price effect by accounting for hospital-specific quality and avoiding potential biases arising from unmeasured differences in care quality across hospitals. For panel data analysis, studying a longer period is necessary to observe sufficient changes in consolidation to identify the price effect of consolidation. Unfortunately, a lack of accurate insurer concentration measures over the period 2003–2005 limits the panel analysis to the examination of only the effect of hospital concentration.

The panel estimates confirm the positive association between rising hospital market concentration and price growth. The IV point estimates show that a 10 percent increase in hospital concentration raises prices by 0.8–3.1 percent, while a consolidation from five equally sized hospitals to four (i.e., “5 to 4” merger) would cause prices to increase by 3–11 percent.

Related Literature

Much of the literature on hospital competition has focused on the unilateral price effect of competition because of the availability of data on hospital markets. The consensus from this body of work is in line with economic theory that lower competition leads to higher prices (e.g., Capps and Dranove 2004; Dranove et al. 2008; Robinson 2011). Relatively few studies have examined the price effect of insurer market concentration because of data limitations (Dafny et al. 2011), but much of prior studies find a negative association between insurer concentration and prices (Moriya, Vogt, and Gaynor 2010).

This paper relates to reduced-form studies that have examined the bilateral price effects of hospital and insurer markets concentration (e.g., Staten, Dunkelberg, and Umbeck 1987, 1988; Melnick et al. 1992; Sorensen 2003; Moriya, Vogt, and Gaynor 2010; Halbersma et al. 2011; Melnick, Shen, and Wu 2011). Within this literature, two recent papers closely relate to this study: Moriya, Vogt, and Gaynor (2010) and Melnick, Shen, and Wu (2011). The first paper, Moriya, Vogt, and Gaynor (2010), examined the correlations between hospital and insurer markets concentration and hospital prices using 2001–2003 inpatient admission data with actual transaction prices, the same data source employed in this study. Their geographic market for hospitals was the Health Service Areas and that for insurers was the state. Their findings showed that higher insurer concentration significantly reduces hospital prices, but higher hospital concentration had an insignificant effect on prices.

The second paper, Melnick, Shen, and Wu (2011), used 2001–2004 data from Medicare hospital cost reports to examine the link between hospital and insurer concentration measures and prices. Unlike Moriya, Vogt, and Gaynor (2010), they considered the local nature of competition by calculating a hospital-specific concentration index for each hospital market and measured insurer competition at the MSA level. Their results show that higher insurer concentration significantly reduces average hospital prices, with much stronger effect in the most concentrated insurer markets. However, contrary to Moriya, Vogt, and Gaynor (2010), they found that higher hospital concentration significantly raises prices.

This study is, however, distinct from these two papers in a few important ways. Unlike Moriya, Vogt, and Gaynor (2010), this study uses an index of hospital concentration that accounts for the local nature of competition in the hospital industry and an index of insurer concentration at the MSA level. The literature on hospital competition has thoroughly documented that travel time of patients to hospitals is a key determinant of a patients' hospital choice. It is therefore essential to consider the local geographic market when measuring hospital market concentration (Kessler and McClellan 2000; Gaynor and Vogt 2003). Not doing so may inaccurately characterize the true effect of hospital competition. Similarly, Moriya, Vogt, and Gaynor (2010) measured insurer competition at the state level, although evidence suggests that competition occurs locally at the MSA level (Pauly et al. 2002; Kopit 2004; Robinson 2004).

While Melnick, Shen, and Wu (2011) considered local geographic markets when constructing their concentration measures, they did not use transaction prices. Rather, they used average prices as a proxy, which they constructed by regressing total net revenue on adjusted patient days (as proxy for hospital volume) and controlling for hospital case-mix and other characteristics. However, a hospital's average price is a vague term as there are many other measures of volume (e.g., admissions, discharges). This may potentially cause substantial measurement error in average prices (Gaynor and Vogt 2000). It is therefore not clear whether a hospital's average price provides accurate information about actual prices received from payers. This study rather uses actual transaction prices for inpatient hospital services negotiated between hospitals and insurers.

In addition to these differences, this study differs from these two papers by addressing potential endogeneity and measurement error concerns associated with measures of market concentration in reduced-form empirical studies. Their failure to address these concerns is potentially problematic. There is reason to believe that some unobservable factors that are not exogenous to prices are also likely to affect differences in hospital and insurer concentrations. For instance, markets with many individuals in high demand for medical care may have high prices because such markets may have more severe cases. However, such markets may attract more providers, leading to lower hospital concentration. Ignoring such a concern may potentially cause the estimated effect of hospital concentration to be downward-biased. This study mitigates the endogeneity concern using an IV technique.

Finally, this study differs from these papers in that it uses panel data analysis to improve identification and thereby contributing to the unilateral effects literature.

Methods

Data Sources

This study uses 2005–2008 data from multiple sources. The main data set is the Inpatient Services files of Truven Health MarketScan Commercial Claims and Encounters Database. The files contain data on more than 6.5 million unique privately insured patients with over 11 million inpatient visits. An observation is an inpatient admission. The measures of hospital and insurer markets concentration are from the American Hospital Association's (AHA) Annual Survey of Hospitals Database and the HealthLeaders-InterStudy, respectively. Besides market-level covariates from the AHA data set, other covariates and instrumental variables come from the Bureau of Health Professions' Area Resource File and the County Business Patterns. A measure of the stringency of certificate-of-need (CON) regulations, which require certain health care providers to seek state government's approval before starting certain large capital projects (e.g., constructing new hospitals), is from the American Health Planning Association.

Dependent Variable

The dependent variable is the logarithm of total gross facility payments to hospitals for an inpatient admission. Gross facility payments—including deductibles and co-payments, but excluding professional-related payments—are what hospitals receive for claims after applying pricing guidelines. It constitutes over 98 percent of paid claims in the MarketScan inpatient data. The analysis excludes nonpositive payments and capitated claims (because payment information is unobserved), and also the top and bottom 1 percent of payments in each DRG-year to reduce the effect of coding errors.

Independent Variables

The key independent variables are the logarithms of hospital and insurer markets concentration. To calculate these variables, a definition of the relevant product and geographic market is necessary and important.

I define the hospital product as "general acute care hospital services" as is typical in the literature for this kind of analysis (Gaynor and Vogt 2000). I use data only on short-term general medical and surgical hospitals that operated for at least 180 days in each survey year, which is over 99 percent of hospitals. The definition of the relevant geographic market for hospitals is an issue of ongoing debate. I define the hospital market as a specific amount of travel time boundary surrounding a census-tract centroid. I use market boundaries ranging from 10 to 80 minutes.

The construction of the measure of hospital concentration follows the methodology of Dunn and Shapiro (2014), which examines physician market concentration. Following these authors, I construct an index of hospital market concentration, termed the "Fixed-Travel-Time Herfindahl-Hirschman Index" (FTHHI), which takes into account the distance and travel time of patients to competing hospitals in the market boundary. The following is a summary of the methodology, details of which are in Dunn and Shapiro (2014).

First, I define the market boundary as fixed distance surrounding the centroid of a census tract, assuming that all patients living in a particular census tract reside at the centroid of the census tract. Then, I calculate the probability of a hospital located at a certain place attracting a patient located at another place, based only on information about the distances between hospitals and patients. Next, I calculate an FTHHI for a census-tract centroid as the sum of squared expected market shares for each hospital system, using the probabilities and the number of staffed beds for a hospital system. Hospital systems that are closer to a patient receive more weight. I treat hospitals belonging to the same system as one, and assign the same market share to them because they are more likely to negotiate collectively with health plans. Finally, I calculate a county-level FTHHI as the weighted sum of the census-tract-level FTHHI, using census-tracts' population share as weights. I match the county-level FTHHI measures to patients in the MarketScan data set who are residents of the county. In effect, the geographic market area becomes a census-tract area surrounding a particular patient.

It is important to note that the measure of hospital concentration in this study does not use the actual hospital shares, but instead a predicted concentration measure based on the distance and travel time of patients to competing hospitals in a market boundary. Incorporating geographic factors is essential for properly measuring competition in hospital markets, and using predicted measures reduces the possibility of endogeneity bias (Kessler and McClellan 2000; Town and Vistnes 2001; Gaynor and Vogt 2003).

For insurers, while some studies have used the state as the relevant geographic market, others recognize the MSA as the relevant market, arguing that the health insurance marketplace is local because employers and consumers value proximity of provider networks to their workplaces or residences (Pauly et al. 2002; Kopit 2004). This study uses the MSA as the relevant market for health insurers. Regarding the relevant product, like Melnick, Shen, and Wu (2011), this study takes the view that health maintenance and preferred provider organization plans are close enough, at least in the eyes of many consumers, for one to substitute for the other, and so calculates a single insurer HHI by combining total enrollment for both plans. The HHI of insurer concentration is calculated as the sum of squared market shares for each insurer operating in a market. The calculation accounts for merger activities that occurred during the sample period and treats all Blue Cross and Blue Shield subsidiaries in markets where they do not compete as one insurer.

Covariates

The covariates include individual-, hospital-, and market-level variables that might affect geographic market demand and supply. Individual-level covariates include gender, age, health plan type, number of diagnoses and procedures, and length of stay, Charlson comorbidity index, indicators for employer-based health plans, and for 50 different condition categories.[1]

Hospital- and market-level variables include average hospital length of stay, median household income and the fraction of college-educated people in patients' county. Others include median gross rent, median household income, fraction of teaching, and medical-school-affiliated hospitals and fraction of various hospital-physician arrangement types in hospitals' county. The analysis includes state-year fixed effects to mitigate potential biases from any unobserved state heterogeneity that may correlate with markets concentration and prices. Table S1 in Appendix SA1 shows the variables, their specification, and data sources.

Statistical Methods

The study uses both OLS and IV regression models to examine the price effects of hospital and insurer concentrations. Conditional on the observed covariates, the prediction is that rising hospital concentration should lead to higher prices and rising insurer concentration should lead to lower prices paid by insurers.

Instruments Variables and Identification

Despite the careful measurement of market concentration and the inclusion of the many covariates, OLS estimates are unlikely to be the true effect of market concentration because of potential endogeneity and measurement error concerns associated with market concentration measures, especially insurer concentration. I mitigate these concerns using an IV approach.

This paper identifies the price effect of market concentration using a new identification strategy: The across-state differences in the stringency of

CON regulations interacted with proxies for market size. Specifically, it uses the patient's county population aged 65 and over, number of business establishments with at least 100 employees in the hospital's MSA, and their interaction terms with a 2-year lag of a categorical variable measuring the comprehensiveness of states' CON regulations as the identifying instruments. The categorical variable contains values of 0, 1, and 2 for states with no CON regulations, states whose weighted rank of CON-regulated services is less than 10, and states whose rank is at least 10, respectively.

These instruments are assumed to be uncorrelated with prices but should uniquely explain the variations in the hospital and insurer markets concentration. For instance, providers are more likely to enter geographic areas with large elderly population, whereas insurers are less likely to be affected by this population because the majority of them are enrolled in Medicare. Similarly, insurers are more likely to enter markets with a higher number of large employers because the majority of the privately insured population obtains their health insurance through the workplace. The CON regulations act as a fixed cost that hospitals incur when entering CON-regulated markets. By creating barriers to market entry, expansion, and competition, they limit the supply of health services. Indeed, some empirical evidence suggest that CON regulations have deter market entry and capacity expansion (see Ford and Kaserman 1993; Caudill, Ford, and Kaserman 1995; Ho 2009). Besides, prior studies like Abraham, Gaynor, and Vogt (2007) have used CON regulations as instruments.

As a check detected the presence of arbitrary heteroscedasticity in the second-stage IV regression, I use a heteroscedasticity-robust estimator, the two-step feasible efficient generalized method of moments, instead of the traditional two-stage least squares estimator.

RESULTS

Descriptive Statistics

Table 1 presents selected summary statistics for the final sample consisting of 4,884,854 observations on 3,554,713 unique patients for the study period. The table reports measures of hospital market concentration for various market boundaries. During the study period, hospital payments rose by about 17.6 percent. Insurer concentration rose by about 6.1 percent, while hospital concentration decreased by up to 8.8 percent. The sample has a high proportion of females, and almost two-thirds of the people had coverage under PPO

Table 1: Selected Descriptive Statistics

Variables	*2005*	*2006*	*2007*	*2008*	*Overall*
Hospital payments ($)	10,310	10,442	11,245	12,122	11,106
	(20,285)	(20,510)	(22,543)	(23,893)	(22,031)
Insurer HHI†	0.2809	0.2850	0.3196	0.2994	0.2979
	(0.1099)	(0.1032)	(0.1278)	(0.1225)	(0.1179)
Hospital FTHHI‡					
10-minute boundary	0.6638	0.6565	0.6505	0.6564	0.6560
	(0.1917)	(0.1765)	(0.1776)	(0.1682)	(0.1771)
20-minute boundary	0.3850	0.3626	0.3541	0.3568	0.3622
	(0.1950)	(0.1895)	(0.1887)	(0.1895)	(0.1905)
30-minute boundary	0.2686	0.2473	0.2405	0.2448	0.2481
	(0.1631)	(0.1640)	(0.1610)	(0.1638)	(0.1632)
40-minute boundary	0.2132	0.1949	0.1906	0.1944	0.1966
	(0.1369)	(0.1408)	(0.1381)	(0.1415)	(0.1398)
50-minute boundary	0.1815	0.1664	0.1641	0.1676	0.1686
	(0.1192)	(0.1249)	(0.1222)	(0.1259)	(0.1237)
60-minute boundary	0.1607	0.1485	0.1476	0.1507	0.1509
	(0.1079)	(0.1141)	(0.1120)	(0.1160)	(0.1132)
70-minute boundary	0.1467	0.1361	0.1360	0.1391	0.1387
	(0.0997)	(0.1059)	(0.1041)	(0.1092)	(0.1055)
80-minute boundary	0.1365	0.1271	0.1272	0.1305	0.1297
	(0.0930)	(0.0990)	(0.0974)	(0.1038)	(0.0991)
Patient characteristics					
Age	38.1	36.9	36.8	37.1	37.2
	(19.6)	(19.8)	(20.0)	(19.8)	(19.8)
% Male	37.9	38.0	37.8	38.1	37.9
	(48.5)	(48.5)	(48.5)	(48.6)	(48.5)
No. of inpatient days	3.9	3.9	3.9	3.9	3.9
	(6.2)	(5.9)	(6.0)	(5.9)	(6.0)
No. of diagnoses	5.7	5.9	6.3	6.4	6.1
	(3.5)	(3.8)	(3.8)	(3.9)	(3.8)
Plan types					
% COMP	10.3	6.4	3.3	3.0	5.3
% EPO	0.5	0.5	0.7	0.8	0.6
% HMO	16.4	13.3	13.2	13.9	14.0
% POS	13.9	12.0	11.7	11.5	12.1
% PPO	57.0	65.5	68.8	67.7	65.5
% CDHP	1.9	2.3	2.3	2.5	2.3
% HDHP	0.0	0.0	0.0	0.6	0.2
No. of observations	904,835	1,279,614	1,336,967	1,363,438	4,884,854

Notes: Mean values with standard deviations in parentheses. Hospital payments are in nominal dollars. Plan types: COMP is Comprehensive, EPO is Exclusive Provider Organization, HMO is Health Maintenance Organization, POS is Noncapitated Point-of-Service, PPO is Preferred Provider Organization, CDHP is Consumer-Driven Health Plan, and HDHP is High Deductible Health Plan.

†The variable is weighted by the MSA population.

‡The variable is weighted by the county population.

plans. Although not representative of the entire U.S. population, the MarketScan data compares well with the U.S. population with employer-sponsored coverage. For example, hospital transaction prices in the data are broadly consistent with those from two other data sources: The Health Care Cost Institute and the states of California and Oregon (Lemieux and Mulligan 2013).

Cross-Sectional Regression Results

Table 2 presents both OLS and IV results for hospital and insurer concentrations using hospital market boundaries definitions ranging from 10-minute to 80-minute drive time (Table S2 in Appendix SA1 reports the full set of results). Unsurprisingly, the OLS estimates show a significant positive correlation between hospital concentration and prices, but the significant positive correlation between insurer concentration and prices is surprising. However, these associations are unlikely to be causal because of endogeneity issues. The IV results address endogeneity concerns.

However, before discussing the IV results, Table S3 (Appendix SA1) presents results from the first-stage and reduced-form regressions assessing the effects of the instruments on market concentration and prices. As expected, higher population of seniors and higher number of large establishments lead to lower hospital and insurer markets concentration but these variables have less or no effect when interacted with the CON variable. The second-stage regression diagnostics jointly assessing the validity and relevance of the instruments, reported in the bottom part of Table 2, suggest that underidentification is not a concern, while weak identification is only of a moderate concern. Indeed, in most boundaries, the weak identification test suggests that the instruments are strong enough for instrumentation to remove a sizable portion of the OLS bias. Furthermore, the validity of the excluded instruments is not in doubt, as shown by the overidentification test, while the endogeneity test indicates that IV estimation, rather than OLS, is more appropriate.

Now turning to the IV results in Table 2, the estimates show that endogeneity problems severely bias the price effect of market concentration. Indeed, OLS estimation biases the estimates for hospital concentration downwards and those for insurer concentration upwards. The hospital concentration estimates are still positive and statistically significant, but they decrease with the size of the market boundary until the 30-minute boundary. A 10 percent increase in hospital concentration raises prices by 1.4 to 5 percent. A hypothetical “5 to 4” hospital merger would cause

Table 2: The Price Effect of Hospital and Health Insurer Markets Concentration

	Hospital Market Boundaries: Driving-Time Radius							
	10-minute		*20-minute*		*30-minute*		*40-minute*	
Variables	*OLS*	*IV*	*OLS*	*IV*	*OLS*	*IV*	*OLS*	*IV*
Hospital FTHHI	0.169***	0.502***	0.116***	0.177***	0.094***	0.142***	0.090***	0.146***
	(0.028)	(0.049)	(0.014)	(0.017)	(0.011)	(0.014)	(0.011)	(0.015)
Health insurer HHI	0.060***	−0.113	0.056***	−0.117	0.057***	−0.128	0.057***	−0.185
	(0.020)	(0.175)	(0.019)	(0.170)	(0.019)	(0.168)	(0.019)	(0.168)
No. of observations	4,884,854	4,884,854	4,884,854	4,884,854	4,884,854	4,884,854	4,884,854	4,884,854
R-squared	0.598	0.397	0.599	0.402	0.599	0.402	0.599	0.400
IV Regression Diagnostics								
Under identification test								
Kleibergen-Paap rk LM	—	21.850	—	25.073	—	24.998	—	23.481
	—	[0.000]	—	[0.000]	—	[0.000]	—	[0.000]
Weak identification test								
Kleibergen-Paap rk Wald F	—	8.475	—	9.201	—	9.203	—	8.790
Over identification test								
Hansen J statistic	—	0.102	—	1.203	—	1.288	—	1.363
	—	[0.950]	—	[0.548]	—	[0.525]	—	[0.506]
Endogeneity test	—	33.615	—	20.196	—	18.850	—	20.711
	—	[0.000]	—	[0.000]	—	[0.000]	—	[0.000]

Continued

Table 2. *Continued*

	Hospital Market Boundaries: Driving-Time Radius							
	50-minute		*60-minute*		*70-minute*		*80-minute*	
Variables	*OLS*	*IV*	*OLS*	*IV*	*OLS*	*IV*	*OLS*	*IV*
Hospital FTHHI	0.089***	0.166***	0.086***	0.191***	0.080***	0.222***	0.072***	0.256***
	(0.012)	(0.018)	(0.013)	(0.023)	(0.015)	(0.030)	(0.016)	(0.040)
Health insurer HHI	0.059***	−0.260	0.060***	−0.346*	0.061***	−0.425**	0.062***	−0.494**
	(0.019)	(0.173)	(0.020)	(0.188)	(0.020)	(0.209)	(0.020)	(0.238)
No. of observations	4,884,854	4,884,854	4,884,854	4,884,854	4,884,854	4,884,854	4,884,854	4,884,854
R-squared	0.599	0.397	0.598	0.392	0.598	0.385	0.598	0.377
IV Regression Diagnostics								
Under identification test								
Kleibergen-Paap rk LM	—	21.383	—	19.132	—	17.045	—	15.172
	—	[0.000]	—	[0.000]	—	[0.001]	—	[0.002]
Weak identification test								
Kleibergen-Paap rk Wald F	—	8.059	—	7.188	—	6.370	—	5.613
Over identification test								
Hansen J statistic	—	1.785	—	2.063	—	2.024	—	1.795
	—	[0.410]	—	[0.357]	—	[0.363]	—	[0.408]
Endogeneity test	—	23.818	—	27.562	—	31.479	—	35.137
	—	[0.000]	—	[0.000]	—	[0.000]	—	[0.000]

Notes: The dependent variable is the logarithm of total gross facility payments for a hospital admission. The instruments, all in logarithms, are *the county population aged 65 and over*, *the number of establishments with at least 100 employees in the MSA*, and their interaction terms with *a 2-year lag of a categorical variable measuring the comprehensiveness of states' CON regulations*. All specifications include patient-specific and county-level patient controls, hospital-specific and county-level hospital cost and quality controls, and fixed effects for condition category and state-year. Robust standard errors (clustered by county) are in parentheses. ***, **, and * indicate significance at 1 percent, 5 percent, and 10 percent. P-values for the IV regression diagnostics are in square brackets. For two endogenous variables and four instruments, the Stock and Yogo (2005) critical values for 5%, 10%, and 20% maximal IV relative bias are 11.04, 7.56, and 5.57. The critical values for 10%, 15%, 20%, and 25% maximal IV size are 16.87, 9.93, 7.54, and 6.28.

prices to increase by 4.9 to 18.4 percent. Putting these figures into perspective, Vogt and Town (2006)'s review of prior studies found that a "5 to 4" merger in the 1990s would have caused a 5 percent price increase. This is roughly the predicted effect at the 40-minute boundary in this study for such a merger.

For insurer concentration, the estimates are now of the expected sign, albeit they are statistically insignificant at conventional levels, except at the 60- to 80-minute boundaries, where a 10 percent increase in insurer concentration decreases prices by 3.5 and 4.9 percent. At the 70-minute and 80-minute boundaries, the estimates imply a price reduction in between 13.3 and 15.3 percent for a "5 to 4" insurer merger. The merger effect predicted by Moriya, Vogt, and Gaynor (2010) for such a merger was 6.7 percent, which is closer to the prediction at the 50-minute boundary in this study.

Overall, the findings in this study support prior studies examining such bilateral effects (e.g., Melnick et al. 1992; Halbersma et al. 2011; Melnick, Shen, and Wu 2011), but the significantly positive price effect of hospital concentration contradicts the finding in Moriya, Vogt, and Gaynor (2010).

Implications for Hospital Geographic Market Boundaries

An important finding for antitrust cases that is worth mentioning is about the relevant market boundary for hospital competition. While the 1990s saw the courts adopting a much broader geographic market boundary for hospital competition, in more recent litigated hospital merger cases the courts have accepted a relatively narrower definition. In a recent ruling to a merger challenge (i.e., ProMedica Health System, Inc. v. FTC), the 6th Circuit Court of Appeals agreed with the FTC contention that the relevant hospital market in that case was Lucas County, Ohio. The evidence in this study highlights such a difficulty in identifying the relevant market. While larger boundaries produce jointly significant coefficients on the concentration measures and thus provide a better description of the marketplace (i.e., the bilateral possession of market power), F-ratio tests suggest that the specification with a 20-minute travel-time boundary fits the data better. That there is a significantly positive price effect for the hospital concentration measure even at the 10-minute boundary suggests that the relevant hospital market boundary may be much smaller than the boundaries typically considered by the courts. This finding is consistent with those of Capps (2010) and Dafny (2009) for hospital markets, and Dunn and Shapiro (2014) for physician markets.

Measuring the Impact of Consolidation on Hospital Expenditures over the 2003–2008 Period

Based on the jointly significant point estimates on the concentration measures, the findings suggest that average increases in hospital market concentration between 2003 and 2008 likely raised hospital prices by between 1.9 to 2.6 percent by 2008, while mean increases in insurer market concentration possibly depressed hospital prices by between 7.6 and 10.8 percent, ceteris paribus. Using National Health Expenditure data from CMS, the effect of hospital consolidation at the 80-minute boundary translates into a cumulative estimated increase of roughly $4.9 billion in annual hospital expenditures by 2008 from the 2003 baseline amount of roughly $192 billion. Similarly, the effect of insurer consolidation translates into a decrease of about $20.7 billion in hospital expenditures. It is therefore likely that the net cumulative estimated effect of mean increases in both hospital and insurer markets concentration is a reduction in hospital prices. Indeed, upstream suppliers like providers are not the only ones to feel the impact of rising insurer consolidation, as downstream buyers may also likely be impacted. In fact, Dafny, Duggan, and Ramanarayanan (2012) find that insurer consolidation between 1998 and 2006 possibly raised annual employer health insurance premiums by an additional $34 billion.

Robustness Checks and Extensions

I conduct several checks to test the robustness of the results. First, I check the sensitivity of the results to a different instrument set, using the following as instruments: the total population in the patient's county, the number of establishments in the hospital's MSA, and their interaction terms with the CON variable. As shown in Table 3, the key results remain qualitatively similar. Second, most provider–insurer contracts are set in the prior year and so prices may respond with a lag to market concentration. I therefore use a specification that assumes a one-year lagged relationship between market structure and price. Table 4 shows both the OLS and IV results, which are quite similar to the main results.

Third, I use an alternative construction of FTHHI, which assumes no travel cost, and so assigns equal weights to hospitals located within the same market boundary regardless of the location of their patients. Indeed, this will be the correct weight if the distribution of patients within the market boundary is even. Table S4 (Appendix SA1) shows that the key results appear reasonably

Table 3: The Price Effect of Hospital and Health Insurer Markets Concentration Using an Alternative Set of Instruments

	Hospital Market Boundaries: Driving-Time Radius							
Variables	*10-minute*	*20-minute*	*30-minute*	*40-minute*	*50-minute*	*60-minute*	*70-minute*	*80-minute*
Hospital FTHHI	0.462***	0.169***	0.137***	0.143***	0.161***	0.186***	0.216***	0.248***
	(0.043)	(0.018)	(0.015)	(0.015)	(0.018)	(0.022)	(0.029)	(0.039)
Health insurer HHI	−0.135	−0.133	−0.145	−0.203	−0.270*	−0.347**	−0.417**	−0.479**
	(0.174)	(0.177)	(0.174)	(0.173)	(0.177)	(0.190)	(0.210)	(0.239)
No. of observations	4,884,854	4,884,854	4,884,854	4,884,854	4,884,854	4,884,854	4,884,854	4,884,854
R-squared	0.396	0.402	0.402	0.400	0.396	0.391	0.385	0.377
IV Regression Diagnostics								
Under identification test								
Kleibergen-Paap rk LM	23.585	26.099	25.962	24.467	22.471	20.211	18.029	15.985
	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.001]
Weak identification test								
Kleibergen-Paap rk Wald F	8.549	9.120	9.139	8.784	8.157	7.348	6.555	5.787
Over identification test								
Hansen J statistic	1.788	3.393	2.951	2.662	2.703	2.713	2.500	2.199
	[0.409]	[0.183]	[0.229]	[0.264]	[0.259]	[0.258]	[0.286]	[0.333]
Endogeneity test	32.027	17.569	16.533	18.767	22.441	26.899	31.207	34.727
	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]

Notes: The dependent variable is the logarithm of total gross payments to hospitals. The instruments, all in logarithms, are *the total population in the patient's county, the number of establishments in the hospital's MSA* and their interaction terms with *a 2-year lag of a categorical variable measuring the comprehensiveness of states' CON regulations*. All specifications include patient-specific and county-level patient controls, hospital-specific and county-level hospital cost and quality controls, and fixed effects for condition category and state-year. Robust standard errors (clustered by county) are in parentheses. ***, **, and * indicate significance at 1 percent, 5 percent, and 10 percent. *p*-Values for the IV regression diagnostics are in square brackets. For two endogenous variables and eight instruments, the Stock and Yogo (2005) critical values for 5%, 10%, and 20% maximal IV relative bias are 17.70, 10.22, and 6.20. The critical values for 10%, 15%, 20%, and 25% maximal IV size are 25.64, 14.31, 10.41, and 8.39.

Table 4: The Price Effect of One-Year Lagged Values of Hospital and Health Insurer Markets Concentration

	Hospital Market Boundaries: Driving-Time Radius							
	10-minute		*20-minute*		*30-minute*		*40-minute*	
Variables	*OLS*	*IV*	*OLS*	*IV*	*OLS*	*IV*	*OLS*	*IV*
Lagged hospital FTHHI	0.169***	0.455***	0.115***	0.159***	0.093***	0.128***	0.089***	0.135***
	(0.029)	(0.045)	(0.015)	(0.018)	(0.012)	(0.015)	(0.012)	(0.016)
Lagged health insurer HHI	0.039**	−0.141	0.036**	−0.077	0.037**	−0.097	0.038**	−0.168
	(0.018)	(0.223)	(0.018)	(0.213)	(0.018)	(0.212)	(0.018)	(0.214)
No. of observations	3,963,234	3,963,234	3,963,234	3,963,234	3,963,234	3,963,234	3,963,234	3,963,234
R-squared	0.608	0.402	0.608	0.408	0.608	0.408	0.608	0.406
IV Regression Diagnostics								
Under identification test								
Kleibergen-Paap rk LM	—	16.640	—	18.406	—	18.408	—	17.647
	—	[0.001]	—	[0.000]	—	[0.000]	—	[0.001]
Weak identification test								
Kleibergen-Paap rk Wald F	—	5.820	—	6.244	—	6.241	—	6.054
Over identification test								
Hansen J statistic	—	1.131	—	1.329	—	1.188	—	1.015
	—	[0.568]	—	[0.514]	—	[0.552]	—	[0.602]
Endogeneity test	—	29.047	—	14.436	—	14.328	—	17.602
	—	[0.000]	—	[0.001]	—	[0.001]	—	[0.000]

Continued

Table 4. *Continued*

	Hospital Market Boundaries: Driving-Time Radius							
	50-minute		*60-minute*		*70-minute*		*80-minute*	
Variables	*OLS*	*IV*	*OLS*	*IV*	*OLS*	*IV*	*OLS*	*IV*
Lagged hospital FTHHI	0.089***	0.155***	0.087***	0.180***	0.082***	0.210***	0.075***	0.242***
	(0.013)	(0.019)	(0.014)	(0.021)	(0.016)	(0.024)	(0.017)	(0.028)
Lagged health insurer HHI	0.039**	−0.245	0.039**	−0.328	0.040**	−0.409*	0.041**	−0.475**
	(0.018)	(0.217)	(0.018)	(0.222)	(0.018)	(0.230)	(0.018)	(0.238)
No. of observations	3,963,234	3,963,234	3,963,234	3,963,234	3,963,234	3,963,234	3,963,234	3,963,234
R-squared	0.608	0.403	0.608	0.397	0.608	0.391	0.608	0.383
IV Regression Diagnostics								
Under identification test								
Kleibergen-Paap rk LM	—	16.678	—	15.561	—	14.294	—	13.036
	—	[0.001]	—	[0.001]	—	[0.003]	—	[0.005]
Weak identification test								
Kleibergen-Paap rk Wald F	—	5.790	—	5.477	—	5.123	—	4.772
Over identification test								
Hansen J statistic	—	1.095	—	1.216	—	1.209	—	0.980
	—	[0.578]	—	[0.544]	—	[0.546]	—	[0.613]
Endogeneity test	—	22.074	—	25.802	—	28.610	—	31.204
	—	[0.000]	—	[0.000]	—	[0.000]	—	[0.000]

Notes: The dependent variable is the logarithm of total gross facility payments for a hospital admission. Hospital and health insurer markets concentration measures are one-year lagged values. The instruments, all in logarithms, are *the county population aged 65 and over*, *the number of establishments with at least 100 employees in the MSA*, and their interaction terms with *a 2-year lag of a categorical variable measuring the comprehensiveness of states' CON regulations.* All specifications include patient-specific and county-level patient controls, hospital-specific and county-level hospital cost and quality controls, and fixed effects for condition category and state-year. Robust standard errors (clustered by county) are in parentheses. ***, **, and * indicate significance at 1 percent, 5 percent, and 10 percent. p-Values for the IV regression diagnostics are in square brackets. For two endogenous variables and four instruments, the Stock and Yogo (2005) critical values for 5%, 10%, and 20% maximal IV relative bias are 11.04, 7.56, and 5.57. The critical values for 10%, 15%, 20%, and 25% maximal IV size are 16.87, 9.93, 7.54, and 6.28.

Table 5: The Price Effect of Hospital Market Concentration: Hospital Fixed Effects Balanced Panel Results

Variable	*10-minute*	*20-minute*	*30-minute*	*40-minute*	*50-minute*	*60-minute*	*70-minute*	*80-minute*
OLS results								
Hospital FTHHI	0.127***	0.081***	0.074***	0.074***	0.073***	0.070***	0.064***	0.056***
	(0.014)	(0.006)	(0.006)	(0.007)	(0.009)	(0.011)	(0.013)	(0.014)
No. of observations	1,910,252	1,910,252	1,910,252	1,910,252	1,910,252	1,910,252	1,910,252	1,910,252
R-squared	0.610	0.610	0.610	0.610	0.610	0.610	0.610	0.610
IV results								
Hospital FTHHI	0.183***	0.083***	0.078***	0.093***	0.121***	0.159***	0.217***	0.301***
	(0.022)	(0.009)	(0.008)	(0.010)	(0.015)	(0.024)	(0.039)	(0.065)
No. of observations	1,910,252	1,910,252	1,910,252	1,910,252	1,910,252	1,910,252	1,910,252	1,910,252
R-squared	0.397	0.397	0.397	0.397	0.397	0.396	0.395	0.391
IV Diagnostics								
Under identification test								
Kleibergen-Paap rk LM	36.434	41.139	38.635	32.985	27.862	24.099	20.188	15.770
	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
Weak identification test								
Kleibergen-Paap rk Wald F	115.055	600.713	491.137	202.255	95.074	54.456	31.772	17.496
Over identification test								
Hansen J statistic	10.184	10.394	10.455	10.581	10.785	10.951	11.213	11.568
	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]
Endogeneity test	3.342	0.178	0.700	4.001	8.649	11.062	12.746	13.248
	[0.068]	[0.673]	[0.403]	[0.045]	[0.003]	[0.001]	[0.000]	[0.000]

Notes: The dependent variable is the logarithm of total gross facility payments for a hospital admission. The instruments, all in logarithms, are *the interaction terms of the county population* and *a 2-year lag of a categorical variable measuring the comprehensiveness of states' CON regulations.* All specifications include patient characteristics and hospitals-year fixed effects. Robust standard errors (clustered by county) are in parentheses. ***, **, and * indicate significance at 1 percent, 5 percent, and 10 percent. *p*-Values for the IV regression diagnostics are in square brackets. For one endogenous variable and two instrumental variables, the Stock and Yogo (2005) critical values for 10%, 15%, 20%, and 25% maximal IV size are 19.93, 11.59, 8.75, and 7.25.

robust. Finally, the results of Melnick, Shen, and Wu (2011) show that although higher insurer concentration leads to lower hospital prices, it is only when concentration values are above 3,200 are prices significantly impacted. Hence, I conduct an extra analysis to allow for nonlinearity in the relationship between insurer concentration and prices by interacting the logarithm of insurer concentration with indicator variables that split insurer HHI into quintiles, with the cutoffs based on the 2005–2008 pooled distribution. Table S5 (Appendix SA1) reports the results, which suggest that the price effect of insurer concentration varies at different levels of concentration, with larger effects at higher quintile distributions. At the 60- to 80-minute boundaries, prices are significantly impacted in this study only when concentration values are above 4,100.

Panel Regression Results

The observed data available make it difficult to avoid potential bias arising from unmeasured differences in the quality of services rendered within and across hospitals overtime, which may affect prices. Therefore, as an extension and alternative to the cross-sectional estimates, I estimate a hospital fixed effects balanced panel regression model with instrumental variables to cleanly identify the price effect of hospital concentration. Table 5 reports both the OLS and IV results. An OLS estimation still biases the hospital concentration estimates downwards. Although the fixed-effect estimates confirm the main cross-sectional results, the magnitudes of the estimates are slightly lower. The IV point estimates show that a 10 percent increase in hospital concentration raises prices by 0.8 to 3.1 percent, while a “5 to 4” hospital merger would cause prices to rise by 3 to 11 percent.

CONCLUSION

This study examines the price effects of hospital and insurer concentrations using a large data set with actual transaction prices. It mitigates potential endogeneity issues using an IV approach with a new identification strategy. I find that hospitals in more concentrated markets are able to use their market power to secure significantly higher prices from insurers and insurers in more concentrated markets are able to exercise a countervailing power and offer significantly lower prices to hospitals.

This study is not without limitations. First, the paper only captures average price effects, although price discrimination is common in insurer–provider negotiations (see Sorensen 2003). Second, for policy purposes, it is also important to examine the effects of hospital and insurer concentration on hospital care quality and not just the price effect. Although recent studies have shown very modest cost-reduction effects of consolidation (see Gaynor and Vogt 2000), I still find evidence of a price increase due to consolidation. Third, this study only considers system consolidation in geographic markets. However, differentiation along clinical service lines may also have important implications for competition and its effect on prices. Further work should examine the impact of both geography and clinical service lines consolidation on prices. These limitations notwithstanding, the findings in this study are consistent with predictions of markets that are less competitive and have important policy implications. The findings provide support for antitrust enforcement and price regulation to curb rising hospital prices and health care costs.

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Disclaimer: None.

Note

1. For these MEPS categories, see http://meps.ahrq.gov/mepsweb/data_stats/conditions07.shtml.

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Supporting Information

Additional supporting information may be found online in the supporting information tab for this article:

Table S1. Variables Definition

Table S2. The Price Effect of Hospital and Health Insurer Markets Concentration.

Table S3. First-Stage and Reduced Form Results.

Table S4. The Price Effect of Hospital and Health Insurer Markets Concentration Using Alternative Construction of Hospital FTHHI.

Table S5. The Price Effects of Markets Concentration Using Alternative Insurer HHI.