THE PRICE AIN'T RIGHT? HOSPITAL PRICES AND HEALTH SPENDING ON THE PRIVATELY INSURED*

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We use insurance claims data covering 28% of individuals with employer-sponsored health insurance in the United States to study the variation in health spending on the privately insured, examine the structure of insurer-hospital contracts, and analyze the variation in hospital prices across the nation. Health spending per privately insured beneficiary differs by a factor of three across geographic areas and has a very low correlation with Medicare spending. For the privately insured, half of the spending variation is driven by price variation across regions, and half is driven by quantity variation. Prices vary substantially across regions, across hospitals within regions, and even within hospitals. For example, even for a nearly homogeneous service such as lower-limb magnetic resonance imaging, about a fifth of the total case-level price variation occurs within a hospital in the cross section. Hospital market structure is strongly associated with price levels and contract structure. Prices at monopoly hospitals are 12% higher than those in markets with four or more rivals. Monopoly hospitals also have contracts that load more risk on insurers (e.g., they have more cases with prices set as a share of their charges). In concentrated insurer markets the opposite occurs—hospitals have lower prices and bear more financial risk. Examining the 366 mergers and

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acquisitions that occurred between 2007 and 2011, we find that prices increased by over 6% when the merging hospitals were geographically close (e.g., 5 miles or less apart), but not when the hospitals were geographically distant (e.g., over 25 miles apart). *JEL* Codes: I11, L10, L11.

I. Introduction

Over 55% of the U.S. population has private health insurance. In 2017, the average insurance premium for employer-sponsored health coverage for a family of four was $18,764, and between 2007 and 2017, premiums increased by about 55% (Kaiser Family Foundation 2017). However, because of data availability, most of our understanding of health care spending comes from the analysis of the Medicare program, which covers less than 15% of the population.[1] For the most part, Medicare pays hospitals using prospectively set, formula-based reimbursements. By contrast, hospital prices for the privately insured are set via negotiations between hospitals and insurers. Unfortunately, private health insurance claims data in general and the results of these hospital/insurer negotiations in particular—hospitals' transaction prices—have been treated as commercially sensitive and have been largely unavailable to researchers.

In this study, we use newly accessible claims data from three of the five largest private insurers in the United States to study the variation in health spending on the privately insured. Notably, the data we use includes hospitals' transaction prices. As a result, we are able to study the role that variation in hospitals' prices plays in influencing health spending variation for the privately insured; describe the variation in hospital prices across regions, within regions, and within hospitals; and analyze the extent to which hospital and insurer market structures are associated with hospital price levels and the design of insurer-hospital payments (henceforth, "contracts").

The main data we use here are claims from Aetna, Humana, and UnitedHealthcare, which were provided by the Health Care Cost Institute (HCCI). Our data capture the claims from the

1. Our discussion of Medicare is focused on the traditional, publicly administered Medicare program. See Curtu et al. (2017) for a comparison of the traditional, public Medicare program and the privately administered Medicare Advantage program. The remainder of the population have coverage from the Medicaid program, other payers (e.g., the Veterans Administration), or are uninsured.

health care services delivered to 27.6% of individuals in the United States with employer-sponsored coverage between 2007 and 2011. The data include more than 88 million unique individuals and capture over $125 billion in health spending a year. The article proceeds in three stages.

First, we present a national picture of the variation in health spending per privately insured beneficiary across all 306 hospital referral regions (HRRs) in the United States.[2] Risk-adjusted health spending per privately insured beneficiary age 18 to 64 varies by a factor of more than three across these regions. The HRR in the 90th percentile of the spending distribution (Grand Junction, CO) spends 47% more than the HRR in the 10th percentile of the spending distribution (Sarasota, FL). Spending per privately insured beneficiary and spending per Medicare beneficiary have a correlation of only 0.044 across HRRs. For the Medicare program (where prices are set administratively), variation in hospital reimbursement rates account for only 13% of the variation in spending across regions, whereas the variation in the quantity of care delivered across regions accounts for 95% of the national variation in spending (these sum to more than 100% because a covariance term accounts for −8%). This fact has motivated research analyzing the factors that drive variation in the amount of care delivered across regions (e.g., Finkelstein, Gentzkow, and Williams 2016; Cutler et al. 2017). By contrast, for the privately insured, about half of the variation in spending is driven by price variation across regions, and half is driven by quantity variation. This motivates us to focus on analyzing the drivers of hospital price variation.

The second stage of our analysis looks at the variation in hospital prices and the structure of hospital payment contracts. Hospital care represents nearly 6% of GDP (Centers for Disease Control and Prevention 2017) and is expensive—the average price of an inpatient case in 2011 is $14,240 in our data. Hospital prices vary significantly across the country and across hospitals within HRRs. For example, hospitals with risk-adjusted knee

2. Hospital referral regions are geographic regions created by researchers at the Dartmouth Institute for Health Care Policy to approximate markets for tertiary medical care in the United States. Each HRR generally includes at least one major referral center, and the United States is divided into 306 HRRs. See http://www.dartmouthatlas.org/downloads/methods/geogappdx.pdf for more information.

replacement prices in the 90th percentile of the national distribution of hospitals are 2.3 times as expensive as hospitals in the 10th percentile. Likewise, in one representative HRR (Philadelphia, PA), the hospital in the 90th percentile of prices in the region is more than twice as expensive as the hospital in the 10th percentile. This variation is also present for plausibly undifferentiated services, such as lower-limb magnetic resonance imaging (MRI), which suggests that the dispersion we observe is not simply a function of differences in hospital quality or patient severity across providers.

Our data also allow us to extend beyond previous analysis and identify the variation in prices for health care services delivered within hospitals. We find that the variation in prices within hospitals for services ranging from joint replacement to lower-limb MRI is substantial. Over a fifth of the total price variation across cases in the average month-year occurs within hospitals for the same procedure, after controlling for hospital fixed effects, insurance plan characteristics, and patient characteristics. That there is such substantial variation in prices for plausibly undifferentiated procedures such as lower-limb MRIs within hospitals suggests that the relative bargaining power of insurers with hospitals can strongly influence price levels.

We then analyze how hospitals are paid. Although there has been recent work looking at how physicians set their negotiated prices with commercial insurers (Clemens, Gottlieb, and Molnar 2017), much less is known about insurer-hospital contracts. We find that about 23% of hospitals' inpatient cases have prices set as a share of hospitals' charges—a form of contract that loads idiosyncratic patient risk onto the insurers.[3] We estimate no more than 57% of cases are on contracts where prices are prospectively set as a percentage of Medicare payment rates. This implies that hospital prices are less closely linked to the Medicare fee schedule than the 75% of cases that Clemens and Gottlieb (2017) observed for physicians' prices.

In the third stage of our analysis, we look at whether there is a link between market structure, hospital prices, and contractual form. Hospital prices and contract form are determined by bargaining between hospitals and insurers. Market structure

3. Hospital charges are the amount hospitals bill for care (i.e., their list prices). Individuals who self-fund their care are typically the only ones who pay hospitals their charges.

is related to bargaining power—hospitals with fewer potential competitors are likely in a stronger negotiating position with insurers, and vice versa. Further motivating this analysis, as we illustrate in Online Appendix Figure I, there has been significant consolidation in the hospital sector between 2001 and 2011. During that period, based on data we collected, there were on average 66 merger and acquisition (M&A) transactions a year.[4] This led the Herfindahl–Hirschman Index (HHI) in hospital markets where mergers occurred to increase by 19% over this period.[5]

In our cross-sectional analysis, we find that hospitals in monopoly markets (relative to hospitals in quadropoly or greater markets) have 12.5% higher prices, 10.5 percentage points more cases paid as a share of charges (over a mean of 18.6%), and 11.3 percentage points fewer of their prospectively paid cases that have prices set as a share of Medicare payment rates (over a mean of 48.3%).[6] By contrast, hospitals located in areas where the three insurers in our data had a high (collective) market share had significantly lower prices and participated in contracts that exposed insurers to less financial risk. A 10 percentage point increase in the insurers' market share is associated with 7% lower prices, 4 percentage points fewer cases paid as a share of charges, and 6 percentage points more prospectively paid cases that have prices set as a percentage of Medicare payments.

To look at events that shifted market structure over time, we use our comprehensive database of hospital mergers combined with the HCCI panel data to examine how hospital prices evolve before and after merger events using difference-in-differences analysis. After mergers occurred, we find that prices increase by over 6% if the merging hospitals were close neighbors (less than or equal to five miles apart). The size of the postmerger price increases declines as the distance between merging parties increases, and there are no significant merger coefficients once merging hospitals are located more than 25 miles apart. We find

4. We have made our roster of hospital mergers available at http://www.healthcarepricingproject.org.

5. We measure a HHI for each hospital in our data within a circular area around each hospital defined by a 15-mile radius. We measure a hospital's market share as its share of total hospital beds in those areas.

6. We measure hospital market structure by counting competitors within a circular area around each hospital defined by a radius of 15 miles. In the results section we show that our results are robust to many alternative measures of hospital market structure and different market definitions.

no premerger differences in trends in prices between merging and nonmerging hospitals and show that our results are robust when we use various procedures to match treated and control hospitals.

Our article builds on a sizable literature that has used Medicare claims data to document large variations in health spending per beneficiary across HRRs (Fisher et al. 2003a,b; Finkelstein, Gentzkow, and Williams 2016). A smaller literature has documented similar variation in spending on privately insured individuals using limited data samples. Both Chernew et al. (2010) and Newhouse et al. (2013) have documented a low correlation between Medicare spending per beneficiary and private spending per beneficiary across HRRs. We add to this literature by using a much larger and more comprehensive national data set to analyze health spending on the privately insured, by analyzing hospitals' transaction prices, and by addressing the key question of why prices are so high in some regions but not in others. Crucially, our data on hospitals' transaction prices allow us to probe more deeply the claim in Chernew et al. (2010) and Philipson et al. (2010) that variation in health spending on the privately insured is driven by differences in hospital prices across regions.

We also add to an existing literature that used limited data sets to analyze variation in hospital transaction prices. Most of this literature has focused on describing differences in prices across regions (e.g., Government Accountability Office 2005; Ginsburg 2010; Coakley 2011; White, Reschovsky, and Bond 2014). We add to this literature by using data that cover the majority of hospitals nationally.[7] This allows us to look at national variation in hospitals' prices and compare hospital prices across and within geographic areas. Likewise, we risk-adjust prices, look at narrowly defined procedures (e.g., joint replacements without complications), and focus on plausibly homogeneous services (e.g., lower-limb MRIs). Collectively, this allows us to more effectively compare prices across hospitals by reducing the potential bias from differences in quality and patient characteristics across hospitals. In addition, this is one of the first publications we are aware of that has described and quantified variation in prices within

7. Our data contain transaction prices for 72% of noncritical access hospitals that are registered with the American Hospital Association (AHA). These 2,358 hospitals in our inpatient sample capture over 88% of total hospital admissions in the United States (based on AHA data). Previous studies have generally relied on data from single states, a single employer, or a small set of urban areas.

hospitals. Analyzing price variation within hospitals for broadly undifferentiated services allows us to hold quality constant. That we observe significant variation in prices across contracts within the same hospital provides evidence that the bargaining leverage of insurers influences hospital prices.

Finally, we add to a large body of literature on hospital competition (see Gaynor, Ho, and Town 2015), which has generally found that hospital prices are higher in more concentrated markets. However, much of this literature has relied on estimates of transaction prices based on hospitals' charges (rather than actual data on transaction prices) or has focused on data from limited areas or single states (often California). Our analysis shows that there is a positive but rather low correlation (0.314) between hospital charges and hospitals' transaction prices. Moreover, we go beyond existing work by looking at the relationship between market structure and transaction prices using data from across the nation and analyzing the relationship between market structure and the design of hospital-insurer contracts. Our findings are broadly consistent with models of insurer-hospital bargaining, such as Gowrisankaran, Nevo, and Town (2015) and Ho and Lee (2017). There is also an existing literature that has examined the effects of single mergers or small groups of mergers.[8] We add to this literature by examining the postmerger price effects of all hospital mergers between 2007 and 2011.

This article is structured as follows. In Section II we outline our data, describe how we measure prices, and present descriptive statistics. In Section III we describe the variation in health spending across HRRs and determine the share of the variation that is a function of price differences across regions and the share that is a function of quantity differences. In Section IV, we describe the variation in hospital prices across HRRs, within HRRs, and within hospitals. In Section V, we describe insurer-hospital contracts. We then analyze the cross-sectional correlates of hospital price levels and contracts in Section VI, analyze mergers and hospital prices in Section VII, and make some concluding comments in Section VIII. Our Online Appendix gives more details on data (A), how we construct risk-adjusted prices (B), our measures of market structure (C), how we identified mergers

8. See Gaynor, Ho, and Town (2015) for a summary of this literature. The exception is Dafny (2009), which examines the effect of 97 mergers that occurred between 1989 and 1996.

(D), econometric matching methods used in our merger analysis (E), and the robustness of our analysis in areas where Blue Cross Blue Shield (BCBS) insurers had high and low market share (F).

II. Data and Variables

II.A. Health Care Cost Institute Data

The main data we use are from the Health Care Cost Institute (HCCI).[9] We discuss the data in more detail in Online Appendix A, but outline some of the main features here. The HCCI database includes health insurance claims for individuals with coverage from three of the five largest insurance companies in the United States: Aetna, Humana, and UnitedHealthcare. The data cover all health services paid for by the insurers from 2007 to 2011. We focus on individuals with employer-sponsored coverage who are aged 18 to 64 and for whom an HCCI payer is their primary insurer. The raw data covers 2.92 billion claims that were delivered to an insured population in our data of 88.7 million unique individuals (Online Appendix Table I).[10]

Online Appendix Figure II shows the proportion of privately insured lives that the HCCI data cover by state.[11] The HCCI database offers a significantly more comprehensive picture of private health spending across the United States than do other private health insurance claims databases. The most prominent alternative data set of private health insurance claims is the MarketScan database. Although MarketScan data include individuals in 90% of HRRs in the United States, some have very thin coverage and include fewer than 200 beneficiaries. By contrast, the HCCI data include individuals in all 306 HRRs, and the

9. HCCI is a nonprofit organization dedicated to advancing knowledge about U.S. health care costs and utilization. See http://www.healthcostinstitute.org for more information.

10. The HCCI data are deidentified and do not include patient identifiers such as Social Security numbers, names, dates of birth, or addresses. Users of HCCI data are not allowed to publish results that identify patients, insurers, or hospitals by name. Because our data is deidentified, our project was exempted by the Yale Institutional Review Board.

11. The data capture more than 30% of the privately insured population in Texas, Arizona, Colorado, Florida, Georgia, Kentucky, Ohio, Wisconsin, New Jersey, and Rhode Island. At the low end, the data capture between 1.9% and 10% of the privately insured in Vermont, Michigan, Alabama, Wyoming, Montana, South Dakota, and Hawaii.

smallest HRR in 2011 has 2,932 beneficiaries. Online Appendix A.1 gives a more detailed comparison between the data sets.

Although we describe the most comprehensive picture to date of health spending on the privately insured, we do not have claims from every insurer, in particular from BCBS insurers. As a result, our analysis does not necessarily generalize to private health insurance spending in the United States as a whole. BCBS plans covered 41% of covered lives across the small, medium, and large group markets in 2011.[12] To address possible concerns about the generalizability of our results, Online Appendix F reproduces all our main results using data from areas where BCBS plans have a high share of privately insured lives and areas where BCBS plans have a low share of privately insured lives.

The HCCI data include a unique hospital identifier, a unique patient identifier, the date services were provided, hospitals' charges (for 2010 and 2011), hospitals' negotiated transaction prices (broken down by facility and physician fees), and payments to hospitals made by patients in the form of coinsurance payments, copayments, and payments made before deductibles were met. As a result, we know the amounts paid to hospitals for all health care encounters recorded in our data.[13] This allows us to analyze how prices vary within and across hospitals and study how insurers reimburse hospitals.

We use an encrypted version of hospitals' National Plan and Provider Identification System (NPI) code in the HCCI data to link to data on hospital characteristics from the American Hospital Association (AHA) annual survey, quality scores from Medicare's Hospital Compare webpage, Medicare activity data from the 100% sample of Medicare claims (accessed via the American Hospital Directory [AHD]), Medicare reimbursement information from the Centers for Medicare and Medicaid Services (CMS), and reputational quality scores from *U.S. News & World Report*. We use hospitals' five-digit postal codes to link to local

12. BCBS is an association of 36 for-profit and not-for-profit health insurance companies in the United States. The BCBS insurance companies are licensees, the largest of which, Anthem, is a for-profit publicly traded firm that has beneficiaries in 14 states. For more information on BCBS, see http://www.bcbs.com. We identify BCBS market share using data from HealthLeaders Interstudy, which is described in more detail in Online Appendix A.

13. We present a sample hip replacement case constructed from claims data online at http://healthcarepricingproject.org/sites/default/files/papers/sample_hip_claims.xlsx.

area characteristics from the census. We use the system ID from the AHA data to identify multiple hospitals that are part of the same health system when we calculate our measures of hospital market structure.[14] The AHA annual survey sometimes consolidates hospital IDs when two hospitals merge, even when the hospitals both remain open. We use various data sources to continue tracking the original hospitals even after consolidation and to create a consistent longitudinal database of hospital sites.[15]

II.B. Sample Definitions

To support our analysis, we create three broad subsamples from the raw HCCI data: the "spending samples," the "inpatient price sample," and the "procedure samples."

The spending samples measure inpatient and overall spending per privately insured beneficiary. Our measure of total spending per beneficiary captures the sum of spending on inpatient, outpatient, and physician services, but excludes drug spending (we exclude prescription drug spending because it is not readily available for Medicare beneficiaries). Our measure of inpatient spending only captures inpatient hospital spending. We calculate spending per beneficiary by summing total or inpatient spending for each individual in our data in each HRR per year. To get the total number of private beneficiaries per HRR, we sum the member months of coverage per HRR per year and divide by 12. We use data from the Dartmouth Atlas for 2008 through 2011 to analyze variation in spending per Medicare beneficiary.[16] Following the approach taken by Dartmouth, we risk-adjust our HCCI spending samples for age and sex.[17] In our decomposition of Medicare spending, we use data from the 100% sample of Medicare claims

14. Hospitals that are part of the same health system are under common ownership (i.e., they are different establishments that are part of the same firm).

15. A complete list of data sources is contained in Online Appendix A.1 and our process for identifying hospitals using their NPI code is outlined in Online Appendix A.2. In Online Appendix A.3, we detail our method for maintaining a consistent hospital-level panel database in the face of merger activity.

16. Data from the Dartmouth Atlas can be downloaded at: http://www.dartmouthatlas.org/tools/downloads.aspx. Information on how Medicare spending per beneficiary is calculated is available in their Research Methods document, accessible at: http://www.dartmouthatlas.org/downloads/methods/research_methods.pdf.

17. Because we do not have data on race, we risk-adjust using age and sex as opposed to Dartmouth, which risk-adjusts using age, sex, and race. Like Dartmouth, we also risk-adjust spending using indirect standardization.

data that identifies how many cases in each diagnosis related group (DRG) case were provided by each hospital in the United States in 2011. Our spending samples include claims for services that were delivered at all providers including, for example, care delivered at critical access hospitals.

The inpatient price sample is derived from hospital claims for all inpatient care provided to our covered population (age 18–64) in AHA registered facilities.[18] In total, there are 3,272 noncritical access hospitals that are registered with the AHA during our sample period (see Online Appendix Table II) and we have all but 70 of them in the HCCI data. We focus our analysis on general medical and surgical hospitals and do not include specialist hospitals (e.g., orthopedic specialty hospitals). We exclude 3 hospitals for which we do not have Medicare payment information and also drop data from 2007 because of incomplete data (this leads to a loss of 10 hospitals). We limit our analysis to providers that delivered 50 or more cases a year, so that we had sufficient data to calculate our inpatient price index. Although this means losing another 831 hospitals, these hospitals only account for 1.5% of our inpatient cases. We are left with 2,358 hospitals in our inpatient sample, which account for 88.4% of the total inpatient cases from the original 3,272 AHA hospitals that were eligible to be included in our analysis (Online Appendix A.4 gives more detail on our sample restrictions).

We also create seven procedure samples, which capture claims for hospital-based surgical or diagnostic inpatient and outpatient procedures. We create procedure samples for hip replacements, knee replacements, cesarean sections, vaginal births, percutaneous transluminal coronary angioplasties (PTCAs), diagnostic colonoscopies, and MRI of lower-limb joints without contrast. These procedures occur with sufficient frequency to

For a detailed discussion of the risk-adjustment methods, see http://www.dartmouthatlas.org/downloads/methods/indirect_adjustment.pdf.

18. Our inpatient data in Online Appendix Table I includes some incomplete records. We exclude the 0.1% of cases that have missing or negative prices. A further 8% of cases are excluded because they are missing a provider identifier or patient characteristics. We exclude cases that have length of stay in the top 1% of the distribution by DRG (these are cases with a length of stay of over six months in some cases). We then remove cases with prices in the top 1% and bottom 1% of the price distribution by DRG. Our results are robust to winsorizing these outliers instead of trimming them.

support empirical analysis and are relatively homogeneous, thereby facilitating comparison across facilities and areas (Centers for Disease Control and Prevention 2010).

Each observation in the seven procedure samples includes all hospital claims from the time the patient entered the hospital until they exited the facility. We limit the observations included in our analysis to those without major medical complications and define the seven procedure samples narrowly using diagnosis and procedure codes to exclude atypical cases (see Online Appendix A.4). We limit our observations to hospitals that deliver at least 10 of a given procedure each year and applied the same cleaning rules we used to define our inpatient sample.[19] In total, from 2008 to 2011, we capture 470 hospitals performing hip replacements, 932 performing knee replacements, 1,163 performing cesarean sections, 1,280 performing vaginal deliveries, 652 performing PTCAs, 1,237 performing colonoscopies, and 1,628 performing lower-limb MRIs who meet our sample restrictions.

Table I reports summary statistics for our inpatient sample.[20] Our sample of hospitals in the inpatient and procedure samples are generally similar to the universe of AHA-registered hospitals, but there are some differences (Online Appendix Table II). These differences are largely due to our requirement that hospitals treat a minimum number of cases in our data annually, which means we are dropping some smaller hospitals. Relative to the universe of AHA-registered hospitals, hospitals in our inpatient sample are larger (an average of 270 beds versus 218 among all AHA hospitals), are located in less concentrated markets, and are more likely to be teaching facilities, nonprofit facilities, and facilities ranked by the *U.S. News & World Report* as top performers.

19. For MRI we also require a separate physician claim for reading the MRI, which we do not include in our main analyses of price. We do this so that the facility portion we analyze only captures the taking of the MRI, as opposed to the reading of the MRI. We also restrict our lower-limb MRI cases to those for which the scan itself was the only intervention occurring during the individual's visit to the hospital. Focusing on MRIs performed during days where nothing else was done to the patient and outside of broader hospital admissions helps attenuate concerns that the scans we analyze are services folded into broader cases.

20. The descriptive statistics for the subsamples for the procedures look qualitatively similar and are available online at http://www.healthcarepricingproject.org.

TABLE I

Hospital and Patient Characteristics

	Mean (1)	Std. dev. (2)	Min (3)	Max (4)
Market characteristics				
Hospital in monopoly market, 15-mile radius	0.163	0.370	0	1
Hospital in duopoly market, 15-mile radius	0.194	0.395	0	1
Hospital in triopoly market, 15-mile radius	0.123	0.328	0	1
Hospital in quadropoly+	0.520	0.500	0	1
Hospital HHI defined by beds in a 15-mile radius	0.461	0.295	0.043	1
HCCI market share measured at the county level	0.178	0.101	0.017	0.571
Blue Cross Blue Shield market share measured at the county level	0.403	0.218	0.001	0.958
Hospital characteristics				
Number of technologies	59	30	0	138
Ranked in *U.S. News & World Report*	0.053	0.225	0	1
Beds	270	203	10	2,264
Teaching hospital	0.380	0.485	0	1
Government owned	0.122	0.327	0	1
Nonprofit	0.693	0.461	0	1
For-profit	0.185	0.388	0	1
Local area characteristics				
Percent of county uninsured	0.171	0.058	0.031	0.389
Median income ($)	51,516	13,153	22,255	119,525
Rural	0.162	0.369	0	1
Other payers				
Medicare payment rate	6,437	1,288	4,590	14,292
Share Medicare	0.446	0.101	0	0.833
Share Medicaid	0.188	0.096	0	0.777
Quality scores				
30-day AMI survival rate	0.840	0.016	0.751	0.898
% of AMI patients given aspirin at arrival	0.975	0.049	0.330	1
% of patients given antibiotics presurgery	0.934	0.082	0.140	1
% of surgery patients given treatment to prevent blood clots	0.881	0.106	0.030	1
Patient characteristics				
Age 18–24	0.074	0.262	0	1
Age 25–35	0.248	0.432	0	1
Age 35–44	0.196	0.397	0	1
Age 45–54	0.219	0.414	0	1
Age 55–64	0.262	0.440	0	1
Female	0.672	0.470	0	1
Charlson Comorbidity index	0.707	1.442	0	6

Notes. These are descriptive statistics for the inpatient pricing sample from the HCCI database. There are 8,772 hospital-year observations representing 2,358 unique hospitals and 4,964,774 unique patients.

II.C. Measuring Hospital-Level Prices

Hospitals vary in the mix of services they offer and the patients they treat. As a result, a general concern when analyzing differences in prices across hospitals is that variation in prices could reflect observed and unobserved differences in the quality of care, mix of care, or the quantity of care provided per case at different facilities. For example, if patients with a given condition at a hospital were more severely ill, they would require more care, which could potentially show up in our data as higher prices. Likewise, providing higher quality care could raise costs, so a hospital that had a higher quality of care could show up in our data as having higher prices.

We work to address these issues in a number of ways. First, we rely on risk-adjusted price measures, described in detail in Online Appendix B. Second, we show that our results are stable when we control for hospital quality using a variety of measures. Third, we measure price variation across plausibly undifferentiated services (like lower-limb MRI) for which there is little variation in how these services are delivered across hospitals or across patients within a hospital. Since MRIs are plausibly homogeneous across patients, studying this procedure provides a reasonable benchmark for price variation that is uncontaminated by unobservable patient heterogeneity. Fourth, we define our procedures narrowly via our choice of clinical codes and exclude cases with complications. Finally, we limit the age of patients we analyze by procedure to fairly narrow age groups (since older patients or atypically young patients may raise costs). For knee and hip replacements, we limit our analysis to cases involving patients between 45 and 64 years old. For cesarean and vaginal delivery, we limit our analysis to mothers who are between 25 and 34 years old.

Our hospital price measures are generated from data on the actual payments patients and insurers make to hospitals. We construct three different measures of hospital prices based on these allowed amounts (i.e., the sum of the patient and insurer payments to hospitals). The first is a private payer overall inpatient price index that is adjusted for the mix of care a hospital delivers (via DRG fixed effects) and the mix of patients that hospitals treat (we risk-adjust for patient age and sex). This hospital-level, regression-based measure is similar to those used previously in the literature (e.g., Gaynor and Vogt 2003; Gowrisankaran, Nevo, and Town 2015). The second is a set of hospital-level and risk-adjusted price measures for each of our seven procedures.[21] Third, we focus on contract-level prices within hospitals for the seven previously identified procedures. We also construct Medicare reimbursement rates for overall inpatient care and for the seven procedures in our analysis. More details on our price and Medicare reimbursement measures are in Online Appendix B.

II.D. Descriptive Statistics on Prices

Online Appendix Table 4 presents summary statistics for our main price measures and the within-hospital correlations of the inpatient hospital price index, the procedure prices, and the Medicare inpatient base payment rates. There is high correlation in prices within hospitals within service lines like orthopedics (e.g., the correlation of hip with knee replacements is 0.923) and a weaker (but still substantial) correlation across service lines (e.g., the correlation of knee replacement with vaginal delivery prices is 0.510). By contrast, there is a low correlation within hospitals between the Medicare base payment rate and the inpatient price index (0.203) and between Medicare procedure-specific reimbursements and private payment rates for the procedures we study (these range from −0.040 to 0.360). Medicare attempts to set administered prices to reflect hospitals' exogenous costs (e.g., local labor costs) and therefore, the low correlation between Medicare and private prices suggests that private price variation is driven by more than simply differences in costs across hospitals.

The difference in the amounts that Medicare and private insurers pay for services is substantial. Figure I shows that in 2011, Medicare payments were 45% lower than private rates for inpatient care, 55% of private rates for hip and knee replacement, 62% for cesarean and vaginal delivery, 51% for PTCA, 37% for colonoscopy, and 25% for MRIs. As an illustration of the magnitude of this difference, we calculate that if private prices were set at 120% of Medicare rates rather than at their current

21. For inpatient procedures, the procedure price captures the combined price on all claims associated with services provided to the patient by hospitals from admission through discharge. For outpatient procedures (colonoscopies and MRIs), the price is the sum of all claims on the day the patient was in the hospital for the procedure. For colonoscopies and MRIs, we further limit our analysis to observations where no other medical care was provided to the patient on the day of the MRI or colonoscopy and exclude MRIs and colonoscopies that were performed within a wider hospital stay. As a robustness check, we examine the sum of hospital and physician prices for inpatient and procedure prices.

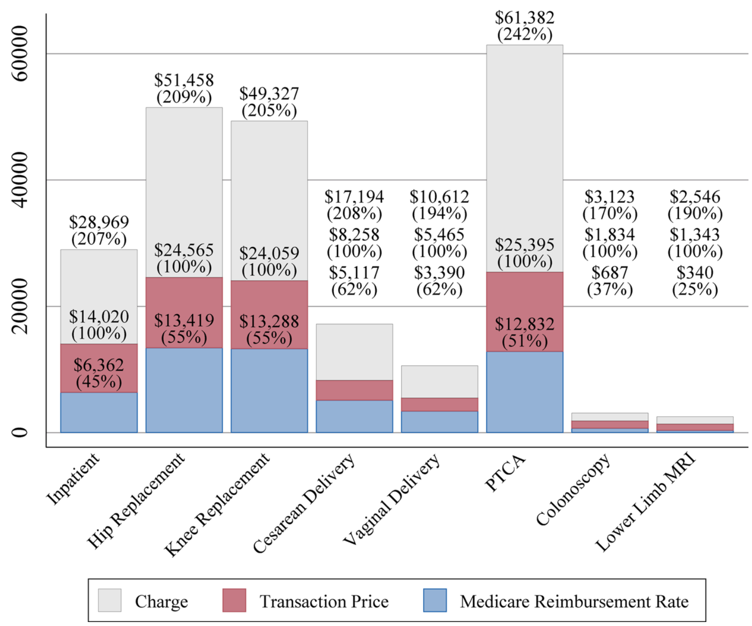


FIGURE I

Average Hospital Facilities Charges, Transaction Prices, and Medicare Reimbursements, 2011

Data drawn from the inpatient and procedures samples. The height of the light gray bars (top) are the average hospital charges. The height of the darker shaded bars (middle) are the transaction prices. Both are risk-adjusted as described in Online Appendix B.1 and B.2. The intermediate shaded bars (bottom) are the Medicare reimbursements as described in Online Appendix B.4. Prices are given in 2011 dollar amounts and as a percentage of the transaction prices (in parentheses).

levels, inpatient spending on the privately insured would drop by 19.7%.[22]

There has also been significant recent interest in hospitals' charges—the list prices for hospital services (e.g., Brill 2013; Bai and Anderson 2015; Hsia and Akosa Antwi 2014). Indeed, in 2013, the Department of Health and Human Services began releasing

22. This thought experiment holds the quantities of care constant (i.e., it assumes no behavioral response). We also find that paying providers for inpatient care at 100% of Medicare rates, 110% of Medicare rates, 130% of Medicare rates, and 140% of Medicare rates would lower spending by 33.1%, 26.4%, 13%, and 6.3%, respectively.

hospital charge information for all inpatient claims billed to Medicare (Department of Health and Human Services 2013). Figure I illustrates that charges are between 170% and 242% of the transaction prices. Online Appendix Figure III presents a scatterplot showing the relationship between hospital charges and transaction prices for the procedures in our analysis in 2011. The correlations are positive, but all below 0.5 in magnitude and range from 0.243 (lower-limb MRIs) to 0.471 (vaginal deliveries).

In the absence of available data on true transaction prices, a number of research papers have used transformations of hospital charges to produce proxies for hospitals' transaction prices. Unsurprisingly, we observe that transformations of charges are not very highly correlated with transaction prices in our data. Using data kindly provided by Dafny, Ho, and Lee (2016), we find that the correlation between our main inpatient price index that is constructed using transaction prices and their price measure constructed using hospital charge data is 0.45. Although the Dafny, Ho, and Lee (2016) measure contains useful information (Garmon 2017), the low correlation illustrates the advantage of using transaction prices if such data are available.

III. Health Care Spending Variation

III.A. Geographic Variation in Spending per Privately Insured Beneficiary

In Figure II, Panel A, we map total risk-adjusted spending per privately insured beneficiary across HRRs. In 2011, mean spending per beneficiary was $4,197. Total spending per privately insured beneficiary in the highest spending HRR (Anchorage, AK) was $6,366, more than three times as much as spending per beneficiary in the lowest spending HRR (Honolulu, HI, spent $2,110 per person). Likewise, the HRR in the 90th percentile of the spending distribution (Grand Junction, CO) spent 47.3% more than the HRR in the 10th percentile of the spending distribution (Sarasota, FL).[23]

23. We also present a map of inpatient spending per privately insured beneficiary in Online Appendix Figure IV. Inpatient spending per privately insured beneficiary has a correlation with total spending per beneficiary of 0.774. Total spending per privately insured beneficiary per HRR has a 0.468 correlation with spending per beneficiary on hip and knee replacements, 0.369 with cesarean sections, 0.335 with vaginal deliveries, and 0.393 with PTCA.

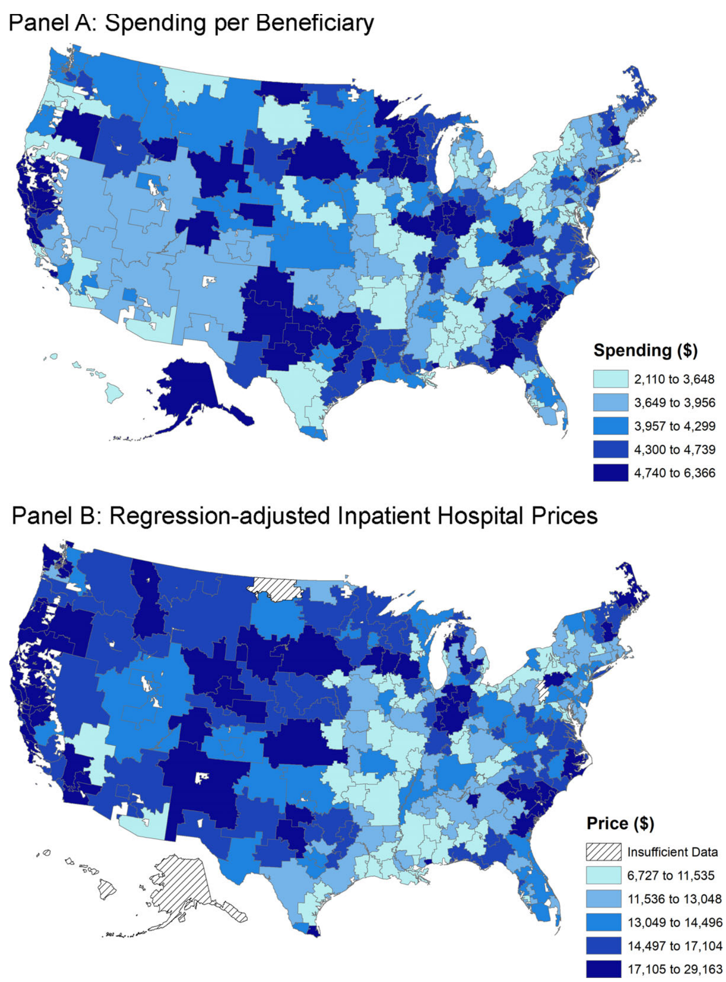


FIGURE II

Total Private Spending by HRR, 2011

Panel A captures risk-adjusted spending per beneficiary by HRR using data from 2011. Each bin captures a quintile of spending per beneficiary. The data are drawn from the spending sample. Spending per beneficiary is risk-adjusted for age and sex. Panel B captures HRR-level average hospital regression-adjusted inpatient prices that are risk-adjusted for DRG, age, and sex, and weighted by hospital activity. In Online Appendix Figure VII, we present this map normalized using the Medicare wage-index to control for local wage costs across the United States. Thatched regions are areas where we do not have sufficient data to calculate prices.

Previous work has found that risk-adjusted Medicare spending per beneficiary also varies by a factor of more than three across HRRs (Fisher et al. 2003a,b). In Online Appendix Figure V, we present maps of total and inpatient spending per Medicare beneficiary across HRRs using data made accessible by the Dartmouth Institute. The correlation between HRR-level total spending per Medicare beneficiary and spending per privately insured beneficiary is only 0.044, and the equivalent correlation for inpatient spending is 0.172. The correlation between HRR-level inpatient spending per privately insured beneficiary age 55 to 64 (i.e., a group with a more similar demographic profile to the Medicare population) and spending per Medicare beneficiary across HRRs is still only 0.165.[24]

III.B. The Contributions of Price versus Quantity to Spending Variation

To what extent is the geographic variation in health spending generated by the variation in the price of care versus the quantity of care delivered across regions? Because the Medicare program's administered hospital prices do not vary significantly across providers, it follows that most of the variation in Medicare spending is mainly driven by differences in the quantities of health care across HRRs.[25] By contrast, variation in spending on the privately insured is likely to be a function of both variation in the quantities of care delivered across regions and variation in the market-determined prices that providers and insurers negotiate.

To analyze the relative contributions of price and quantity to spending variation for Medicare and the 55–64-year-old private patients from HCCI we decompose the variance of ln(inpatient spending per beneficiary) for each DRG d into three components:

(1)
$$Var\left(\ln\left(p_r q_r\right)\right) = Var\left(\ln\left(p_r\right)\right) + Var(\ln(q_r)) + 2Cov\left(\ln\left(p_r\right), \ln\left(q_r\right)\right),$$

where p_r is the average price in HRR r and q_r is the number of inpatient visits (quantity) divided by the number of beneficiaries

24. Chernew et al. (2010) find a correlation between private spending per beneficiary measured using MarketScan data and Medicare spending per beneficiary in 2006 of −0.17.

25. Finkelstein, Gentzkow, and Williams (2016) find that 47% of the geographic variation in Medicare use is driven by patient characteristics. The remainder is driven by place-specific factors.

in each HRR. The component $\frac{Var(\ln(p_r))}{Var(\ln(p_r q_r))}$ represents the share of the variance in spending attributable to differences in price across HRRs; the component $\frac{Var(\ln(q_r))}{Var(\ln(p_r q_r))}$ represents the share attributable to differences in quantity and $\frac{2Cov(\ln(p_r),\ln(q_r))}{\text{Var}(\ln(p_r q_r))}$ is the share attributable to the covariance of price and quantity.[26] We obtain these components per DRG.

In Table II we report results for the top 10 DRGs in the data individually, and the final row presents the decomposition results for spending samples averaged across *all* DRGs (where each DRG-observation is weighted by spending on that DRG in the private population in the first three columns and the Medicare population in the last three columns).[27] The bottom row of column (1) shows that averaged across DRGs, just under half of spending variation on the privately insured is due to price and almost the same is due to quantity in column (2) with the covariance term accounting for essentially zero in column (3).[28] Columns (4)–(6) show that for Medicare spending, quantity differences across HRRs account for 95.3% of the variation whereas only 12.7% is attributable to price variation (the residual is a −8.1% covariance term). These results suggest that variation in health spending on the privately insured is a function of variation in both the price and quantity of care delivered across HRRs, while variation in spending on the Medicare population is driven almost exclusively by differences in the quantity of care delivered across regions.[29]

Overall, both populations have similar levels of quantity variation across HRRs where quantity is defined as spending with hospital prices fixed at the mean (we refer to this as fixed-price

26. We focus on inpatient spending because we do not have reimbursement and quantity measures for Medicare outpatient services.

27. Results for the top 25 DRGs are presented in Online Appendix Table V.

28. Later, we focus on two outpatient procedures (colonoscopy and lower-limb MRI) and five inpatient procedures (hip replacement, knee replacement, vaginal baby delivery, cesarean baby delivery, and PTCA). Price explains 29% of the variation in spending on hip and knee replacements, 42% for vaginal deliveries, 40% on cesarean sections, and 34% on PTCAs. In contrast, price variation explains 12% and 10% of the variation in Medicare spending on hip and knee replacement and for PTCAs, respectively.

29. The results are not driven by the particular weighting scheme used. For example, using the Medicare spending weights (by DRG) in the private spending decomposition generates an overall contribution of price of 52% instead of 50% in the final row of column (1).

TABLE II

PRICE/QUANTITY DECOMPOSITION OF MEDICARE AND PRIVATE HEALTH SPENDING, 2011

	Private			Medicare		
	Share price (1)	Share quantity (2)	Share co-variance (3)	Share price (4)	Share quantity (5)	Share co-variance (6)
Respiratory system diagnosis w/ ventilator support 96+ hours	0.650	0.415	−0.064	0.102	0.771	0.127
Percutaneous cardiovasc proc w/ drug-eluting stent w/o MCC	0.465	0.681	−0.146	0.153	1.113	−0.265
Major small & large bowel proc w/ MCC	0.676	0.299	0.025	0.213	0.888	−0.101
Major small & large bowel proc w/ CC	0.474	0.453	0.073	0.193	0.811	−0.005
Esophagitis, gastroent, & misc digest disorders w/o MCC	0.387	0.637	−0.024	0.164	1.028	−0.192
Spinal fusion except cervical w/o MCC	0.334	0.512	0.154	0.085	1.067	−0.152
Major joint replacement or reattachment of lower extremity w/o MCC	0.381	0.645	−0.026	0.213	0.973	−0.186
Infectious & parasitic diseases w/ OR proc w/ MCC	0.701	0.360	−0.061	0.112	0.769	0.119
Septicemia w/o MV 96+ hours w/ MCC	0.536	0.365	0.099	0.120	0.815	0.064
Rehabilitation w/ CC/MCC	0.460	0.430	0.109	0.056	1.164	−0.219
Average shares (weighted by spending)	0.496	0.495	0.009	0.127	0.953	−0.081

Notes. The decomposition of ln(spending per beneficiary) is carried out on the 2011 Medicare and HCCI inpatient spending samples. The Medicare analysis is based on the 100% sample of Medicare claims accessed via the AHD. HCCI data includes all inpatient claims from our spending sample for those aged 55–64. "CC" is short for with "complication or comorbidity"; "MCC" is short for with "major complication or comorbidity"; "proc" = "procedure"; "w/" = "with"; "w/o" = "without". Because of space constraints, we show only the top 25 highest spending DRGs in the HCCI data; the "average shares" in the final row are the average decomposition results by DRG (weighted by spending, i.e., first three columns use spending weights for private and last three use weights based on Medicare) across the 735 DRGs (HCCI) and 562 DRGs (Medicare).

spending, see Online Appendix Table VII).[30] Furthermore, although Medicare and private prices are only weakly correlated at the HRR level (recall that this correlation is only 0.203), the correlation is much stronger for quantities. The correlation of fixed-price spending (quantity) per private beneficiary and fixed-price spending (quantity) per Medicare beneficiary is 0.536 for the private sample of 55–64-year-olds. Similarly, we observe that the correlation in hip and knee replacements delivered per Medicare beneficiary and per privately insured beneficiary per HRR is correlated at 0.570 across HRRs. Finally, we observe that the correlation in hospitals' case-mix indexes—a measure of the average DRG weights at hospitals—across Medicare and privately insured beneficiaries is 0.659. All this suggests, perhaps unsurprisingly, that the quantities of care delivered to Medicare and privately insured beneficiaries are much more correlated than the payment rates from the two sets of payers.

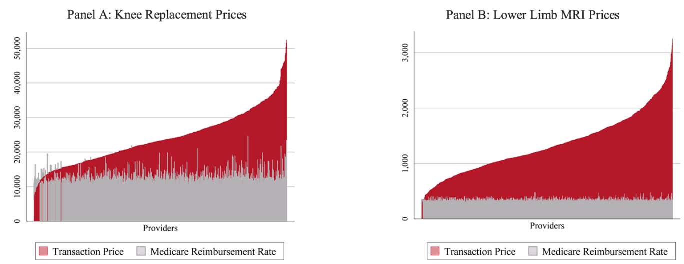
IV. Variation in Hospital Prices

Given the importance of prices for the privately insured, we turn to describing the overall variation in hospital prices and then decompose the amount of variation that occurs in the cross-section (i) across HRRs, (ii) within HRRs across hospitals, and (iii) within hospitals.

IV.A. Quantifying How Much Hospital Prices Vary

Previous research has shown substantial geographic variation in hospital prices for subnational geographies. For example, the U.S. Government Accountability Office (2005) analyzed health care claims data from the Federal Employees Health Benefits Program and found that hospital prices varied by 259% across metropolitan areas. Likewise, the Massachusetts Attorney General's Office (Coakley 2011) found that hospitals' prices varied by over 300% in the state. Ginsburg (2010) used insurance claims data to measure average hospital prices in six cities. Similarly, White, Reschovsky, and Bond (2014) used claims data from autoworkers to examine hospital prices in 13 Midwestern markets. They found that the highest priced hospitals in a market were typically paid 60% more for inpatient care than the lowest priced

30. Online Appendix A.5 describes how these price-fixed and quantity-fixed measures of spending are constructed.



Summary Statistics for Transaction Prices

$24,059	Mean	$1,343
$6,404 – $52,503	Min - Max	$270 – $3,251
$15,115 – $34,554	p10 - p90	$719 – $2,104
$18,260 – $28,698	Inter Quartile Range (IQR)	$965 – $1,671
2.29	p90/10 ratio	2.93
0.32	Coefficient of Variation	0.40
0.18	Gini Coefficient	0.22
664	Number of Hospitals	1,304

FIGURE III

National Variation in Hospital Prices for Knee Replacement and Lower-Limb MRIs, 2011

Each darkly shaded bar represents a single hospital's regression-adjusted transaction price based on 2011 cases. The Medicare payment (lightly shaded bars) is based on the PPS fee schedule described in Online Appendix B.4. The bars are ordered by private price. The summary statistics in the left column refer to knee replacements and those in the right column refer to MRIs.

hospitals.[31] Although extremely valuable, these analyses do not rely on national data, often do not risk-adjust prices for patient case mix, and do not analyze within-hospital price variation.

In Figure III we present the variation in hospital-specific, risk-adjusted private-payer prices for knee replacements across all hospitals in our sample (Panel A). We also include the corresponding hospital-specific Medicare reimbursement rates. Hospitals were paid $24,059 on average for knee replacements in 2011 (Medicare reimbursed these same hospitals $12,986 on average). Across the nation, the ratio of the transaction price for a knee

31. Although notable, this sort of variation is not unique to health care. Many other industries exhibit price variation. Pratt, Wise, and Zeckhauser (1979) find large price variation for a range of services in the Boston area; Hortaşçu and Syverson (2004) document extensive variation in mutual fund fees; Kaplan and Menzio (2015) find significant variation for 36 oz. plastic bottles of Heinz ketchup in Minneapolis in 2007.

replacement at hospitals in the 90th percentile of the price distribution relative to hospitals in the 10th percentile is 2.29.

It is possible that the variation in knee replacement prices across the United States reflects differences in unobserved patient severity or quality across hospitals. Consequently we examine lower-limb MRIs as a plausibly homogeneous procedure free of any contamination due to unobserved heterogeneity. In Figure III, Panel B, we present a histogram of risk-adjusted hospital transaction prices for lower-limb MRIs and show variation that is on a similar scale to knee replacements—the coefficient of variation for knee replacements is 0.32 and for lower-limb MRIs is 0.40. The ratio of the price for a lower-limb MRI at the hospital in the 90th percentile relative to the hospital in the 10th percentile is 2.93 (similar figures for our other procedures are reported in Online Appendix Figure VI).

To determine whether the bulk of the price variation in the cross section occurs across HRRs, within HRRs (across hospitals), or within hospitals, we use our case-level data for 2010 and 2011, add various combinations of control variables into a regression, and observe the subsequent changes in the R^2.[32] In Table III the dependent variable is the price level, $p_{i,p,h,r,t}$, for a case (e.g., a knee replacement) delivered to patient i with insurance plan characteristics p, at hospital h, located in HRR r, in month-year t. In all columns we include month-year dummies, which account for only a trivial fraction of the variance (less than 0.001). Column (1) introduces patient characteristics (sex and age). We then sequentially add in fully interacted insurance plan characteristics, HRR fixed effects, hospital fixed effects, and controls for the hospital charges for each case.[33] We allow HRR fixed effects and hospital fixed effects to vary by month-year pair.

Table III shows that a substantial amount of variation in hospital prices exists across HRRs, within HRRs, and even within hospitals. In column (1) we find that controlling for patient characteristics explains very little of the variation in hospital prices. Indeed, dropping these characteristics would reduce the

32. We focus on these years as we do not have hospital charge information prior to 2010. Results are very similar for the first four columns of Table III for other years.

33. Insurance plan characteristics include the product type (health maintenance organization [HMO], preferred provider organization [PPO], point of service [POS], exclusive provider organization [EPO], indemnity plan, and other), the funding type (administrative services only [ASO] or fully insured plan), and market segment (large versus small group).

TABLE III
DECOMPOSITION OF HOSPITALS' TRANSACTION PRICE VARIATION

	R^2					Unexplained within hospital-month variance (%)	Observations	Within hospital-month coefficient of variation
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Hip replacement	0.006	0.017	0.502	0.763	0.776	22.4	15,122	0.174
Knee replacement	0.006	0.016	0.416	0.728	0.756	24.4	37,157	0.206
Cesarean section	0.011	0.029	0.432	0.726	0.755	24.5	81,482	0.170
Vaginal delivery	0.012	0.030	0.381	0.647	0.701	29.9	108,794	0.192
PTCA	0.005	0.019	0.478	0.724	0.760	24.0	16,636	0.239
Colonoscopy	0.010	0.024	0.412	0.759	0.820	18.0	66,017	0.165
Lower-limb MRI	0.001	0.008	0.331	0.774	0.784	21.6	113,914	0.157
Mean						23.5		0.186
Patient characteristics	Yes	Yes	Yes	Yes	Yes			
Plan characteristics	No	Yes	Yes	Yes	Yes			
HRR fixed effects	No	No	Yes	—	—			
Hospital fixed effects	No	No	No	Yes	Yes			
Control for charges	No	No	No	No	Yes			

Notes. Columns (1)–(5) have transaction-level procedure prices (2010–2011) as the dependent variable and display the R^2 of a regression that includes the relevant right-hand-side variables indicated in the lower rows. All regressions use case-level data and control for month-year dummies. Patient characteristics include fixed effects for sex and five age bands (as in Table I). Plan characteristics include the full interaction of market segment (i.e., large versus small group), and product (HMO, PPO, POS, EOP, indemnity plan, and other) and funding type (fully insured or ASO). "Hospital fixed effects" indicates a full set of hospital dummies interacted with month-year dummies. "HRR fixed effects" indicates a full set of HRR dummies interacted with month-year dummies. Column (6) = 1 − column (5), and the mean is the unweighted average across the seven procedures. Column (8) reports the within-hospital-month coefficient of variation, averaged across hospital-months. The data are drawn from the procedure samples.

R^2 by less than 1.2% across all procedures. In column (2), introducing insurance plan characteristics explains no more than an additional 3%. In column (3), including HRR fixed effects substantially increases the R^2 to between 0.331 (lower-limb MRI) and 0.502 (hip replacements). Column (4) includes hospital fixed effects, which increase the R^2 to between 0.647 (vaginal delivery) and 0.774 (lower-limb MRIs). Although this is a large increase, it still leaves between 22% and 35% of price variation unexplained. In column (5) we include the total charge for each individual case. This is a further control for the patient-specific amount of care that was delivered within a case, since hospitals bill for each unit of service they deliver. Even in this demanding specification, between 18% and 30% of the cross-sectional variation still occurs within hospitals (column (6)) implying that unobserved differences in the cost of providing care cannot account for the unexplained spread of within-hospital prices in column (5).

The sizable variation in prices that we observe within hospitals seems likely to be due to differential insurer bargaining leverage, but potentially it could also be due to measurement error or contract renegotiations that occur within a hospital-month. To address these issues, we focus on MRIs and identify specific hospital/insurer contracts (as described in more detail in Section IV.C). Limiting our analysis to identified contracts excludes cases that have unusually high or low prices due to pure measurement error. This lowers the unexplained variance only slightly (from 21.6% to 19.9%) which is unsurprising because this is administrative (rather than survey) data.[34] Furthermore, because we observe contracts, we also can drop the hospital-month observations when a contract renegotiation occurred. Doing this reduces the unexplained variation to 15.3%. Thus we conclude that over 70% $(=\frac{15.3}{21.6})$ of the unexplained within-hospital MRI price variation in column (6) of Table III is due to cross-insurer price variation within hospitals, rather than measurement error or (within month) contract renegotiation. Although the HCCI data do not identify the specific insurer that covers each beneficiary, these results are suggestive of the substantial degree to which differential insurer bargaining power affects hospital prices.

34. We can classify 97% of the 113,914 MRI cases in Table III to contracts in this way (a higher fraction than for the other procedures). Note that this 1.7 (= 21.6% −19.9%) percentage points is an upper bound for measurement error as it also excludes singleton observations for which we cannot find two matching prices (see Online Appendix B.3).

IV.B. Hospital-Level Price Variation Within and Across HRRs

Figure II, Panel B presents a map of private-payer inpatient prices across HRRs. The map demonstrates that there is substantial variation in prices across geographic areas. Normalizing prices using the Medicare wage index, which captures local labor costs, does not reduce this variation by much (Online Appendix Figure VII). To illustrate the extent of the price variation, Salinas, CA has the highest average inpatient private-payer prices—more than four times as high as the least expensive HRR (Lake Charles, LA). Likewise, the HRR with average hospital inpatient prices in the 90th percentile of the national distribution of HRRs (Eugene, OR) is 1.84 times as expensive as the average inpatient prices for the HRR in the 10th percentile (Lafayette, LA).

Online Appendix Table IX presents the mean prices and coefficients of variation in private-payer prices for our inpatient price index and the seven procedures we analyze for the 25 HRRs with the greatest number of HCCI covered lives. The national averages of the within HRR coefficients of variation range from 0.162 (hip replacement) to 0.249 (MRI). To illustrate how large this variation is, consider the following thought experiment. If each patient paying above the median price in their HRR instead went to the hospital in their HRR with the median price, total inpatient spending for the privately insured would be reduced by 25.8%.[35]

Figure IV illustrates the extent of the variation in hospital prices within a single HRR (Philadelphia, PA) for knee replacements and lower-limb MRIs. The coefficients of variation across hospital-level prices within Philadelphia for knee replacement and lower-limb MRIs are 0.308 and 0.482, respectively. There is a substantial amount of variation in prices for all of these procedures, including lower-limb MRIs (note that there is virtually no variation in Medicare's administered payments across hospitals within HRRs). We find similar variation in hospital prices for all procedures within all HRRs and present the figures for every other HRR online.[36]

35. We calculated this number in the following way. Using data for 2011, we identified the median price for every DRG in the data across all HRRs. For any patient who paid a price above the median for that DRG, we substituted the median price for the actual price and then recalculated average spending per beneficiary. This counterfactual ignores behavioral responses.

36. Our data use agreement precludes us from publicly reporting information about HRRs with fewer than five providers in the data. Within-market price

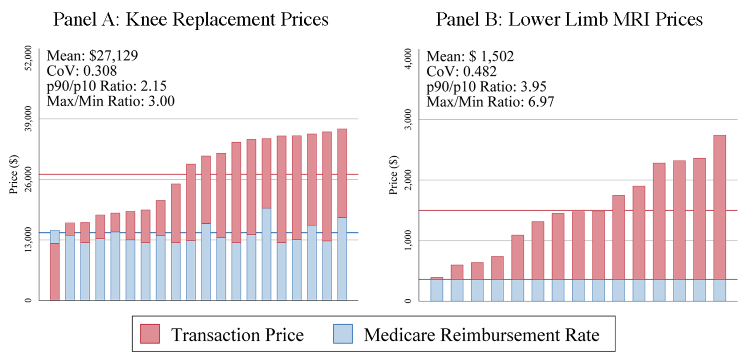


FIGURE IV

Within-Market Hospital Price Variation for Philadelphia, PA, 2011

These panels present average hospital-level regression-adjusted private-payer prices for knee replacements and lower-limb MRIs using data from 2011. Each column captures a hospital in the Philadelphia HRR. We include similar graphs for all procedures and HRRs that include five or more providers at http://www.healthcarepricingproject.org.

IV.C. Within-Hospital Variation in Prices

Table III showed that the amount of within-hospital price variation in the cross-section is substantial. Column (8) shows the within-hospital coefficient of variation by procedure, averaged across every hospital-month, which ranges from 0.157 (lower-limb MRIs) to 0.239 (PTCAs). For reference, the average within-HRR coefficient of variation in MRI prices across hospitals is 0.249 (Online Appendix Table IX).

As a result, to delve into the patterns of contracts within hospitals, we developed an algorithm to identify ongoing hospital/insurer contracts (see Online Appendix B.3 for details). To do so, we find repeated prices at hospitals over time (for a given DRG or procedure) and then pair claims into larger contracts by grouping those that have similar combinations of insurance product characteristics (e.g., HMO versus PPO, or large group products versus small group products). To illustrate these matches, in Figure V we present within-hospital contracted prices for

variation graphs are available for all HRRs with five or more providers for all procedures at http://www.healthcarepricingproject.org.

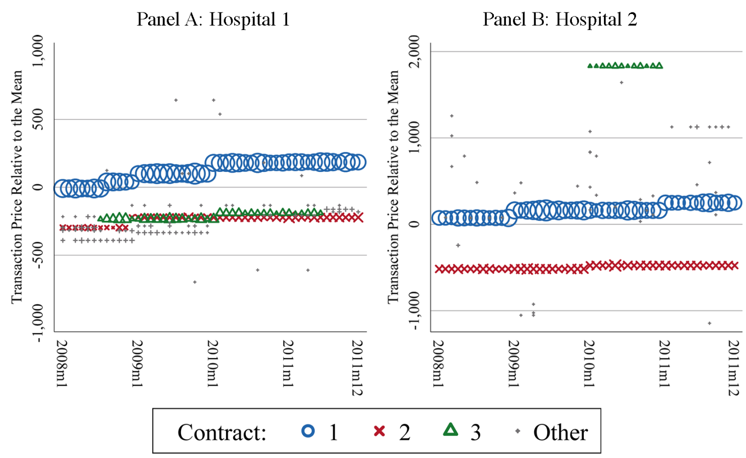


FIGURE V

Within-Hospital Prices for Lower Limb-MRIs at Two High-volume Hospitals, 2008–2011

These figures highlight the top three linked contracts (circles, crosses, and triangles) within the two highest-volume hospitals in our data in 2008–2011. Each point represents a unique price paid for a lower-limb MRI in a given hospital-month, where the size of the point corresponds to the volume of MRIs paid at that price. Repeated prices are linked across renegotiation events using information on the plan characteristics of the patients whose episodes were paid at that price. For more information on the methods used to link contracted prices see Online Appendix B.3.

lower-limb MRIs from 2008 to 2011 at the two highest-volume hospitals in our data. Each point is an exact price paid for a case; the size of the dots is proportional to the number of patient cases at that price (exactly to the cent).[37] We highlight the three highest-volume contracts at each hospital (these capture 92% and 98% of all lower-limb MRI cases at these hospitals, respectively). The figure clearly demonstrates that there are significant differences in MRI prices within hospitals at single points in time. For example, in January 2011, the ratio of the price of the highest volume contract (circles) to the price of the second highest volume contract (triangles) is 1.39 at Hospital A and 1.65 at Hospital B. We also

37. We present these amounts as dollars from the hospital mean to remain consistent with publishing rules in our data use agreement. The hospital mean is fixed across all time periods, so a flat line reflects an unchanging absolute price.

see that the main contract prices are stable for extended periods (usually one year) before being updated, although the updates occur at different times across contracts.

The analysis in this section provides the first national evidence that insurers pay substantially different prices for the same services at the same hospitals. This finding is consistent with insurer-hospital bargaining models of price determination where stronger insurers can negotiate lower prices.[38]

V. Analysis of Insurer-Hospital Contracts

V.A. *Types of Insurer-Hospital Contracts*

When a hospital joins an insurer's network, the hospital signs a contract that stipulates how and what they will be paid. Unfortunately, because most of these contracts contain clauses that prohibit their terms from being released, little is known about precisely how insurers pay each hospital (Reinhardt 2006; Gaynor and Town 2011). However, in addition to analyzing price levels, the richness of the HCCI data also enables us to estimate the *types* of insurer-hospital contracts that are being struck.

In general, there are two main ways hospitals are paid for inpatient services (Moody's Investors Service 2017). The first is using prospectively set prices that pay a fixed dollar amount based on the patient's DRG (or sometimes a more disaggregated coding framework like ICD-9 codes). The second method sets payments as a percent of hospital charges, which we call a share of charges contract. Note that there are also hybrid payments that blend elements of both payment types. These hybrid payments are prospective payment contracts that include outlier adjustments that allow hospitals to be paid more when costs for a particular case are significantly higher than average costs.[39] Furthermore, within the class of prospective payment contracts, some may have their payment levels set as a percentage of Medicare payments, whereas

38. See for example Town and Vistnes (2001), Capps, Dranove, and Satterthwaite (2003), Sorensen (2003), Farrell et al. (2011), Gowrisankaran, Nevo, and Town (2015), and Ho and Lee (2017).

39. There is another type of contract that has been used historically where some inpatient payments were made on a per diem basis. However, our data contributors report that virtually none of the cases in our data are paid on a per diem basis. They also report that they aim to have less than 5% of cases subject to outlier adjustments.

others will have payment levels independent of the Medicare fee schedule.

There are two main reasons hospitals are likely to prefer share of charges contracts to prospective payment contracts.[40] First, hospitals bear less risk with share of charges contracts. With this type of contract, a hospital gets paid for every service they provide to a patient. As a consequence, if a patient (in a particular DRG, for example) requires more care and is therefore more expensive, the hospital gets paid more and the insurer bears this additional cost. Of course, if the patient requires fewer services and is thus cheaper, the hospital receives less payment. By contrast, under a prospective payment the amount a hospital will receive is fixed ex ante. As a consequence, the hospital bears the risk associated with uncertainty over the cost of treatment (Burns and Pauly 2018). With risk aversion, this uncertainty is unattractive (Ellis and McGuire 1988; Town, Feldman, and Kralewski 2011). A second reason hospitals prefer share of charge contracts is that it places them under less pressure to reduce costs, since they get paid for all the services provided (presuming that the prices at least cover hospitals' marginal costs of providing services). As a result, prospective payments give stronger incentives for the hospital to contain costs (Shleifer 1985).

From our discussions with insurers, it seems that when prospective payment contracts exist, insurers will often offer a simple standardized boilerplate contract tied to the Medicare fee schedule (i.e., prospective payments at a fixed percentage of Medicare payments). This saves them the costs of negotiating with each hospital. The patient profile in a hospital may mean true costs depart significantly from Medicare reimbursement. However, it may be difficult for a hospital to credibly demonstrate this to an insurer due to asymmetric information, even if a deviation from the boilerplate contract were worthwhile for both parties (net of negotiating costs). Hence, hospitals with high bargaining power may be able to move away from the insurer's standard Medicare-related prospective scheme, but it will be harder for a weaker hospital to persuade an insurer to do this.

These considerations suggest that the differential bargaining power of hospitals and insurers will affect not only the hospital

40. See Newhouse (1996) for a more general discussion of contract form and trade-offs. Basically, share of charge contracts are like cost-plus contracts and prospective payments are like fixed-price contracts.

price level but also the form of the contract. In particular, we expect that hospitals with greater bargaining power will have more share of charge contracts and, if they have prospective contracts, a lower share of them will be tied to Medicare reimbursement. Before examining this hypothesis in the next section, we turn to how we identify contract types and provide some basic descriptive statistics.

V.B. Estimating the Percentage of Cases Paid as a Share of Hospital Charges

Online Appendix B.3 details exactly how we classify contracts, but we sketch the method here. The HCCI data do not specify whether cases are paid prospectively, as a share of charges, or using a hybrid payment. As a result, we developed a strategy to identify how cases were paid. To do this, we group separate claims within hospitals for a procedure (e.g., knee replacement) into single contracts if cases are paid at identical dollar amounts (down to the cent) or paid at identical percentages of hospital charges (down to the hundredth of a percent).[41] We categorize hospital payments as either (i) share of charges (contracts where two or more cases are paid at an identical percentage of hospital charges), (ii) prospective payments (two or more cases are paid at identical dollar amounts), or (iii) unclassified cases. Unclassified cases are a mix between those using one of the hybrid contracts (e.g., those involving outlier payments) and others that fall under one of the main two contract classes but where the data is not rich enough to identify which one. The latter occurs, for example, when we only observe one case under a contract so we cannot "price match" it to another case.

We find evidence that even within a month a hospital can have prospective payments with one payer and a share of charge contract with another for the same procedure. To illustrate this, consider Figure VI. Here we group cases into contracts for vaginal delivery at a large hospital using the methods described above. Two insurer contracts are clearly visible—contract 1 is shown in

41. Our approach to identifying contracts is similar to the bunching analysis that Clemens, Gottlieb, and Molnar (2017) use to study physician pricing. We identify cases that are paid as a repeated percentage of hospitals' charges or as a repeated dollar amount. For more discussion of how we identify contracts, see Online Appendix B.3.

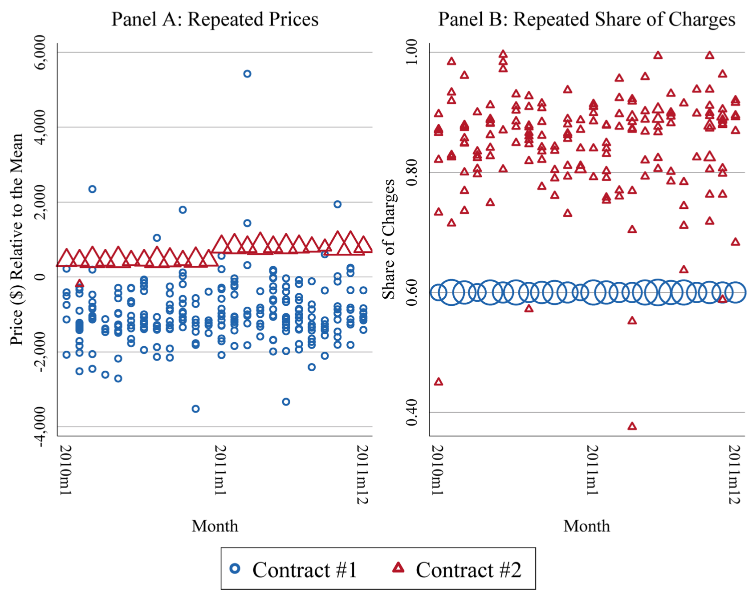


FIGURE VI

Repeated Price and Share of Charge Agreements at a Hospital for Vaginal Delivery, 2010–2011

These figures highlight the top two linked contracts within a high volume hospital for 2010–2011. Circles represent contract 1; triangles represent contract 2. The size of the point corresponds to the volume of cases at that price. Repeated prices and price-to-charge ratios are linked across renegotiation events using information on the plan characteristics of the patients whose episodes were paid at that price or rate. For more information on the methods used to link contracted prices see Online Appendix B.3. In Panel A the prices on the y-axis are relative to the average hospital price over the entire period which is constant across all observations (in order to avoid revealing a particular price).

circles and contract 2 in triangles in both panels.[42] In Panel A, we plot the contracted prices in dollars from the mean price at that hospital. As can be seen, there is one absolute dollar amount for contract 2, but there is significant heterogeneity in the dollar amounts paid for contract 1. Contract 2 is paid using a prospective payment set at a fixed payment amount, whereas the payment amounts for contract 1 clearly vary. In Panel B, we plot all of these

42. To make it easier to visualize, we only show the two highest-volume contracts at this hospital.

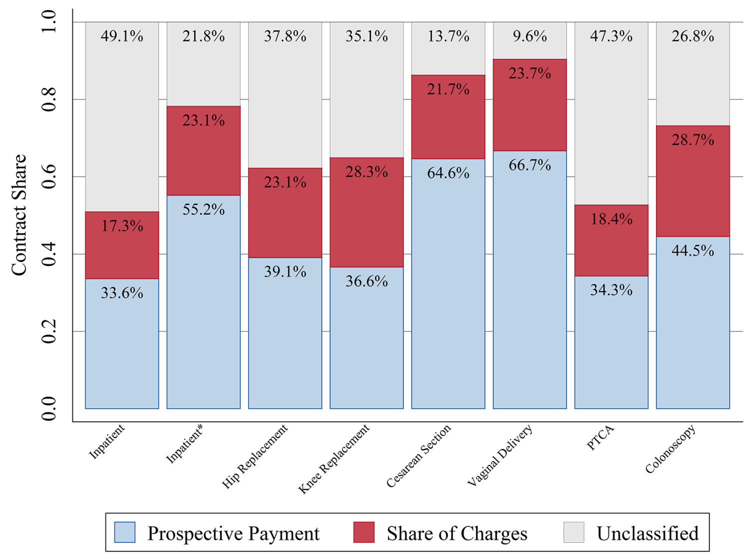


FIGURE VII

Contract Classifications Overall and by Procedure, 2010–2011

The bars present the share of the claims by procedure (or inpatient sample) classified into each type of contract using case-level data from 2010 to 2011. The bottom bars display the percent of cases classified as prospective payments. The middle bars display the percent of cases paid as a share of charges. The top bars display the percent of cases not classified. The numbers of hospitals (cases) underlying each bar are 2,253 (2,288,907) for the inpatient sample, 404 (15,122) for hip replacement, 809 (37,157) for knee replacement, 1,041 (81,482) for cesarean section, 1,136 (108,794) for vaginal delivery, 501 (16,636) for PTCA, and 1,008 (66,018) for colonoscopy. Inpatient* presents a restricted subsample of the inpatient cases for hospital-DRG pairs that represent at least 20 admissions from 2010 to 2011. This sample represents 1,841 hospitals and 1,078,697 admissions.

payments as a percent of the hospital's charges. What is clear is that contract 1 is paid at a constant percent of charges (60%). For contract 2, the percent of charges varies in this panel because although the absolute price is constant, the precise charges vary for each case.

In Figure VII we show the breakdown of cases for the inpatient sample (first two bars) and procedure sample (other bars). Among inpatient cases, about a third are on prospective payments contracts and 17% are share of charge contracts. Almost half were unclassified, but when we restrict our sample to hospital-DRG pairs in higher volumes, we see a big reduction in unclassified cases. For example, in the second bar, we restrict to hospital-DRG pairs that have at least 20 admissions and observe that 22% of cases are unclassified. This is because the more cases a hospital treats, the higher the likelihood we correctly identify two cases paid at the same constant rate. As Online Appendix Figure VIII details, as we alter count restrictions, we maintain a robust estimate of about 23% of all cases being share of charge payments. There is a little more uncertainty about the exact proportion of cases on prospective payments, but we know the upper bound is 77% ($= 100 - 23$), and Online Appendix B.3 suggests that the true fraction is not far from this level.[43]

We also observe large variation in the fraction of share of charge contracts across hospitals and across procedures (see Online Appendix Figure IX). For vaginal deliveries (our highest volume service with the lowest fraction of unclassified cases), the hospital in the 90th percentile has 91% of cases paid as a share of charges, whereas the 10th percentile has zero. It may seem surprising that a single hospital has multiple forms of contracts given their patient mix. The fact that they do is consistent with the idea that different insurers have different degrees of bargaining power within a single hospital.[44]

There have been, to our knowledge, only two other attempts to identify hospital-insurer contracts, both trying to reverse engineer contracts from price (as we do here). Baker et al. (2016) estimate that around three-quarters of inpatient payments were paid prospectively (see Online Appendix B.3 for details). Gift, Arnould, and DeBrock (2002) examined hospital contracts from

43. The proportion of cases classified as prospective payments rises (and the proportion unclassified falls) almost monotonically with the minimum case threshold. For example, the proportion of cases classified as prospective rises from 55% at a threshold of 20 cases to 72% at a threshold of 200 cases. Note that for the procedures (with zero minimum case threshold restrictions), estimates range from 18% of cases on a share of charge contract for PTCA up to 30% for colonoscopies. Because nearly all lower-limb MRIs in our data have identical charges inside facilities, we cannot differentiate between cases paid prospectively and those paid as a share of hospital charges.

44. In Online Appendix Figure X, we plot ln(prices) on the y-axis against ln(charges) on the x-axis for the same DRG for cases paid as a share of charges at a large hospital in our data. It shows that there tends to be a single share of charges per contract applied across all DRGs. In other words, an insurer will tend to negotiate the same level of discount off charges for all DRGs in the same hospital.

a single insurer with hospitals in Washington state in financial year 1994/1995 and found only 41% of the contracts had prospective payment contracts. We are able to extend beyond these papers by the ability to differentiate between cases paid prospectively and those paid as a share of charges and show the existence of different contracts within the same hospital. As we describe in the next subsection, we are also able to analyze whether prospectively paid cases have payments set as a percentage of Medicare payments. This allows us to extend work by Clemens, Gottlieb, and Molnar (2017) and Clemens and Gottlieb (2017) on physicians and analyze the relationship between hospitals' prices and Medicare payments.

V.C. Prospective Payment Contracts and Their Link to Medicare Hospital Payments

To estimate the share of prospective cases tied to Medicare, we calculate each prospective price as a percentage of the Medicare PPS payment rates. We identify other private cases with different DRGs at the same hospital that are paid at the same percentage of Medicare PPS rates. These cases are then grouped into contracts. We calculate the share of a hospital's prospectively set inpatient cases that have another case of a different DRG that is paid at the same percentage of Medicare payment rates (down to the hundredth of a percent). We find that among all inpatient prospective payments, 74% are set as a share of Medicare rates. There is significant heterogeneity across hospitals—the unweighted mean is 48% with a standard deviation of 32.

To illustrate this heterogeneity, in Figure VIII we plot ln(prospective payments) on the y-axis against ln(Medicare payments) on the x-axis for the same DRG at two large hospitals in our data. Each circle is a unique case that we have classified as being under a prospective payment contract for a specific DRG. If hospitals were paid a fixed percentage of Medicare payment rates, the points on the graph would have a slope of 1.[45] Indeed, we observe that the private payment rates for the hospital in Panel A, for example, are predominantly set as a percentage of Medicare

45. To formalize this point, when the price P paid at hospital h, for DRG d, for an admission that occurs at time t, is set as a percentage of the DRG-specific Medicare rate M, assume it takes the form of a percentage markup $\Theta_{h,t}$ over Medicare payments: $P_{h,d,t} = \Theta_{h,t} * M_{d,t}$. Thus, $\ln(P_{h,d,t})$ is additively separable: $\ln(P_{h,d,t}) = \ln(\Theta_{h,t}) + \ln(M_{d,t\cdot})$.

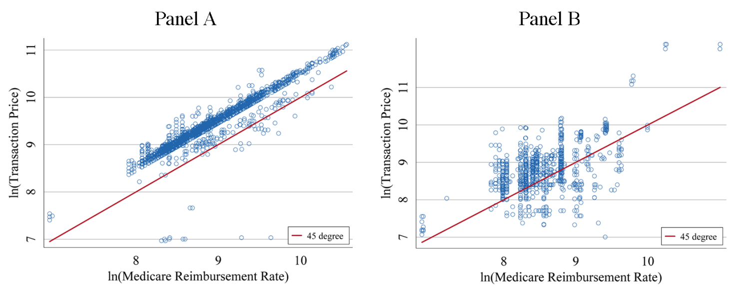


FIGURE VIII

Medicare Reimbursements and Transaction Prices at Two High-Volume Hospitals, 2011

The panels represent two large hospitals in the data. Each circle is a unique, privately paid prospective-payment amount for a DRG (y-axis). The x-axis is the corresponding logged Medicare reimbursement rates based on 2011 data. The diagonal line is the 45° line.

rates (they parallel the 45° line). By contrast, the payment rates at the hospital in Panel B are not highly correlated with Medicare rates.

When we look across all inpatient cases in our data, our results suggest the share of hospitals' private prospective payments that are linked to Medicare is likely to be lower than the 75% estimate Clemens, Gottlieb, and Molnar (2017) observed among physicians. First, about 23% of cases are share of charge payments, which are therefore directly not linked to Medicare. Second, since no more than 77% of cases are paid prospectively and 74% of prospective cases are linked to Medicare, this implies that the upper bound for total cases linked to Medicare payment levels is 57% (= 77 * 0.74).

VI. FACTORS ASSOCIATED WITH HOSPITAL PRICES AND CONTRACT TYPES

VI.A. *Cross-Sectional Analysis of Hospital Prices and Contracting Type*

We have identified substantial differences across hospitals in their prices and contract structures, and we now turn to identifying the factors associated with these differences. Prices and contract forms are determined by negotiations between hospitals and insurers, and a number of factors may affect the outcomes of these negotiations. These include demand shifters (e.g., hospital quality), supply shifters (e.g., labor costs), and the respective bargaining leverage of insurers and hospitals.

We begin by examining the cross-sectional relationship between hospital and insurer market structure and hospital prices and contracts. To do so, we use the following estimating equation:

$$y_{h,t} = \alpha' M_{h,t} + \gamma' x_{h,t} + \tau_t + \upsilon_{h,t}, \tag{2}$$

where $M_{h,t}$ is a vector of measures of hospital and insurer market structure for hospital h in year t, $x_{h,t}$ is a vector of control variables (described below), τ_t are year dummies, and $\upsilon_{h,t}$ is the error term. The $y_{h,t}$ outcomes we consider are (i) the inpatient hospital price index ($\hat{p}_{h,t}$) described above and in Online Appendix B.1; (ii) our procedure-level prices described in Online Appendix B.2; (iii) the percent of cases paid as a share of the hospital's charges described in Section V.B; and (iv) the percent of prospective payments linked to the Medicare fee schedule described in Section V.C.

We construct several measures of market structure. Our main measure of hospital market structure is made by drawing a circular area with a radius of 15 miles around each hospital. We label hospitals in these areas that do not have competitors as monopolies; those in areas with two hospitals as duopolies; and those in areas with three hospitals as triopolies. Our omitted base category is hospitals in areas with four or more hospitals (i.e., quadropolies or greater). We also show that our main results are robust to a large range of alternatively defined measures of hospital market structure, such as measures with alternative market size definitions (e.g., fixed-distance radii of various distances) and alternative measures of market structure (e.g., counts of hospitals and Herfindahl–Hirschman Indexes [HHIs]). Our main measure of insurer market structure is the HCCI data contributors' market share of privately insured lives at the county level. Further details of how our market structure measures are constructed are contained in Online Appendix C. We present correlates of our hospital concentration measures and key covariates in Online Appendix Figure XI. These concentration measures are not strongly associated with other covariates, such as hospital quality or average population characteristics, although we do find that rural areas have more concentrated hospital markets.

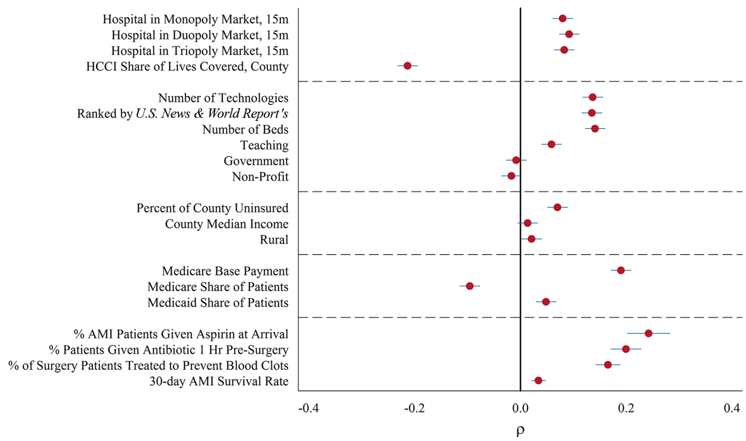


FIGURE IX

Bivariate Correlations of Hospital Price with Observable Factors, 2008–2011

The *x*-axis reflects the level of the bivariate correlations between key variables featured in our regressions and hospitals' regression-adjusted inpatient prices that are risk-adjusted for DRG, age, and sex. The bars show the 95% confidence intervals surrounding the correlations. Because these are bivariate correlations, "duopoly" is duopoly or monopoly and the implicit omitted category is triopoly or greater. "Triopoly" is triopoly, duopoly, or monopoly. For government and nonprofit, the omitted category is private for-profit hospital.

We begin by examining the bivariate correlations between our hospital inpatient price index and other key variables in Figure IX. Relative to hospitals in markets with four or more competitors, hospitals in markets with fewer competitors have significantly higher prices. By contrast, prices are considerably lower at hospitals in counties where HCCI insurers have a higher market share. Apart from market structure, the other covariates are generally of the expected signs. Hospitals using more technologies, teaching hospitals, and larger hospitals (number of beds) have higher prices. Nonprofit and government hospitals have slightly lower prices than for-profit hospitals. Hospitals with higher quality measured either by a mention in *U.S. News & World Report* or via process scores tend to have higher prices.[46]

46. These process scores are the percentage of AMI patients given aspirin at arrival, the percentage of patients given an antibiotic before surgery, and the percentage of patients treated to prevent blood clots. The sole exception is hospitals' 30-day AMI survival rate, which is negatively correlated with hospital prices.

Hospitals with higher Medicare base payment rates or those located in high-income counties have higher prices, consistent with these being high-cost areas.[47] The higher the share of Medicare patients a hospital treats, the lower its private prices. Counties with more uninsured individuals also have higher prices.

VI.B. Cross-Sectional Analysis of Hospital Prices and Contract Form

In Table IV, we present estimates of equation (2) and report the coefficients on the market structure variables where an observation is a hospital-year (full results with coefficients on the other covariates are reported in Online Appendix Tables X–XII). In Panel A the dependent variable is the inpatient price index, in Panel B it is the percent of each hospital's inpatient cases paid as a share of charges, and in Panel C it is the percent of prospective payments paid as a percentage of Medicare payments.

Table IV, Panel A shows that there is a significant and positive association between hospital price and whether a hospital is located in a monopoly, duopoly, or triopoly market. Conversely, hospital prices fall as the HCCI insurers' market share increases. Column (1) presents the simplest specification, column (2) adds insurer market share, and column (3) further adds HRR fixed effects, so the coefficients are identified from the variation in market structure within HRRs. Introducing HRR fixed effects reduces all the hospital concentration coefficients, but with the exception of the triopoly dummy, all coefficients remain significant at conventional levels. The coefficients in column (3) indicate that monopoly hospitals are associated with prices that are 12.5% $(= e^{0.118} - 1)$ higher than places where there are four or more hospitals. Duopolies are associated with 7.6% higher prices. Furthermore, a 10 percentage point increase in the market share of the HCCI insurers (i.e., a one standard deviation increase) is associated with a statistically significant 7% fall in hospital prices. Note that the hospital market structure indicators are quantitatively the most important variables in our cross-sectional price analysis. Our hospital market structure indicators capture 19.6% of the explained variance from estimates presented in Table IV,

47. By contrast, the higher the percentage of Medicaid patients a hospital treats, the higher its prices. However, this is the only coefficient which is significantly reversed in our multivariate regression estimates of equation (2)—see Online Appendix Table X.

TABLE IV

Hospital Concentration, Prices, and Contract Form, 2008–2011

	(1)	(2)	(3)
Panel A: ln(hospital price); mean = 9.42, obs = 8,772, number of hospitals = 2,358			
Monopoly	0.234***	0.190***	0.118***
	(0.024)	(0.024)	(0.024)
Duopoly	0.161***	0.130***	0.073***
	(0.021)	(0.020)	(0.024)
Triopoly	0.115***	0.083***	0.036
	(0.023)	(0.023)	(0.023)
HCCI market share		− 0.006***	− 0.007***
		(0.002)	(0.002)
Panel B: Percent of cases paid as share of charges; mean = 18.6%, obs = 4,344, number of hospitals = 2,253			
Monopoly	17.335***	15.241***	10.455***
	(1.828)	(1.823)	(1.778)
Duopoly	9.979***	8.424***	5.702***
	(1.760)	(1.740)	(1.596)
Triopoly	7.804***	6.235**	4.909**
	(1.909)	(1.938)	(1.608)
HCCI market share		− 0.288***	− 0.403***
		(0.077)	(0.120)
Panel C: Percent of cases of prospective payments tied to Medicare; mean = 48.3%, obs = 3,669, number of hospitals = 1,936			
Monopoly	− 16.849***	− 11.275***	− 11.293***
	(2.882)	(2.696)	(3.160)
Duopoly	− 8.791***	− 4.272*	− 5.595**
	(2.441)	(2.443)	(2.316)
Triopoly	− 7.111**	− 2.422	− 5.747**
	(2.866)	(2.727)	(2.790)
HCCI market share		0.890***	0.616***
		(0.091)	(0.174)
HRR fixed effects	No	No	Yes

Notes. * $p < .10$, ** $p < .05$, *** $p < .01$. OLS estimates of equation (2) with standard errors clustered at the HRR level in parentheses. Market structure variables are described in Online Appendix C. The dependent variable in Panel A is ln(hospital inpatient prices) that are regression risk-adjusted for DRG, age, and sex; in Panel B the dependent variable is the percent of cases paid as share of charges (i.e., nonprospective payments); in Panel C the dependent variable is the percent of cases tied to the Medicare reimbursement rate. An observation is a hospital-year. In Panel A, the data covers 2008 to 2011; in Panels B and C it covers 2010 to 2011 because charge data are unavailable for earlier years. All regressions include controls for the number of technologies, a dummy for being ranked in *U.S. News & World Reports*, size (number of beds), hospital ownership (government, nonprofit, or for-profit), whether a teaching hospital, % of county uninsured, county median income, the Medicare payment rate, share of Medicare, share of Medicaid, and year dummies. Online Appendix Tables X–XII report full sets of results. Hospitals in quadropoly or greater markets are the omitted category.

Panel A, column (2) (when we exclude these measures, the R^2 drops from 0.170 to 0.137). The market share of the HCCI insurers captures the second highest share of the explained variance, with an associated decrease in R^2 from 0.170 to 0.143. No other variables in the analysis capture more than 10% of the explained price variance.

The results in Table IV, Panel A are robust to measuring prices in a multitude of ways, such as (i) risk-adjusting our inpatient price measure with patients' Charlson score; (ii) risk-adjusting our inpatient price using ICD-9 diagnosis codes instead of DRG fixed effects (about 9,235 ICD-9 codes versus 746 DRG codes), and measuring price in levels instead of logarithms (see Online Appendix Table XIII).[48] Our results are consistent with earlier, single-state studies of hospital prices and market structure (mostly using data from California), which have found strong positive and statistically significant correlation between hospital market concentration and prices (see Vogt and Town 2006; Gaynor and Town 2012).

Table IV, Panel B has the same specification as in Panel A but changes the dependent variable to the percent of cases paid as a share of hospital charges.[49] Because data on charges are only available in 2010 and 2011, the sample size roughly halves. Across the various specifications, we consistently find that the share of inpatient cases paid as a share of charges declines monotonically as the number of rival hospitals per market increases. Focusing on the estimates from column (3), we find that a monopoly hospital has 10.5 percentage points more cases paid as a percent of charges than do hospitals in areas with four or more hospitals (over a mean of 18.6%). Hospitals in counties where the HCCI insurers have a larger market share have significantly lower rates of cases paid as a share of charges (a 10 percentage point increase in the HCCI share is associated with a 4% lower share of cases on these contracts).

One might be concerned that the coefficient on monopoly in the price regressions of Table IV, Panel A reflects some form of

48. For example, when we use prices in levels as the dependent variable instead of logarithms in Table IV, Panel A, we obtain a coefficient on the monopoly indicator of 1,605 in the equivalent of column (3). Since the average inpatient case's cost is $14,020, this estimate implies an effect of 12%, nearly identical to the baseline estimate. This is reported in Online Appendix Table XIII.

49. The bivariate correlations are illustrated in Online Appendix Figure XI.

prospective contract where the hospital obtains a higher price because it is bearing more risk than the insurer. For example, perhaps there are more patients with unobservable idiosyncratic costs in places with concentrated hospital markets, which (under a prospective pay contract) would leave hospitals bearing more financial risk. The fact that monopoly hospitals receive both higher prices and have a disproportionately larger share of price contracts (where insurers bear more of the risk) is inconsistent with this explanation.[50]

Table IV, Panel C uses the share of prospective payments that are tied to Medicare payment levels as the dependent variable.[51] The pattern is familiar: hospitals in markets with fewer potential competitors have significantly fewer cases paid as a percent of the Medicare payments. In column (3), monopoly hospitals are associated with having 11.3 percentage points fewer cases on contracts of this type (over a mean of 48%). We also find that hospitals in areas where the HCCI insurers have bigger market shares have a higher share of their cases paid based on the Medicare fee schedule (a 10 percentage point increase in insurer share is associated with 6% more Medicare-linked contracts).

The results in Table IV paint a consistent picture of bargaining power. At least descriptively, when hospital markets are concentrated (and/or insurer markets are fragmented), hospital prices are higher and hospitals are able to obtain contracts that shift more risk on to insurers.

VI.C. Results for Individual Procedures

A concern with the regressions in Table IV is that because we aggregate over many different procedures, we may fail to account for unobserved heterogeneity in hospitals' care. For example, prices in monopoly hospitals may be higher because their procedures are more complex and costly, even after we risk-adjust.

50. If we control for contract type on the right-hand side of the price regressions the coefficient on monopoly falls by about a tenth which implies that monopolies have higher prices even on the same type of contract. To investigate this we ran a case-level price regression on 2010 and 2011 data (where we have charge data) analogously to column (3) of Table IV Panel A where we include a dummy reflecting whether the case is paid as a share of charges or not. Without this control the coefficient on monopoly was 0.137, but with the control the coefficient falls to 0.125.

51. Bivariate correlations are in Online Appendix Figure XII.

Consequently, in Online Appendix Table XIV we reestimated the models of Table IV using our seven procedures.[52] In column (1) we reproduce the baseline inpatient estimates in the final column of Table IV. Looking across the different procedures, it is striking that despite the smaller sample sizes, the results are qualitatively very consistent with the overall inpatient results. For all procedures, we find that areas with a monopoly hospital have higher prices than those with four or more hospitals. This positive association is significant at the 5% level for all procedures except hip replacements and PTCA (which have our smallest sample size) and colonoscopy (significant at the 10% level). The coefficients imply that a hospital located in a monopoly market has prices that are between 5.5% (hip replacements in column (3)) and 23.4% (lower-limb MRIs in column (9)) higher than hospitals in markets with four or more hospitals. The coefficient on the HCCI insurer market share is less precisely estimated, but it is negative for all procedures except cesarean sections and hip replacements. Column (2) summarizes the effects by pooling across all the procedures in columns (3)–(9) and adding a dummy variable for each procedure. The pooled results confirm that hospitals facing fewer potential competitors have significantly higher prices.[53]

In Online Appendix Table XIV, Panel B, we perform the same exercise for each procedure sample but use the percent of cases paid as a share of charges as the dependent variable. We again find that hospitals with fewer potential competitors have a higher proportion of their cases paid as a share of charges. As with price, we find that hospital concentration is positively associated with the percentage of cases paid as a share of charges for all procedures and is significant for all procedures except hip replacements and PTCA (which have the smallest samples). The coefficient on HCCI insurer share is negative for five of the six procedures. There is almost no variation in hospital charges for MRIs within a facility, so we cannot estimate the structure of contracts for this

52. See Online Appendix B.2 for construction of these prices. Note that we cannot perform an analysis of the share of prospective payments tied to Medicare at the procedure level because the variable is constructed by linking payment rates across procedures (DRGs), and thus does not exist for any specific procedure.

53. As hospitals increasingly purchase physician groups, there may be concerns that some portion of physician fees show up in facility prices. Consequently, we reestimate our analysis using prices measured as the sum of hospital and physician prices in each claim (see Online Appendix Table XV). The results are qualitatively similar to what we observe in our main specifications.

procedure. When we pool our procedures into a single estimate, we confirm a positive association between hospital market concentration and the fraction of cases paid as a share of charges. We also find that HCCI insurer market share is negatively and significantly associated with the fraction of cases paid as a share of charges.

VI.D. Robustness of Cross-Sectional Analysis

We conducted a large number of robustness tests on the results in Table IV, some of which we describe here. First, the main cross-sectional estimates are robust when we use alternatively constructed measures of hospital market structure, such as continuous or binned HHIs, allowing many alternatively sized radii to define markets, and/or allowing differential market definitions in rural and urban areas. Likewise, our results are also robust to different measures of insurer market structure.[54] Second, our pricing analysis could be sensitive to omitted quality if, in particular, quality is correlated with market structure. Consequently, we include four additional measures of clinical quality to the price regression. Consistent with Figure IX, three of the four measures are correctly signed, but the coefficients on market structure were largely unchanged. We also included all 41 measures of quality published by Medicare Hospital Compare in a regression, which again did not meaningfully shift the hospital market structure coefficients. Third, we show that our results are not driven exclusively by extremes by dropping observations from monopolies or hospitals in markets with six or more providers. Fourth, we show that our results are not sensitive to the exact sample size cutoffs we use (e.g., hospitals must perform at least 50 cases a year to be in the inpatient sample) by showing results where we use many alternative cutoffs from between 0 to 100 cases a year.[55]

Finally, as discussed previously, we do not have data from BCBS plans. If hospital market structure is correlated with

54. For example, the coefficients on our main hospital market structure measures are broadly unchanged when we include cubic polynomials of the market shares of the three HCCI contributors and/or individual shares of the top 10 insurers in each market.

55. The analysis of alternative market structure is in Online Appendix Tables XVI–XVIII; quality in Online Appendix Table XIX; extreme market structures in Online Appendix Table XX; and alternative cutoffs in Online Appendix Table XXI.

omitted BCBS presence, this could present a problem. Online Appendix F conducts an extensive analysis of this and does not find it to be a major issue. First, note that the correlation between hospital HHI and the county-level BCBS market share is only 0.222. Second, we estimated all our models solely in areas with high (above-median) and low (below-median) BCBS market shares. Although the exact magnitudes of some of our coefficients differ in areas where BCBS have high and low market share, our main finding that having fewer hospitals in a market is associated with higher prices, a higher proportion of cases paid as a share of hospital charges, and a lower fraction of prospectively paid cases paid as a share of Medicare rates remains robust.[56]

VII. Hospital Mergers

VII.A. Introduction to our Merger Analysis

Our cross-sectional regressions in the previous section suggest that hospital market structure is strongly associated with hospital prices. Here we analyze mergers and hospital prices using the panel aspect of our data. Over the past few decades, there have been hundreds of mergers between hospitals across the United States. Economic models of competition in the hospital sector predict that mergers between hospitals that are close geographic competitors will lead to price increases, making mergers of direct interest (see the Gaynor, Ho, and Town 2015 review). Furthermore, examining the impact of mergers on hospital prices provides another lens through which to view the relationship between market structure and prices, and complements our cross-sectional analysis.

A number of papers have estimated the impacts of specific mergers that were suspected to be anticompetitive. One strand of this literature uses estimates from structural (or semi-structural) models and ex ante simulation methods to generate estimates of predicted price changes from a single or a small number of transactions.[57] Although these models allow for a more

56. As we discuss in more detail in Online Appendix F, it becomes difficult to precisely estimate the impacts of the market structure variables in areas with high BCBS share when HRR fixed effects are included because very few of those HRRs have monopoly hospitals and hospitals facing four or more competitors that meet our sample restrictions.

57. See Town and Vistnes (2001), Capps, Dranove, and Satterthwaite (2003), Gaynor and Vogt (2003), and Gowrisankaran, Nevo, and Town (2015).

sophisticated modeling approach to competition and bargaining between insurers and hospitals, they would be difficult to estimate for the hundreds of mergers we have in our data. Instead, we follow a second strand of the literature that uses ex post econometric methodologies to analyze the effects of consummated mergers.[58] This kind of modeling is coarser but has the advantage of looking at what happens after mergers occur. Historically, this strand of the literature has also focused on analyzing individual mergers or small numbers of mergers. We extend the literature by examining the impact of hospital mergers that occurred in the United States during the five years covered by our data (which is also a more recent time period than covered in previous studies).

VII.B. Hospital Merger Data

We created a database of nearly all U.S. hospital mergers between 2007 and 2011 (see Online Appendix D for details) and found 366 transactions involving more than 2,000 hospitals. For example, as Online Appendix Table XXII shows, there were 55 transactions involving 84 hospitals where the merging parties were less than 5 miles apart and 121 transactions involving 260 hospitals within 15 miles of each other.

VII.C. Modeling Hospital Mergers

To estimate the effects of mergers on hospital prices, we use the following specification:

$$\ln(\hat{p}_{h,t}) = \beta MERGE^{D}_{h,t} + \eta_h + \delta_t + \upsilon_{h,t}, \tag{3}$$

where $\hat{p}_{h,t}$ is the usual risk-adjusted hospital inpatient price for hospital h in year t. We include hospital fixed effects (η_{h}) and year dummies (δ_{t}). The key variable of interest is the binary indicator, $MERGE^{D}_{h,t}$. In our baseline specification, this indicator is 0 until the year a hospital becomes involved in a merger, when it then takes a value of 1 and retains a value of 1 for the remainder of our sample period. We categorize mergers based on the physical distance (superscript D) between the merging entities (whether the merging parties were separated by 5 miles or less, 10 miles or less, etc.). Since hospital location is a key factor

58. See Vita and Sacher (2001), Krishnan (2001), Capps and Dranove (2004), Dafny (2009), Kemp, Kersten, and Severijnen (2012), Haas-Wilson and Garmon (2011), Tenn (2011), and Thompson (2011).

determining demand (and hence potential patient substitutability between hospitals), we expect mergers between hospitals that are geographically closer to result in larger increases in prices than mergers between hospitals separated by large distances.[59] We use a variety of different control groups, including all hospitals not involved in mergers and matched controls using a number of different matching methods. In some specifications we also include the same set of control variables included in our cross-sectional regressions in Table IV.[60]

There are differences in the characteristics of the merging versus nonmerging hospitals (see Online Appendix Table XXIII). Merging hospitals tend to be located in less concentrated markets (this is unsurprising due to antitrust scrutiny and a mechanical limit to how concentrated a market can get), are more likely to be nonprofit and teaching hospitals, are larger (more beds), and have higher reputational average quality (*U.S. News & World Report* quality rankings). However, merging and nonmerging hospitals look broadly comparable in terms of their share of Medicare and Medicaid admissions, the technologies they possess, and their area characteristics (county uninsured and median income). Most of these characteristics vary little over time so the hospital fixed effects in equation (3) will largely control for them. More important, as we demonstrate below, we do not find any evidence that merging hospitals have different premerger trends in prices relative to nonmerging hospitals.

VII.D. Results on Mergers and Hospital Prices

Table V, Panel A contains the baseline specifications where we vary the distance between merging hospitals from 5 to 50 miles. There are positive coefficients on the merger dummies at every distance, and these are almost all significant for mergers

59. We recognize that mergers between hospitals farther apart may have impacts on prices through more subtle forms of multimarket conduct behavior. Our specification flexibly allows for mergers to have impacts at any distance, although we are not testing specifically for cross-market merger effects like those analyzed by Dafny, Ho, and Lee (2016) and Lewis and Pflum (2017).

60. Because the Department of Justice and the Federal Trade Commission occasionally allow failing or "flailing" firms to merge, we want to exclude these firms from our analysis. To do that, we exclude 53 hospitals that have the largest share of unused capacity defined as the average daily census divided by the total number of hospital beds (e.g., those in the 99th percentile of unused capacity). Our results are robust to including these 53 hospitals in our analysis.

TABLE V
HOSPITAL PRICES AND MERGERS.

Distance (merger within given number of miles):		5	10	15	20	25	30	50
Dependent variable: ln(price)		(1)	(2)	(3)	(4)	(5)	(6)	(7)
Panel A: Baseline								
	Post merger	0.060**	0.039**	0.021	0.023*	0.024**	0.014	0.008
		(0.025)	(0.019)	(0.013)	(0.013)	(0.011)	(0.011)	(0.009)
Panel B: Add controls								
	Post merger	0.062**	0.040**	0.021	0.024*	0.024**	0.014	0.008
		(0.025)	(0.019)	(0.013)	(0.013)	(0.011)	(0.010)	(0.009)
Panel C: Separately controlling for neighbors								
	Post merger	0.062**	0.040**	0.021	0.022*	0.024**	0.013	0.008
		(0.025)	(0.019)	(0.013)	(0.013)	(0.011)	(0.011)	(0.009)
	Merging neighbor	−0.016	0.024*	0.005	0.013	0.003	0.003	−0.005
		(0.028)	(0.014)	(0.010)	(0.008)	(0.007)	(0.007)	(0.006)
Panel D: Merger effects over time ($t - 2$ and before omitted base)								
	One year prior to merger	0.019	0.015	0.015	0.015	0.015	0.012	0.013
		(0.03)	(0.018)	(0.013)	(0.012)	(0.011)	(0.012)	(0.010)
	Year of merger	0.074**	0.035	0.021	0.025	0.028*	0.017	0.011
		(0.034)	(0.025)	(0.017)	(0.017)	(0.015)	(0.015)	(0.012)
	One year after merger	0.070**	0.064**	0.041**	0.044**	0.041**	0.028	0.024
		(0.035)	(0.027)	(0.020)	(0.019)	(0.018)	(0.017)	(0.015)
	2+ years after merger	0.056	0.088***	0.068***	0.063***	0.059***	0.041**	0.036*
		(0.040)	(0.033)	(0.026)	(0.024)	(0.022)	(0.021)	(0.019)
	Observations	8,655	8,655	8,655	8,655	8,655	8,655	8,655

Notes. * $p < .10$, ** $p < .05$, *** $p < .01$. OLS estimates of equation (3) with standard errors in parentheses (clustered by hospital and system). All regressions include hospital fixed effects and year dummies. The dependent variable is our risk-adjusted inpatient price index. Controls: share of the privately insured covered by the HCCI insurers, number of technologies, a dummy for being ranked in *U.S. News & World Reports*, size as measured by number of beds, hospital ownership (government, nonprofit, or for-profit), whether a teaching hospital, percent of county uninsured, county median income, the Medicare payment rate, share of Medicare, share of Medicaid, year dummies, and HRR fixed effects. Post merger is dummy equal to 1 in the year a hospital merges and in all years afterward and 0 otherwise. "Neighbor" = 1 if a hospital was not involved in the merger, but within the distance indicated in the column head of a hospital where a merger took place (and 0 otherwise).

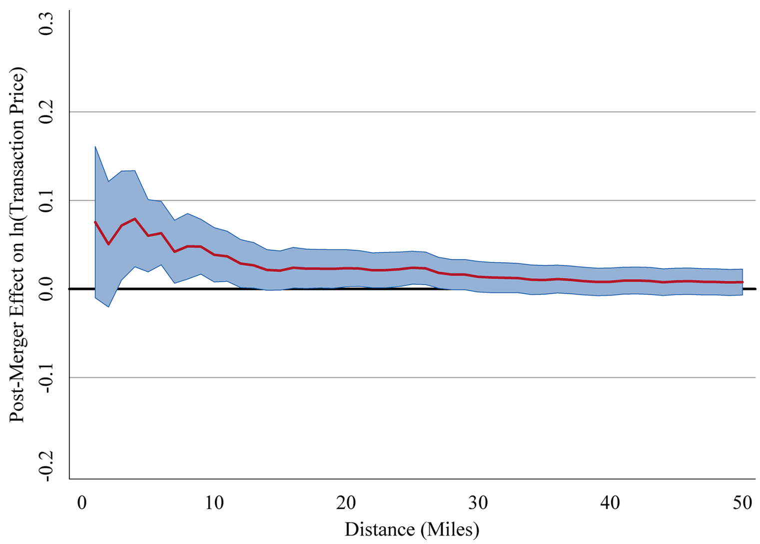


FIGURE X

How Merger Coefficient Changes for Mergers Between Hospitals of Different Geographical Proximity

These are the regression coefficients from equation (3) of postmerger effects on the log of regression-adjusted price for the sample of inpatient admission. These prices are risk-adjusted for DRG, age, and sex. We estimate the model separately for 50 specifications identical to that of Panel A in Table V. We allow the merger definition to vary in including merging hospitals within the distances shown on the x-axis. So a value of 10 corresponds to a merger of hospitals within 10 miles of each other. The shaded area presents the 90 percent confidence interval for each estimate.

between hospitals up to 25 miles apart. The magnitude of the merger coefficient declines as the distance between the merging parties increases. Mergers within 5 miles are associated with price increases of 6%, whereas the coefficients decline to 2% for mergers involving hospitals located up to 25 miles apart. In Figure X, we present the estimates of merger effects by 1-mile bins for all mergers up to those 50 miles apart. The estimates are noisy for very close mergers (because there are few such events), but the coefficient on mergers is broadly monotonically decreasing as the distance between the merging parties increases.

In Table V, Panel B, we add the control variables we included in our cross-sectional analysis, which makes almost no difference to the results. It is also possible that nonmerging neighboring hospitals may be affected by mergers (Dafny 2009). We test for this by adding a dummy for neighboring hospitals, which switches on after a neighboring hospital is exposed to a nearby merger (in the relevant distance bin). As we illustrate in Panel C, although the coefficients on neighboring mergers are usually positive, they are generally statistically insignificant.

It is possible that our estimates are capturing intertemporal factors other than the mergers themselves. Given the short time series in our panel, we examine price trends for two years before and after the merger event in Table V, Panel D, and Online Appendix Figure XIV. Reassuringly, there does not appear to be evidence of pretrends prior to the merger, as prices in the year before the merger are not significantly different from two years before (or earlier) in any of the columns. By contrast there are significant postmerger price increases, with higher prices in all columns two years after mergers occurred. The coefficients seem to generally build up from the year of the merger, but given the size of the standard errors, it is hard to be certain.

The merger coefficients we observe are economically significant.[61] A horizontal merger price effect of 5% is often used as an indicator of (enhanced) market power (U.S. Department of Justice and Federal Trade Commission 2010). Furthermore, this estimate represents the average effects of all mergers, not just those thought to be anticompetitive (as in previous ex ante studies). In addition, because we examine the impacts of consummated mergers, we are looking only at transactions that passed antitrust scrutiny. Since it is likely that the mergers with the largest potential effects on price are not attempted due to concerns over antitrust litigation or are blocked by enforcement authorities, those that we observe should be expected to have a smaller impact on price.

VII.E. Robustness of Merger Results

We subject our merger analysis to a large number of other robustness tests, some of which we discuss here.[62] First, instead of using the simple merger dummy, we estimate the cumulative

61. We note that our estimates are of the same or similar order of magnitude to the bulk of studies of merger price effects in other industries (Ashenfelter, Hosken, and Weinberg 2014).

62. The tests discussed here are contained in Online Appendix Table XXIV, where Panel A reproduces the baseline results from Table V.

merger effects by hospital for all mergers that hospitals were exposed to from 2007 to 2011.[63] Our postmerger price coefficients remain similarly scaled. Second, we used various matching procedures to identify alternative control groups for our analysis (see Online Appendix E), such as Mahalanobis distance matching between hospitals, the Dranove and Lindrooth (2003) procedure, and K-nearest neighbor matching. These tend to show slightly larger price effects for mergers within five miles than we observe in our baseline estimates. Third, we varied the 50 patients per year sample cutoff. This does not alter our main results. Fourth, it is possible that the price increases we observe following a merger could be due to improvements in management (e.g., hospitals doing a better job at price setting) rather than increased bargaining leverage (on the importance of management for hospital performance; see Bloom et al. (2015, 2017)). To test for this, we allow the merger coefficient to be different for targets and the acquirers and do not find statistically significant differences between the two. Finally, we also attempted to estimate merger effects for the seven procedures. Unfortunately, because those samples have fewer hospitals, there are fewer treated hospitals, so we cannot estimate merger effects with any precision.

VIII. Conclusions

Using insurance claims from three of the five largest commercial insurers in the United States, we find that health spending on the privately insured varies by a factor of three across the nation. Approximately half of the variation in private spending across HRRs is driven by differences in hospitals' prices and half by quantity (Medicare spending variation is almost all accounted for by quantity variation). Since previous research has focused on understanding the drivers of differences in the quantity of health care delivered across regions (Finkelstein, Gentzkow, and Williams 2016; Cutler et al. 2017), we focus on analyzing the variance in hospital prices.

Historically, the prices hospitals negotiate with insurers have been treated as commercially sensitive and have been largely unavailable to researchers on a national basis. Our data include

63. For example, of the 514 hospitals involved in at least one merger involving hospitals located less than 30 miles apart, 47 were involved in more than one merger from 2007 to 2011.

hospitals' transaction prices, and we are able to observe substantial variation in prices across hospitals, even for plausibly undifferentiated services like lower-limb MRIs. Moreover, a significant amount of the national variation in prices occurs within hospitals. This suggests that insurers' bargaining leverage influences the prices they negotiate with hospitals.

We use our data to characterize insurer-hospital contracts. When price is set as a share of charges (rather than prospectively paid), it offers hospitals weak incentives to lower costs, and it transfers the financial risk from idiosyncratically expensive cases to insurers. We find that approximately 23% of inpatient cases are paid as a share of charges and estimate that no more than 57% of inpatient cases are set as a percentage of Medicare rates.

Market structure appears strongly associated with hospitals' price levels and contract structure. Monopoly hospitals are associated with 12% higher prices, 10 percentage points more cases paid as a share of charges, and 11 percentage points fewer of their prospectively paid cases set as a percentage of Medicare payments compared to hospitals located in quadropoly or greater markets. In concentrated insurer markets we find the opposite correlations—hospitals have lower transaction prices and operate under contracts where they bear more risk. We also analyze the 366 hospital mergers that occurred between 2007 and 2011 and find that after mergers involving hospitals located less than five miles apart, prices at the merging parties increased by over 6%. As the distance between the merging parties' increases, the size of the postmerger price increases is attenuated. This set of results around market structure suggests that bargaining leverage is an important component of the dispersion we see in transaction prices.

Collectively, our research highlights the importance of studying hospital pricing and contracts when analyzing health spending on the privately insured. Our findings suggest that policy makers should continue to analyze whether potential hospital mergers could harm consumer welfare. Likewise, although we cannot draw strong normative conclusions, quantifying the scale of the variation in prices is still important. Given the variation in prices that we observe (particularly for undifferentiated procedures), our results suggest that patients and payers could save significant amounts of money if patients attended lower-priced providers. This suggests that policies aimed at steering patients toward low-cost providers (e.g., reference pricing, incentivizing referring physicians) could lower spending. Finally, there is widespread agreement that payment reform (shifting to contracts where providers bear more risk) is crucial to increasing hospital productivity (McClellan et al. 2017). Our analysis suggests that providers who have fewer potential competitors will be more able to resist attempts at such payment reform.

Further research should be focused on understanding the economic forces behind the patterns and correlations we have identified in the data. Given the growing availability of insurance claims data, there is scope for a rich and broad variety of research that takes on these important tasks.

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Supplementary Material

An Online Appendix for this article can be found at *The Quarterly Journal of Economics* online. Data and code replicating tables and figures in this article can be found in Cooper et al. (2018), in the Harvard Dataverse, doi:10.7910/DVN/ERXASS.

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