Hospital Systems and Bargaining Power: Evidence from Out-of-Market Acquisitions*

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Forthcoming in RAND Journal of Economics

Final Version: August 13, 2016

Analyses of hospital mergers typically focus on acquisitions that alter local market concentration. However, as prices are negotiated between hospital systems and insurers, this focus may overlook the impact of cross-market interdependence in the bargaining outcome. Using data on out-of-market acquisitions occurring across the U.S. from 2000-2010 we investigate the impact of cross-market dependencies on negotiated prices. We find that prices at hospitals acquired by out-of-market systems increase by about 17% more than unacquired, stand-alone hospitals; and confirming that out-of-market mergers result in a relaxation of competition, the prices of nearby competitors to acquired hospitals increase by around 8%.

JEL Codes: L10, L41, I11

Over the last 20 years, there has been a transformation in the way researchers and antitrust authorities assess the competitiveness of hospital markets and evaluate mergers. Theoretical and empirical models of hospital competition now incorporate more directly the contract negotiations between managed care organizations (MCOs) and hospitals. These models reveal that equilibrium reimbursement rates depend heavily on the relative bargaining positions of the MCO and hospital, which are determined by the profits each party would earn if they fail to agree to a contract that includes the hospital in the MCO's provider network. When two or more potentially substitutable hospitals merge and jointly negotiate prices with an MCO, the relative bargaining position of the hospital system increases because the MCO now finds it increasingly difficult to offer enrollees a sufficiently attractive

* The authors benefitted greatly from the comments and suggestions of David Balan, Samuel Kleiner, Aviv Nevo, and three anonymous referees, as well as seminar participants at the University of North Carolina, Clemson University, the University of Alabama, the Bureau of Economic Analysis, the Congressional Budget Office, the 2013 International Industrial Organization Conference in Boston, the 2014 American Society of Health Economists Biannual Conference in Los Angeles, and the 2015 American Economic Association Annual Meeting in Boston. We also thank Yaxuan He for excellent research assistance. An online appendix containing additional results tables and robustness checks is available from the author's website (`http://kevinpflum.com/papers/Lewis_Pflum_RAND_online_appendix.pdf`).

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provider network without this group of hospitals. Recent empirical studies have consistently found that mergers between local rival hospitals result in significantly higher reimbursement rates (e.g. Capps, Dranove, and Satterthwaite, 2003; Gowrisankaran, Nevo, and Town, 2015), and antitrust authorities have adopted these new approaches to more successfully challenge local mergers (Farrell, Balan, Brand, and Wendling, 2011).

Mergers between hospitals located in different markets are also fairly common and have contributed significantly to the persistent expansion of large regional and national hospital systems in recent years. Despite this, relatively little is known about the potential competitive effects of these out-of-market mergers and acquisitions. Existing theoretical and empirical approaches have not incorporated the potential for cross-market dependencies, focusing only on the impact of changes in local market structure. Yet, there are mechanisms that can generate cross-market interdependencies that may have an impact on negotiated reimbursement rates. For example, Vistnes and Sarafidis (2013) consider situations in which employers attempt to contract with a single MCO to cover employees in multiple patient markets and illustrate how a hospital in such a situation might be more valuable to the MCO if it is a member of a system that operates hospitals across many of these different markets. Alternatively, Gowrisankaran et al. (2015) suggest that mergers involving hospitals in different markets could generate price changes by allowing an MCO operating in both markets to arrange adjustments in relative reimbursement rates across cities that are mutually beneficial to both the MCO and the hospitals. Finally, Lewis and Pflum (2015a) highlight that systems may also provide their member hospitals with additional information and support during contract negotiations, allowing the hospital to extract from MCOs a larger share of the profits generated by a successful contract (i.e., hospitals in systems may have higher bargaining power).

Despite this growing awareness of these additional competitive effects, there has been essentially no empirical investigation of their relative importance or overall impact on hospital prices. To begin to fill in this gap, we investigate the degree to which system membership may influence negotiated reimbursement rates though channels unrelated to local market concentration. Rather than attempting to specify a particular model or test a particular mechanism we identify the overall importance of these additional mechanisms by examining the impact of out-of-market mergers that do not alter local market structure. Using a difference-in-differences approach to study 81 such hospital acquisitions occurring across the United States during the years 2000–2010, we find that the average net reimbursement rates at these hospitals increase by about 17 percent after joining an out-of-market hospital system with some specifications suggesting even larger effects.

We adopt a variety of strategies to control for unobservable market conditions that may influence hospital pricing or be correlated with the likelihood of a hospital acquisition. To ensure that treatment effects are not driven by differing price trends, we estimate specifications that allow for separate sets of year fixed effects for acquired hospitals and control hospitals and specifications that include leads and lags of the treatment variable to reveal discrete jumps or kinks in price trends at the time of acquisition. We also estimate a difference-in-difference-in-differences (triple-differences) specification which compares the difference-in-differences estimate of the price change experienced by acquired hospitals with the difference-in-differences estimate of the change in price experienced by other hospitals in the markets where these hospitals are acquired.

Additional evidence is provided by examining the impact of out-of-market acquisitions on the prices of nearby rival hospitals. Neighboring hospitals do not experience a change in ownership, eliminating any possibility that observed changes in their price mistakenly reflect some type of acquisition-related accounting or administrative change in hospital operations rather than an actual price change. The findings confirm that out-of-market mergers result in a relaxation of competition by revealing that prices rise at nearby rival hospitals in response to price increases by acquired hospitals. As expected, the price increases at rival hospitals are smaller than at the acquired hospital and decrease in magnitude as distance from the acquired hospital increases, providing additional evidence that the estimated price effects are directly caused by the acquisition rather than reflecting some confounding change in the local market.

We also consider the possibility that the identified increases in reimbursement rates are associated with other cost-related changes that may occur at hospitals during acquisition. A collection of auxiliary regressions show that the observed price increases do not appear to be a result of changes in patient case-mix or hospital quality, or a change in the cost of providing care more generally. Moreover, the results of the rivals analysis further contradict the possibility that average prices at acquired hospitals increase because patients that require more intensive treatment substitute to the hospital in response to some change in quality or service, as we would then expect patient complexity and prices to fall at nearby rival hospitals.

Our findings represent the first direct empirical evidence of cross-market dependencies in hospital competition and should serve as a warning sign to researchers and policymakers that have previously focused their scrutiny exclusively on in-market mergers.[1] Although many potential cross-market mergers might not involve cross-market dependencies, the acquisitions that systems have chosen to undertake over the last decade substantially increased prices at acquired hospitals on average. It is crucial to better understand and account for these effects when analyzing hospital competition more generally. The implications for competition policy depend on the nature of the cross-market dependencies. For example, if prices increase because large employers have fewer MCOs with which to negotiate coverage for employees spread across multiple markets, then this reflects a standard "lessening of competition" due to the elimination of substitutes (similar in many ways to the elimination of competitors within markets). On the other hand, if prices increase because merged hospitals are simply able to extract from MCOs a larger share of the profits generated from a successful contract (i.e., they benefit from an enhanced bargaining power or bargaining weight), it is not entirely clear that a lessening of competition has occurred.

We conclude our analysis with an investigation of how acquisition price effects vary by: the size of the acquiring system, the size of the acquired hospitals, and the proximity to other markets with system partners. The results reveal that acquisition-related price increases remain large even when acquired hospitals are very far from existing system partners and are located in seemingly unrelated geographic markets. Interestingly, however, prices do increase somewhat more when a hospital is acquired by a larger system or when the acquired hospital is relatively small. We cautiously interpret these patterns to be consistent with the idea that out-of-market mergers impact prices by improving the acquired hospitals

[1]Current antitrust evaluation of hospital mergers relies on establishing how substitutable hospitals are from the perspective of patients (FTC and USDOJ, 2004) implying that cross-market mergers are unlikely to ever be evaluated beyond the initial screening phase.

negotiating ability (or bargaining power) and that the observed price effects are not likely to be entirely explained by employers attempting to insure employees in multiple markets. More research is necessary, however, to investigate the different potential sources of cross-market pricing power and determine the extent to which such mergers warrant greater antitrust scrutiny.[2] In addition, because some mechanisms capable of generating price increases in out-of-market mergers have the potential to also impact in-market mergers, existing models should be re-evaluated to assure that assessments of local mergers are not biased by the exclusion of other relevant factors.

I. Empirical Strategy and Data

Several existing studies have found suggestive evidence that cross-market system affiliation can lead to higher prices. Melnick and Keeler (2007) discuss this possibility and present reduced-form evidence that system hospitals seem to receive higher prices than non-system hospitals, even after controlling for local market share and concentration. Lewis and Pflum (2015b) extend the standard structural bargaining model to identify differences in bargaining power across hospitals and show that system hospitals appear to have higher bargaining power. Both studies rely largely on cross-sectional variation, however, and consequently are only able to establish evidence of a correlation between system membership and cross-market price effects.

To more directly estimate a causal effect of system affiliation on reimbursement prices, we examine the observed changes in reimbursement rates associated with actual system acquisitions using a difference-in-differences approach and focus our investigation on out-of-market system acquisitions. Difference-in-differences methods have been used in several case studies focusing on individual hospital mergers (e.g.; Vita and Sacher, 2001; Tenn, 2011; Haas-Wilson and Garmon, 2011) and by Dafny (2009) to analyze mergers of co-located hospitals (within .5 miles). However, these studies exclusively examined mergers between local competitors. Dafny's study is unique in that it incorporates an instrumental variables (IV) estimation to control for the potential endogeneity of the acquisition. Unfortunately, as even the best instruments available to predict the likelihood of acquisition do

[2]We discuss the implications for antitrust policy in more detail in Section VI.

not vary significantly over time, using IV estimation requires collapsing the data to adopt a cross-sectional approach (as Dafny (2009) does), ignoring panel variation in the timing of the observed mergers. As a result, identification requires the potentially strong assumption that expected price trends (in the absence of a merger) are uncorrelated with the market characteristics (i.e., instruments) that make acquisition more likely. In contrast, by using a panel our difference-in-differences approach relies more heavily on the timing of the merger. This allows price trends to be correlated with the likelihood of a merger and instead requires only that the exact timing of the merger is uncorrelated with any remaining unobserved factors that influence prices.

As in Dafny (2009) and other observational merger studies, our estimates can only reveal the impact of system acquisitions that firms have chosen to undertake. These choices are not random. It is likely that acquisitions of other hospitals might not generate the same changes in reimbursement rates or profit margins. However, given that these out-of-market mergers have become more common and have been largely ignored in previous analysis, measuring the effects of these mergers is a crucial first step in understanding the influence of systems across markets and informing future work that may be able to more accurately predict the effects of a broader set of proposed mergers.

A. Data

We construct a panel spanning the years 1998 to 2010 primarily using data from the American Hospital Association's (AHA) Annual Survey of Hospitals and the Healthcare Cost Report Information System (HCRIS) maintained by the Centers for Medicare and Medicaid Services' (CMS). Data from these two sources are matched using the CMS Medicare provider numbers and observations are at the hospital-year level. We augment the data with county-level characteristics using the U.S. Census County Intercensal Estimates (proportion of population above 65 years old), the U.S. Census Bureau Small Area Income and Poverty Estimates (median income and proportion in poverty), and the BLS Local Area Unemployment Statistics (unemployment rate).

Data on hospital system status come from the AHA annual survey of hospitals. The AHA annually surveys all of the approximately 6,000 hospitals within the U.S. and its territories soliciting information such as system membership, ownership type, and service offerings. We only consider hospitals from the non-territorial United States that provide general short term care and have an average of at least 25 beds over the sample. A hospital is included in our analysis based on one of two criteria: if it was never part of a system during our sample period, or is observed to have joined a system sometime between 2000–2010 and remained in a system for the remainder of the sample period. If a hospital enters the market during this period they are included in the sample as long as they are not part of a system for at least two years before their acquisition. In addition, the first year of data is dropped for these entering hospitals as it generally reflects a partial year of operation and is typically quite different from subsequent years. We also exclude any hospital that appears to exit and re-enter the data. Based on the data these appear to reflect hospitals that have been closed and reopened, often as a specialty hospital. The sample used for estimation includes 1,859 general acute care hospitals that are not in a system at the beginning of our sample period, so the control group for our analysis becomes all hospitals that do not join a system at any time between 1998 and 2010.

We want to distinguish between those system acquisitions in which the system has no other general acute care hospitals in the patient market of the acquired hospital—out-of-market acquisitions—and those in which the acquiring system already has other acute care hospitals—in-market acquisitions. Identifying a hospital's market is often done with discharge-level data (e.g., Melnick et al., 1992; Melnick and Keeler, 2007; Capps et al., 2003); however, as we do not have discharge data for all of the hospitals in the sample and because we are not attempting to measure market shares or HHIs we simply categorize hospitals as belonging to the same patient market if they are within 45 miles from one another.[3] As we discuss in Section VI our results are robust to expanding this cut-off. Distance is calculated as the straight-line distance between hospitals' latitude and longitude. Figure 1 reports the number of in- and out-of-market system acquisitions in our sample.

All Medicare-certified hospitals must submit an annual cost report to CMS. The data from the reports are maintained in HCRIS and are available for download by the public at

[3]Based on the discharge abstracts for 2000-2010 provided by the California Office of Statewide Health Planning and Development, fewer than 1% of the patients would be predicted to choose a hospital more than 45 miles from their chosen hospital if it was removed from their choice set.

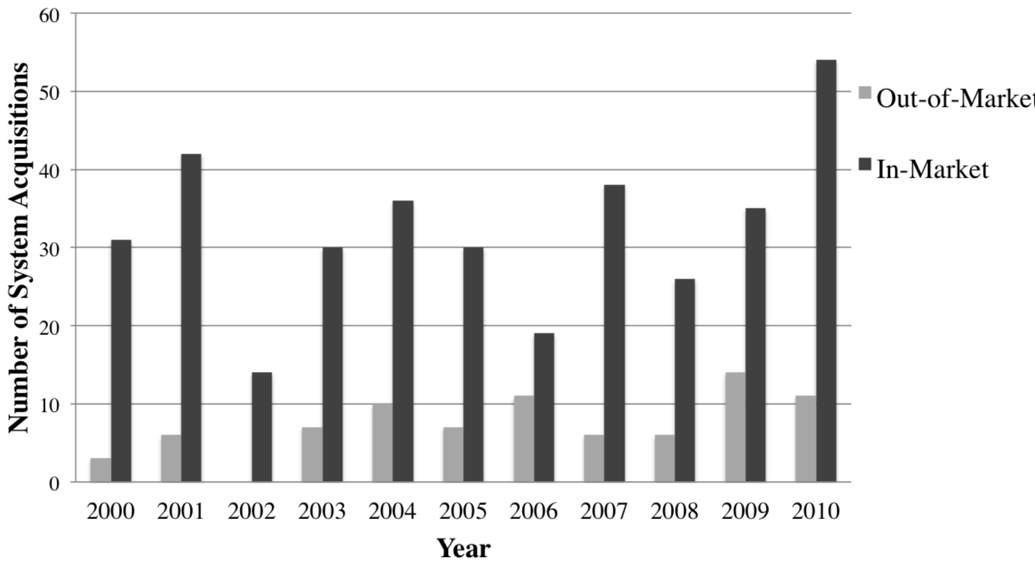


FIGURE 1. SYSTEM ACQUISITIONS BY TREATMENT GROUP, 2000-2010

the HCRIS website at `cms.gov`. The HCRIS data include the total gross charges of each hospital in each year, reported separately for inpatient and outpatient care. Gross charges represent the revenue the hospital would receive if it was paid its list price. As Medicare and Medicaid pay an administratively established rate and most insurers negotiate special rates that are less than the list price, a hospital receives its list price for very few, if any, discharges. HCRIS also contains the total revenues of each hospital net of its contractual deductions from list price. Net revenues are not reported separately for inpatient and outpatient charges so gross charges must be used in conjunction with net payments to formulate an inpatient price.

Using an approach similar to Dafny (2009) we construct a measure of the net revenues from inpatient discharges for a hospital in a given year by multiplying a hospital's gross charges for inpatient care by the ratio of the hospital's total net revenues to gross charges. We then subtract the total (net) amount received from Medicare for inpatient services from the total net inpatient revenue to calculate the total net inpatient revenue from non-Medicare patients.[4] Information on the number of inpatient discharges are reported by payer type. To obtain our measure of average price per discharge for non-Medicare patients, we divide the total net inpatient revenue from non-Medicare patients by the observed number of non-

[4]The HCRIS data on revenue and gross charges are not separated by payer type, except that the total (net) amount received from Medicare for inpatient services is reported.

Medicare inpatient discharges. This represents a weighted average of the price per discharge for privately insured and Medicaid patients weighted by the share of inpatient discharges coming from each type.[5]

Although we are unable to net out Medicaid patient revenues like we do for Medicare, in most cases Medicaid patients represent a relatively small share of a hospital's discharges, so the average net revenue per discharge largely reflects negotiated prices paid on behalf of privately-insured managed care patients. Moreover, as described in the next section, we adjust the treatment effects to correct for the presence of Medicaid discharges at the hospital when analyzing the estimated impact of hospital mergers on these average prices. Nevertheless, to ensure that the hospitals in the sample have a sufficient number of privately insured patients so that their revenues substantially reflect the outcomes of price negotiations, we include only those hospitals in which at least 10% of the patient population is privately insured.[6]

We also utilize the Medicare case mix index (CMI) assigned by the CMS in some auxiliary specifications to analyze how case mix may be affected by an acquisition.[7] The CMI measures the relative weight for all of the Medicare discharges at a hospital for a given year and represents the differences in clinical complexity and resource use required to treat discharges belonging to different diagnosis-related groups (DRGs). The CMIs are collected from Medicare's inpatient prospective payment system (PPS) final rules which are also available online at `cms.gov`.

The standardized collection procedure that includes all hospitals nationwide over a long time-period provides substantial advantages for the study of acquisitions and their effects. Average price measures derived from hospital-level HCRIS reports are almost certainly measured with error; however, so we put considerable effort into minimizing the impact of noise in the data and showing that our findings are robust across a variety of different tests.[8] To eliminate outliers generated by data entry errors, like Dafny (2009), we trim the

[5]A detailed description of the price derivation we use is provided in Appendix A.

[6]Although 10% appears small, our prices are based only on non-Medicare payments, so the more important number is the private share of non-Medicare patients. Only 1% of the sample has a non-Medicare private share below 29% and the sample average is 77%. The results are robust to a 25% private share cut-off as well.

[7]We don't use the CMI as a control in the main specifications because it reflects the case mix of Medicare patients and not the privately insured. In fact we find that the Medicare CMI has virtually no correlation with the average price per discharge for privately insured patients.

[8]Exploratory analysis described in Online Appendix 3 also suggests that these HCRIS-based average price measures

lower and upper tails at the 5th and 95th percentiles of the price distribution. We also consider the possibility that there may be outliers with respect to how a hospital's price evolves over time.[9] To help insure that these are not driving the results we calculate a hospital's average price over the sample period and remove those hospitals in which the absolute value of the difference in the price for a given year and its average price ($|p_{ht} - \overline{p}_h|$) is above the 90th percentile; i.e., we remove hospitals that exhibit unusually large jumps in price.[10] We test the robustness of the results by performing the analyses with data that has not been trimmed in any way and with data that excludes hospitals exhibiting a price difference from its average that is above the 80th percentile and has observations in the lower 10th and upper 90th percentiles removed. Estimates from these additional analyses (reported in Online Appendix 2) confirm that the general pattern of acquisition price effects is similar regardless of the trimming procedures. The estimated effect of an out-of-market acquisition is larger and less precise when no trimming is performed but is only slightly smaller when larger trim levels are used. Table 1 reports the summary statistics for the hospitals in our sample.

II. ESTIMATION

The main treatment group, *Group1*, consists of 81 stand-alone hospitals that become affiliated with a system that is completely out-of-market to that hospital (has no members within 45 miles) during the period 2000-2010. By definition these are hospitals that have no affiliated partner hospitals within their local market at any time before or after merger during the sample period. These 81 hospitals are acquired by 45 distinct systems. For comparison purposes we also consider a second treatment group, *Group2*, that includes the 355 stand-alone hospitals that join a system in which there are other same system members within the same patient market at the time (or shortly after) they joined. These 355 hospitals are acquired by a total of 186 systems, 24 of which are also involved in out-of-market acquisitions.

do exhibit similar patterns of price variation to those implied by the more disaggregate discharge-level data sources that are becoming available for more recent years in certain states.

[9]Prices are logged and hospital fixed effects are used, so treatment effects are identified by changes in price (rather than relative price levels), and large changes will have a more substantial impact on estimates.

[10]In the online appendix we present results based on the removal of hospitals exhibiting a standard deviation in price that is in the upper 90th percentile. The results are similar to removing hospitals exhibiting a large jump in price.

TABLE 1—HOSPITAL SUMMARY STATISTICS

	Full Sample (N = 19,022)			
	Mean	Standard Dev.	Min.	Max.
ln(Price/Discharge)	8.773	0.591	7.331	10.715
ln(Cost/Discharge)	9.052	0.441	6.670	12.599
ln(# Beds)	4.880	0.910	1.792	9.746
Bed Utilization	0.436	0.209	0.001	1.000
% Private	0.364	0.142	0.100	0.996
% Medicare	0.505	0.158	0.000	0.893
% Medicaid	0.130	0.106	0.000	0.824
% OP Revenue	0.479	0.147	0.000	0.983
	Hospitals Acquired by Out-Of-Market Systems (N = 750)			
	Mean	Standard Dev.	Min.	Max.
ln(Price/Discharge)	8.842	0.506	7.481	10.590
ln(Cost/Discharge)	9.043	0.334	8.118	11.381
ln(# Beds)	4.867	0.728	2.708	6.368
Bed Utilization	0.421	0.179	0.013	1.000
% Private	0.366	0.134	0.100	0.745
% Medicare	0.512	0.134	0.139	0.859
% Medicaid	0.122	0.074	0.000	0.410
% OP Revenue	0.449	0.136	0.000	0.897

Notes: Summary statistics are based the sample used in the main analysis in which the tails of the Log(Price/Discharge) distribution have been trimmed at the 5th and 95th percentiles. Average costs are frequently higher than the average reimbursements due to the fact that average costs are likely biased upward slightly as we do not observe inpatient costs separately and instead disaggregate total operating costs based on the proportion of total patient revenues generated by inpatient care, while average reimbursements are biased downward slightly due to the presence of Medicaid patients.

Our model for estimation can be expressed as:

$$r_{ht} = \alpha + \beta_1 \mathit{T1}_{ht} + \beta_2 \mathit{T2}_{ht} + \kappa_{ht} + d_{ht}\delta_1 + g_{ht}\delta_2 + m_{ht}\delta_3 + \eta \mathit{FP}_{ht} + \mu_{Gt} + \xi_h + \epsilon_{ht}. \tag{1}$$

The dependent variable, r_{ht}, is the natural log of hospital h's reimbursement price at time t. $\mathit{T1}_{ht}$ is an indicator taking the value of 1 when hospital h is in *Group1* and is in a system at time t. In other words, $\mathit{T1}_{ht}$ represents the interaction of the indicator $\mathit{Group1}_h$ and an indicator System_{ht}, which is 1 when hospital h is in a system in year t and 0 otherwise (i.e., $\mathit{T1}_{ht} = \mathit{Group1}_h \times \mathit{System}_{ht}$). Similarly, $\mathit{T2}_{ht}$ is an indicator taking the value of 1 when hospital h is in *Group2* and is in a system in year t (i.e., $\mathit{T2}_{ht} = \mathit{Group2}_h \times \mathit{System}_{ht}$). The average price over the year in which a merger occurs will not accurately reflect the pre-merger or post-merger price level. System affiliation is collected by the AHA in the summer, so a hospital first appearing in a system in year t of the AHA records may have been acquired in calendar year t or $t-1$ of the HCRIS data. As a result, we include a merger-period indicator

κ_{ht} which is set to 1 for these two periods so that the difference-in-differences estimate more accurately compares reimbursement prices before merger to prices in the years following the merger. The $j \times 1$ vector d_{ht} represents characteristics of hospital h's discharges at time t that impact the cost of care; the $k \times 1$ vector g_{ht} represents characteristics of hospital h's patient population at time t that may bias the estimated reimbursement price; the $m \times 1$ vector m_{ht} represents characteristics of the county that hospital h is in at time t that could affect reimbursement prices; FP_{ht} is a dummy indicating whether or not hospital h is for-profit at time t to control for any confounding effects of changing objectives;[11] μ_{Gt} are treatment-group-specific year fixed effects that allow the prices of hospitals in *Group1*, *Group2*, and the control group to potentially evolve differently over time; ξ_h are hospital fixed effects; and ϵ_{ht} is a mean zero, heteroskedastic disturbance term capturing unobserved heterogeneity in hospitals' prices. We assume this disturbance, ϵ_{ht}, is independent across MSAs but may be correlated across hospitals and years for hospitals in the same MSA.[12]

The specification reported in (1) includes three groups of control variables. The d_{iht} includes the natural log of the hospitals' average cost per discharge and a measure of the hospital's capacity usage. Average cost per discharge is calculated by multiplying the hospital's total operating expenses by the ratio of gross inpatient revenues to total gross revenues and dividing by the number of discharges. We include the hospital's capacity usage to control for differences in opportunity costs that the accounting cost does not capture.

The second set of control variables, the g_{jht}, control for factors that could bias the measure of reimbursement price due to the data limitations. Although we are able to net out Medicare inpatient revenues, our calculated measure of net inpatient revenue for privately insured patients will still be distorted by the fact that the contractual discount measure is based partially on revenues from outpatient care, which also includes Medicare and Medicaid outpatient revenues. If contractual discount rates for outpatient care systematically differ from those for inpatient care, then our measure of inpatient net revenue will be biased in the direction of the difference. To help control for these distortions our specification allows the average reimbursement rate measure to vary as a function of the fraction of the hospital's pa-

[11]Of the hospitals in our sample, 13 acquired by an out-of-market system and 23 acquired by an in-market system change to for-profit status after acquisition. Fifty of the control hospitals also change to for-profit status.

[12]For hospitals that are located outside of an MSA we allow for correlation at the three-digit zip code level.

tients insured by Medicare and Medicaid and the fraction of revenues from outpatient care. As HCRIS does not report the number of outpatient visits, we use the overall proportion of gross revenues that outpatient revenues represent and the proportion of inpatient discharges that are from patients insured by Medicare and Medicaid to proxy for their outpatient proportions. In 2001 Medicare made a substantial change to the way it pays for out-patient services and research has found that this change had some impact on patient volumes and average payments (He and Mellor, 2012). To account for this we include the Medicare share of OP revenue as a control and allow the coefficient to have a different value in the pre-2001 period.

The third set of control variables, the m_{mht}, control for county-level characteristics that could affect the reimbursement price either directly or indirectly through limitations of the data used in the estimation. We include as controls the unemployment and poverty rates, which could impact the amount of uncompensated care provided by a hospital; the median income level, which measures the affluence of patients likely to visit the hospital and may be correlated with their price elasticities and the reimbursement price hospitals can secure; and the percent of the population above 65, which captures the importance of the Medicare eligible population to the hospital.

Recall that because only Medicare inpatient revenues are netted out, the average price per discharge represents a discharges-weighted average of the private and Medicaid reimbursements. In consequence, the treatment effects are also attenuated by the presence of inpatient discharges from Medicaid patients in our average revenue measure. As reimbursement rates paid by Medicaid are set by state regulators, any increase in market power resulting from a system acquisition will impact only the share of patients that are privately insured. Therefore, the extent to which the true treatment effect for privately insured patients is reflected in this average price will be proportional to the share of the hospital's non-Medicare discharges that are privately insured. Assuming that system acquisitions have no impact on rates paid by Medicaid patients, we can correctly identify the treatment effect on prices for the privately insured by interacting the treatment indicator variables with the share of the hospital's non-Medicare discharges that are privately insured. To avoid endogeneity concerns we utilize the 1998 value of private patient share in all subsequent years.[13]

When this correction is implemented it results in the following alternative model:

$$(2) \quad \begin{aligned} r_{ht} = \alpha + [\beta_1 \mathit{T1}_{ht} + \beta_2 \mathit{T2}_{ht}] \times \mathit{PrvtShare}_h + \kappa_{ht} \\ + d_{ht}\delta_1 + g_{ht}\delta_2 + m_{ht}\delta_3 + \eta \mathit{FP}_{ht} + \mu_{Gt} + \xi_h + \epsilon_{ht}, \end{aligned}$$

where $\mathit{PrvtShare}_h$ represents the share of a hospital h's non-Medicare inpatient discharges that are from privately insured patients in 1998 and the other variables are the same as in eq. (1). The estimated values of β_1 and β_2 from this corrected model more accurately represent the impact of system acquisition on the prices a hospital negotiates with MCOs.

Our main difference-in-differences model includes separate sets of year fixed effects for the treatment groups and control group to allow for differences in trends between groups that are not explained by observable hospital characteristics and might otherwise introduce bias. As a result, the identification of treatment effects relies on the assumption that price changes observed at other hospitals in the treatment group that were not acquired by a system within the previous year represent a valid counterfactual for how prices would have changed at acquired hospitals had they not been acquired by a system in that year. In other words, the timing of acquisition is assumed to be uncorrelated with unobservable non-acquisition-related factors that also affect price changes.[14]

We also estimate a variety of additional specifications that alter the control group and/or counterfactual to require less stringent assumptions. For example, we estimate a triple-differences specification that is also robust to the presence of unobservable factors that might impact prices of hospitals in the acquisition area more generally around the time of acquisition. Identification in this case only requires that the timing of acquisition is uncorrelated with unobserved factors that might cause the acquired hospital's prices to deviate from others in the city (after also conditioning on treatment-group trends). Additional specifications introduce leads and lags of the acquisition indicator into the model, relaxing the assumption that the impact of acquisition on prices is completed within the first year and

[13]For those hospitals that enter the sample after 1998, we use the hospital's non-Medicare share from its second observation year as the first data year does not represent a full year of operation.

[14]For comparison, we also include specifications in which there is only one set of year fixed effects so that identification of the treatment effect is also identified off of the difference in the prices at a treated hospitals relative to both treatment hospitals that were not acquired by a system within the previous year and non-system control hospitals.

instead allowing the merger price effect to continue evolving over subsequent years. The presence of acquisition leads also help to confirm that prices did not begin increasing before acquisition, further reducing the likelihood that unobserved factors or reverse causality might be responsible for generating positive acquisition price effects. As a result, any endogeneity in the treatment would likely require that certain stand-alone hospitals periodically experience exogenous, permanent jumps in their average reimbursement rates and that systems identify these jumps beforehand and systematically acquire these hospitals during the year when these price increases were to occur. We view this to be unlikely and believe that the ability to relax the underlying difference-in-differences identification assumptions along a variety of dimensions strengthens our ability to convincingly uncover causal acquisition price effects.

III. Results

Table 2 reports the results of several specifications based on eq. (2). All specifications include hospital and year fixed effects, whereas columns 4 – 6 additionally include treatment-group-specific year fixed effects. Columns 2, 3, 5, and 6 include the natural log of the cost per discharge, and Columns 3 and 6 additionally include controls for the proportion of patients that are insured by Medicare and Medicaid and the proportion of a hospital's gross revenues that can be attributed to outpatient care, all of which will affect the calculated price per discharge; the number of beds at the hospital the utilization rate and utilization rate squared to help control for differences in the hospital's opportunity costs; and county-level characteristics that could influence the price level.

Absent any cost controls the treatment effects are smaller, the standard errors are larger, and, in the case of out-of-market acquisitions, the estimates are not statistically significant. The lack of significance is not surprising as hospitals exhibit a fairly large amount of variation in the case mix of the patients it treats from one year to the next. Not controlling for cost differences introduces more noise to an already noisy approximation of price.

Including the average cost per discharge significantly improves the fit of the model, the estimates for the treatment effects are larger, and the standard errors are reduced. Columns 2 and 3 indicate that both out-of-market and in-market acquisitions generate about a 9.6 per-

TABLE 2—THE IMPACT OF SYSTEM MEMBERSHIP ON THE PRICE/DISCHARGE

ln(Price/Discharge)	(1)	(2)	(3)	(4)	(5)	(6)
T1× Private Share	0.037	0.096^{b}	0.096^{b}	0.128^{a}	0.170^{c}	0.173^{c}
	(0.057)	(0.041)	(0.042)	(0.075)	(0.055)	(0.058)
T2× Private Share	0.067^{b}	0.099^{c}	0.096^{c}	0.060^{a}	0.102^{c}	0.109^{c}
	(0.026)	(0.020)	(0.019)	(0.036)	(0.025)	(0.026)
ln(Cost/Discharge)		1.205^{c}	1.124^{c}		1.204^{c}	1.124^{c}
		(0.035)	(0.036)		(0.035)	(0.036)
FP× Private Share			0.023			0.018
			(0.029)			(0.028)
% Medicare			0.288^{a}			0.276
			(0.174)			(0.175)
% Medicaid			-0.183^{c}			-0.180^{c}
			(0.069)			(0.069)
% OP Revenue			-0.939^{c}			-0.952^{c}
			(0.172)			(0.172)
ln(# Beds)			-0.026			-0.025
			(0.028)			(0.027)
Bed Utilization			0.200^{c}			0.205^{c}
			(0.066)			(0.066)
Bed Utilization Sqrd.			-0.066^{a}			-0.068^{a}
			(0.038)			(0.039)
Unemployment			-0.005^{b}			-0.005^{b}
			(0.003)			(0.003)
Poverty rate			0.003			0.003
			(0.002)			(0.002)
Median income (/1000)			0.066^{c}			0.066^{c}
			(0.013)			(0.013)
Percent over 65			0.632^{a}			0.675^{a}
			(0.383)			(0.384)
Year Fixed Effects						
Year Only	X	X	X			
Trmt×Year				X	X	X
Adj. R^2	0.133	0.450	0.474	0.134	0.451	0.475
N	19,009	19,009	19,009	19,009	19,022	19,009

Notes: All specifications include hospital and year fixed effects while columns 4 – 6 also include treatment-specific year fixed effects. Private Share is the non-Medicare private share and accounts for the Medicaid patients that are contributing to the hospital's calculated price. Standard errors in parentheses are clustered by MSA. Significance Levels: a =p $<$ .10, b =p$<$ .05, c =p $<$ .01

cent increase in reimbursements. The impact of an out-of-market acquisition jumps to about 17 percent whereas the impact of an in-market acquisition rises slightly to about 10 to 11 percent when treatment-group-specific year fixed effects are included. This is because prices at hospitals that are acquired by out-of-market systems exhibit a slight downward trend prior to acquisition, resulting in an underestimation of the treatment effect absent treatment-group-specific year fixed effects.[15] This downward trend could suggest that these hospitals were underperforming, hence better targets for acquisition. Given the importance of controlling for the differences in how the prices at treatment hospitals evolve over time prior to acquisition we also estimate a specification that includes one set of year fixed effects but allows for separate hospital year trends.[16] Those estimates are reported in Appendix Table B7 and support the estimates reported in Table 2.

The patient population controls indicate that prices are increasing with the share of a hospital's patients that are insured by Medicare; and, reflecting that Medicaid payments are not netted out, an increase in the proportion of patients insured by Medicaid results in a lower estimated reimbursement price. The negative coefficient estimates on outpatient revenue share suggest that outpatient prices are discounted more heavily than inpatient prices.[17] The estimates also suggest that hospitals that are near their capacity limit have higher prices. These higher prices could reflect some increase in bargaining power that comes from an ability to play insurers off of one another because the hospital will not need to contract with all insurers to maximize utilization. Similarly, it could capture the fact that the opportunity cost of utilizing an inpatient bed becomes very high when the hospital is near its capacity limit.

Interestingly, the average price effect generated by the out-of-market system acquisitions observed in our sample is larger than the average price effect for observed in-market acquisitions. We should not conclude from this result, however, that a particular hospital would have been able to increase its price by more if it had been acquired by an out-of-market system. The underlying sources of market power generated by these two types of acquisitions are likely to differ and hospitals will be selected based on these and other factors. Whereas most of the gains in market power for in-market acquisitions may come from reductions in the degree of competition within that market, any market power generated by out-of-market mergers occurs through different mechanisms, which we discuss further in

[15]Table 4 in the online appendix reports the estimates for the year fixed effects for column 6.

[16]We thank an anonymous referee for suggesting these specifications.

[17]Recall that we separately observe gross revenues from inpatient and outpatient care, but observe only an aggregate measure of the discounts provided to insurers.

Section VI.

Given the antitrust authorities' scrutiny of local hospital mergers, it is not surprising that most of the in-market acquisitions we observe do not substantially increase market concentration and may have little impact on market power.[18] These acquisitions were typically only pursued (and allowed to take place) in very competitive markets (e.g., larger cities) or when the merging parties were not neighbors or direct competitors. For the in-market acquisitions in our data, the average distance to the closest system partner is 22 miles. Using a rough measure of HHI based on hospital discharge shares within a 45 mile radius of the acquired hospitals, the average acquisition in our sample resulted in an increase in HHI of around 55, and for about 95 percent of the in-market acquisitions the increase in HHI is less than 100. These increases are well below the threshold for further scrutiny under the U.S. Horizontal Merger Guidelines (Section 5.3). In the relatively few cases where neighboring hospitals merge, price increases are somewhat larger. Appendix Table B6 reports the results of specifications that separately estimate the price effect of in-market mergers that involve hospitals less than two miles away, between two and five miles away, and more than 5 miles away.[19] When hospitals within two miles of each other merge, prices increase by roughly 50% more on average than for mergers involving hospitals more than five miles away, though given the small number of such mergers this difference is not statistically significant.

The treatment indicators in Table 2 are interacted with the share of patients that are privately insured in the non-Medicare patient population to account for the fact that our measure of price is diluted by the inclusion of Medicaid patients in particular. Not interacting the treatment effects with the private share will downwardly bias the estimates. To check the importance of this correction we re-estimate the specifications reported in Table 2 but based on eq. (1), in which the treatment dummies are not interacted with the non-Medicare private share. Table 3 reports the pure treatment effects for these specifications (full results are reported in Online Appendix Table 3). Privately insured patients represent about 73% of the non-Medicare patient population on average, and not controlling for the fact that the treatment effect only impacts these patients decreases the estimates by about 25 to 30 percent,

[18]The Federal Trade Commission (2012) reports several proposed mergers of hospitals that were dropped after the FTC challenged them. Recent examples include the proposed acquisition of Prince William Hospital by Inova Health System (D. 9326) and OSF Healthcare System's proposed acquisition of Rockford Health System (D. 9349).

[19]We choose such close distances because we do not have good measures of market concentration.

TABLE 3—THE PURE TREATMENT EFFECT FOR SYSTEM MEMBERSHIP ON THE PRICE/DISCHARGE

ln(Price/Discharge)	(1)	(2)	(3)	(4)	(5)	(6)
T1	0.019	0.073^{b}	0.070^{b}	0.082	0.129^{c}	0.127^{c}
	(0.049)	(0.034)	(0.035)	(0.062)	(0.045)	(0.047)
T2	0.047^{b}	0.074^{c}	0.072^{c}	0.032	0.071^{c}	0.080^{c}
	(0.021)	(0.016)	(0.016)	(0.029)	(0.020)	(0.021)
Year Fixed Effects						
Year Only	X	X	X			
Trmt×Year				X	X	X
Adj. R^2	0.132	0.449	0.474	0.134	0.451	0.475
N	19,009	19,009	19,009	19,009	19,009	19,009

Notes: Each specification includes the same control variables reported in the corresponding column in Table 2 as well as hospital and year fixed effects while columns 4 – 6 also include treatment-specific year fixed effects. Standard errors in parentheses are clustered by MSA. Significance Levels: a =p $< .10$, b =p$< .05$, c =p $< .01$

depending on the specification. The standard errors are also lower in magnitude by similar amounts so there is little change in the statistical significance. The fit of the model is also slightly lower in each specification.

A. *Market-wide Price Effects*

In our main analysis we are careful to consider the possibility that estimated treatment effects could be biased if underlying price trends at hospitals targeted for acquisition differ systematically from other hospitals absent the merger taking place. To control for such differences we estimate specifications that include treatment-group-specific year fixed effects. One might, however, have a related concern that acquisitions may be more likely to occur in markets where the prices of all hospitals are increasing (or decreasing) more quickly than in other markets, or that changes occurring in a particular market may impact both the local price trajectory and the likelihood of an acquisition occurring in the market. For example, a change in market structure on the insurer-side of the market (e.g., entry by additional insurers) could generate higher bargaining power for all hospitals in the area. The resulting reimbursement price increases could make independent hospitals in the market more attractive acquisition targets for a system wanting to expand. We want to be careful not to attribute these price increases to the acquisitions that they might induce.

One way to control for this type of bias is to compare the post-acquisition prices of acquired hospitals with the prices of other hospitals located in markets where acquisitions occur, as these hospitals will experience the same unobserved market-wide shocks as our treatment hospitals. We estimate this using a triple-differences specification which compares our difference-in-differences treatment effect estimate for acquired hospitals with an analogous difference-in-differences "treatment" effect experienced by hospitals when another hospital in the city is acquired. Specifically, let $T1\_RivalT1_{ht}$ indicate whether hospital h has either been acquired by an out-of-market system or a nearby rival within 25 miles has been acquired by an out-of-market system before time t and define $T2\_RivalT2_{ht}$ similarly for in-market acquisitions. The triple-differences specification then takes the form:

$$\begin{aligned} r_{ht} = \alpha + \beta_1 T1_{ht} + \beta_2 T2_{ht} + \beta_3 T1\_RivalT1_{ht} + \beta_4 T2\_RivalT2_{ht} + \kappa_{ht} \\ + d_{ht}\delta_1 + g_{ht}\delta_2 + m_{ht}\delta_3 + \mu_{Gt} + \eta FP_{ht} + \mu_t + \xi_h + \epsilon_{ht}, \end{aligned}$$

where r_{ht}, d_{ht}, g_{ht}, γ, κ_{ht}, μ_t, ξ_h, ϵ_{ht} are defined as in eq. (1). Care should be taken in interpreting the results, however, as the prices of rival hospitals are likely to represent an imperfect counterfactual. Any price increase enjoyed by an acquired hospital is likely to also allow rivals in the same market to increase their prices somewhat[20], so the triple-differences specification might be viewed as a conservative estimate of the price increase generated by a hospital acquisition.

Column 2 of Table 4 reports the results of the triple-differences estimation (all specifications include treatment-specific time fixed effects). For comparison, column 1 reports the corresponding difference-in-differences estimate that appeared in column 6 of Table 2. The estimates indicate that prices at non-acquired hospitals in markets where an out-of-market acquisition occurs tend to increase by 4.3 percent when the acquisition takes place. Similarly, rival hospitals in markets where an in-market acquisition occurs increase by 5.1 percent. However, prices at acquired hospitals increase significantly more than the other hospitals in their market. Those acquired by out-of-market systems increase prices by an additional 13.4 percent and those acquired by in-market systems raise prices by an additional 7.4 percent.

An alternative triple-differences specification can be estimated by simply including

[20]This results because an MCO's value of adding a particular hospital to its network will increase when a nearby substitute hospital raises its price.

TABLE 4—MARKET-WIDE PRICE EFFECTS

ln(Price/Discharge)	(1)	(2)	(3)	(4)
T1× Private Share	0.173^{c}	0.134^{b}	0.109	0.175^{c}
	(0.057)	(0.064)	(0.084)	(0.057)
T2× Private Share	0.109^{c}	0.074^{c}	0.052	0.112^{c}
	(0.025)	(0.026)	(0.037)	(0.025)
Hospital or rival acquired by an...				
out-of-market system		0.042^{a}		
		(0.024)		
in-market system		0.051^{c}		
		(0.016)		
Rival acquired by an out-of-market system that is...				
$\leq$ 7.5 miles				0.078^{a}
				(0.040)
$>$ 7.5 and $\leq$ 25 miles				0.033
				(0.027)
Rival acquired by an in-market system that is...				
$\leq$ 7.5 miles				0.055^{a}
				(0.032)
$>$ 7.5 and $\leq$ 25 miles				0.038^{b}
				(0.016)
MSA×Year Fixed Effects			X	
Trmt×Year Fixed Effects	X	X	X	X
Adj. R^2	0.475	0.474	0.633	0.474
N	19,009	19,009	19,009	19,009

Notes: All specifications include the same control variables reported in the corresponding column in Table 2 as well as hospital and treatment-group-specific year fixed effects. Private Share is the non-Medicare private share and accounts for the Medicaid patients that are contributing to the hospital's calculated price. Standard errors in parentheses are clustered by MSA. Significance Levels: a =p $<$ .10, b =p$<$ .05, c =p $<$ .01

MSA-specific year fixed effects into our baseline specification.[21] This approach allows both pre-existing price trends and the price response of non-acquired hospitals when an acquisition occurs in their MSA to differ across MSAs. Unfortunately, in this model many of the counterfactual price levels (that would have been expected absent an acquisition) are identified using a very small number of hospitals (those within the same MSA), and as a result, the estimated treatment effects are less precise. Nevertheless, the findings (presented in column 3) suggests acquisition price effects that are similar though somewhat smaller than those in column 2.

These findings appear to reject the possibility that price increases at acquired hospi-

[21]For hospitals located in rural areas (outside of any U.S. Census Metropolitan or Micropolitan area) share a common fixed effect with all other rural hospitals in the same state.

tals are a result of market-wide unobserved shocks rather than true merger effects. This is particularly notable given that these triple-differences specifications are likely to be overly conservative as a result of rivals within the same market also having the ability to increase prices somewhat following an acquisition. The magnitude of this competitive reaction can be examined more carefully by noting that the impact of a hospital's price on its rivals will depend on how close of substitutes the hospitals are to one another. In consequence, if rival price effects follow from a competitive response to an increase in price at the acquired hospital we would expect rivals that are closer to an acquired hospital to exhibit a larger increase in price than rivals that are further away. In contrast, if rival price effects are generated by some market-wide change, such as a decrease in the concentration of insurers, then hospital prices within a particular market should change more uniformly.

In column 4 of Table 4 we alter our specification, allowing the price of each hospital to be a function of whether it was acquired by a system itself and also whether a nearby rival hospital (within a particular distance range) has been acquired. The estimates reveal that rivals nearer to an acquired hospital experience larger increases in price. When a hospital is acquired by an out-of-market system, its rivals within 7.5 miles increase their prices by an average of 7.8 percent, whereas rivals over 7.5 miles away increase price by only 3.3 percent. Similarly, rivals with 7.5 miles of a hospital acquired by in-market systems raise price by around 5.5 percent when the acquisition occurs, but those over 7.5 miles away increase price by only 3.8 percent.[22]

The patterns of observed price effects at rival hospitals support the interpretation that price increases are a direct result of the acquisition and originate from the acquired hospital itself rather than reflecting some unobserved market-wide shock. In addition, the presence of a positive price effect for nearby rival hospitals provides additional evidence that the price increases identified in our main analysis are real and not a result of mismeasurement or a failure to effectively control for changes in cost and patient mix that might have occurred during the acquisition.

[22]An alternative specification estimated with three distance categories (less than 5 miles, between 5 and 10 miles, and over 10 miles) exhibited a very similar declining pattern for both acquisition types, although smaller sample sizes in each category caused the coefficients to be somewhat less precisely estimated.

TABLE 5—LEAD AND LAG PRICE EFFECTS FOR OUT-OF-MARKET ACQUISITIONS

	(1)	(2)	(3)	(4)
Years Relative to T1 Acquisition				
t-6 or less	-0.060	-0.036	-0.087	-0.043
	(0.072)	(0.079)	(0.077)	(0.088)
t-5	-0.093	-0.049	-0.120	-0.055
	(0.064)	(0.073)	(0.076)	(0.091)
t-4	-0.010	0.001	-0.027	-0.000
	(0.048)	(0.050)	(0.056)	(0.058)
t-3	-0.013	-0.020	-0.016	-0.009
	(0.033)	(0.038)	(0.041)	(0.047)
t-2	0	0	0	0
	(.)	(.)	(.)	(.)
Acquisition Period				
t+1	0.173^{b}	0.157^{b}	0.226^{c}	0.210^{b}
	(0.071)	(0.075)	(0.080)	(0.083)
t+2	0.208^{c}	0.184^{b}	0.266^{c}	0.242^{c}
	(0.079)	(0.082)	(0.087)	(0.090)
t+3	0.197^{a}	0.190^{a}	0.256^{b}	0.249^{b}
	(0.102)	(0.106)	(0.111)	(0.116)
t+4	0.218^{a}	0.238^{b}	0.270^{b}	0.297^{b}
	(0.114)	(0.118)	(0.133)	(0.137)
t+5 or more	0.294^{b}	0.323^{b}	0.363^{c}	0.400^{c}
	(0.133)	(0.134)	(0.136)	(0.140)
Leads and Lags interacted with				
% Non-Medicare Private Share	No	No	Yes	Yes
Market-wide Acquisition Leads/Lags	No	Yes	No	Yes
Adj. R^2	0.475	0.475	0.475	0.476
N	19,009	19,009	19,009	19,009

Notes: A hospital first reports belonging to a system in year t, however, they may have joined anytime in the last year, between surveys. All specifications include hospital and treatment-specific year fixed effects as well as the controls reported in column 6 of Table 2. All estimates are relative to the omitted years, which are the year prior to joining a system and the first year the hospital was reported in a system as the acquisition took place sometime during these two years. Standard errors in parentheses are clustered by MSA. Significance Levels: a =p $<$ .10, b =p$<$ .05, c =p $<$ .01

B. *Pre- and Post-Acquisition Price Effects*

Treatment-group-specific year fixed effects help to assure that treatment effects are not driven by unobserved factors that differentially impact the prices of acquired hospitals as a group relative to control hospitals. Our triple-differences specification further mitigates concerns that unobserved market-wide shocks in cities experiencing acquisitions might be driving the price effects. Neither of these methods, however, can control for unobserved idiosyncratic shocks that impact hospitals at different times and might cause an independent hospital's reimbursement prices to increase suddenly and raise the likelihood that it is acquired. In this section we address this possibility by examining the behavior of prices at treatment hospitals in the years immediately preceding and following acquisition..

We identify price movements in more detail by including indicator variables based on the number of years between the observation year and the year the hospital is acquired. Pre-acquisition indicators will reveal whether prices are increasing in any meaningful way prior to acquisition, and post-acquisition indicators will provide some insight into how rapidly a hospital's reimbursements increase after acquisition. Tables 5 and 6 report the treatment leads and lags for several specifications for out-of-market and in-market acquisitions, respectively.[23] Each specification contains a treatment lead indicator for each year prior to the acquisition period with all leads at $t-6$ or less aggregated together for a total of 5 treatment lead indicators. Each specification also contains a treatment lag indicator for each year following the acquisition with all lags at $t+5$ or more aggregated together for a total of 5 treatment lag indicators. In columns 3 and 4 these leads and lags are interacted with the non-Medicare share of privately insured patients as in our main specification and represent the preferred specifications. Columns 2 and 4 are analogous to the triple-differences specifications reported in Table 4 in that they also include lead and lag indicators for hospitals that are nearby (< 25 miles) an acquired hospital. We choose the excluded year of the lead/lag indicators to be $t-2$, so all of the pre- and post-acquisition treatment effects reflect the average difference in price relative to the year right before the acquisition period.

Table 5 reports the treatment effects for out-of-market acquisitions. In all specifications, prices are quite stable in the 3 years prior to acquisition ($t-4$ through $t-2$), with no significant upward trend. Immediately following the acquisition, however, we see a sharp increase in prices, which then continues to rise somewhat in subsequent years. Given the larger number of parameters estimated, the standard errors on these acquisition leads and lags are relatively large, but the estimated prices in all post acquisition periods are significantly different from the price in the pre-acquisition year ($t-2$) at the 5% level. The post-acquisition price increases from this specification are also noticeably larger in magnitude than those from the simple difference-in-differences estimation. Without acquisition lags, the post-acquisition price trend is partially absorbed by the treatment-group specific year fixed effects. Once acquisition-year lags are included, these year fixed effects can more accurately reflect the negative trend in prices observed at yet-to-be-acquired hospitals.

[23]Note that the out-of-market and in-market leads and lags are estimated together but the estimates are reported in separate tables because of space.

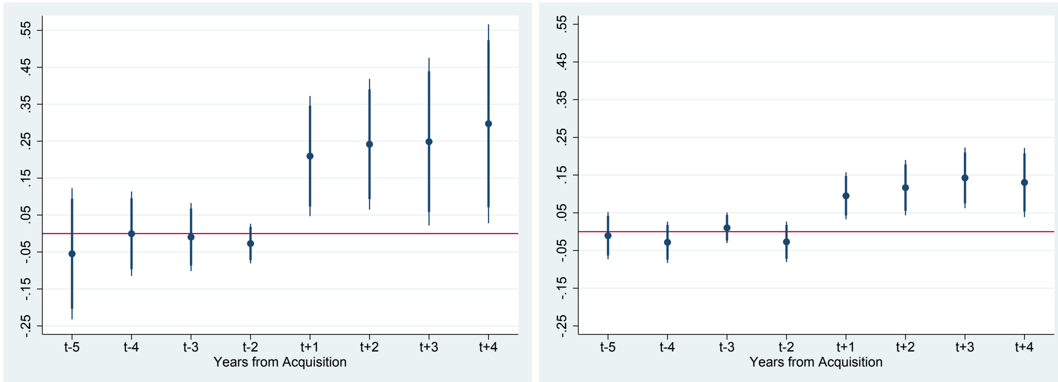
TABLE 6—LEAD AND LAG PRICE EFFECTS FOR IN-MARKET ACQUISITIONS

	(1)	(2)	(3)	(4)
Years Relative to T2 Acquisition				
t-6 or less	−0.010	−0.011	−0.033	−0.025
	(0.032)	(0.034)	(0.043)	(0.044)
t-5	−0.010	0.002	−0.024	−0.011
	(0.025)	(0.026)	(0.030)	(0.032)
t-4	−0.011	−0.014	−0.024	−0.028
	(0.020)	(0.023)	(0.024)	(0.028)
t-3	0.015	0.015	0.013	0.010
	(0.015)	(0.017)	(0.019)	(0.021)
t-2	0	0	0	0
	(.)	(.)	(.)	(.)
Acquisition Period				
t+1	0.077^c	0.073^c	0.106^c	0.095^c
	(0.023)	(0.026)	(0.030)	(0.032)
t+2	0.083^c	0.093^c	0.113^c	0.117^c
	(0.026)	(0.029)	(0.033)	(0.037)
t+3	0.120^c	0.115^c	0.162^c	0.143^c
	(0.032)	(0.033)	(0.041)	(0.041)
t+4	0.128^c	0.107^c	0.171^c	0.130^c
	(0.036)	(0.039)	(0.045)	(0.047)
t+5 or more	0.119^b	0.106^b	0.164^c	0.137^b
	(0.048)	(0.050)	(0.060)	(0.062)
Leads and Lags interacted with				
% Non-Medicare Private Share	No	No	Yes	Yes
Market-wide Acquisition Leads/Lags	No	Yes	No	Yes
Adj. R^2	0.475	0.475	0.475	0.476
N	19,009	19,009	19,009	19,009

Notes: A hospital first reports belonging to a system in year t, however, they may have joined anytime in the last year, between surveys. All specifications include hospital and treatment-specific year fixed effects as well as the controls reported in column 6 of Table 2. All estimates are relative to the omitted years, which are the year prior to joining a system and the first year the hospital was reported in a system as the acquisition took place sometime during these two years. Standard errors in parentheses are clustered by MSA. Significance Levels: a =p $< .10$, b =p$< .05$, c =p $< .01$

Table 6 reports the treatment effects for in-market acquisitions. These estimates also indicate that prices are relatively stable prior to acquisition and exhibit a discrete jump afterwards. Estimated acquisition price effects are slightly smaller for either type of acquisition when market-wide leads and lags are included. As in our standard triple-differences specification, this may reflect the fact that rival hospitals will also enjoy a price increase as a result of the acquired hospital increasing its price.

Figures 2(a) and 2(b) plot the point estimates of the treatment leads and lags and their respective 90 and 95 percent confidence intervals from column 4 of Tables 5 and 6, respectively. Given the lack of evidence for any pre-acquisition price effects and the large jump in prices that occurs surrounding the acquisition year, endogeneity in the selection of hospitals would require that systems anticipate idiosyncratic price increases that are about to occur at



(a) Out-of-market Acquisitions (b) In-market Acquisitions

FIGURE 2. LEAD AND LAG TREATMENT EFFECTS OF ACQUIRED HOSPITALS

particular independent hospitals and manage to acquire these hospitals precisely when these price increases are to occur. We believe such a sequence of events to be reasonably unlikely.

Overall, our estimates consistently suggest that hospitals acquired by out-of-market systems increased average prices by around 17 percent, with some specifications suggesting even larger increases. Though this effect is quite substantial, several previous studies have also found evidence that system hospitals enjoy a price premium of similar magnitude over non-system hospitals, even after controlling for differences in local market concentration. For example, Melnick and Keeler (2007) find that hospitals belonging to a large system enjoy prices that are about 34 percent higher than non-system hospitals, independent of the level of market concentration, and Lewis and Pflum (2015a) find that system hospitals in California have stronger bargaining power than non-system hospitals resulting in prices that are about 20 percent higher on average. Ho (2009) similarly finds that system hospitals have markups that are about $3,200 higher than non-system hospitals.[24] Although these studies rely on cross-sectional data and cannot explicitly identify a causal effect of system membership, our estimates suggest that much of their observed price differences are likely to have been a direct result of system membership, rather than simply representing some type of positive selection effect.

[24]Ho (2009) does not report an average price of a discharge in her study, but the average value from our sample is $14,200, suggesting that system hospitals had prices that were about 23% higher than non-system members having similar costs.

IV. Cost of Care and Profit Margins

The price regressions in the previous section include the average cost per discharge and the hospital's capacity as a measure of opportunity cost to assure that increases in observed reimbursement rates following a merger are not simply a result of an increase in the illness complexity of the patients they treat or a change in the cost effectiveness with which they treat those patients. As a result of including these cost controls, the treatment effects describe how the profit margins of these hospitals change when they are acquired by a system. If the cost efficiency of the hospital is relatively unaffected by the system acquisition, then an observed increase in profit margin largely translates to an increase in reimbursement rates. It is nevertheless possible that an increase in profit margin could be partially (or predominantly) driven by a reduction in costs, which would obviously generate very different policy implications. In fact, firms often try to justify proposed mergers by claiming that they will generate cost efficiencies.

A number of existing studies including Connor, Feldman, and Dowd (1998), Dranove and Lindrooth (2003), and Harrison (2011) investigate the impact of hospital consolidation on costs and generally find that hospitals exhibit significant cost savings post-merger. But these studies examine mergers occurring during the 1980s and 90s and mainly focus on cases in which two independent hospitals in the same patient market consolidate and continue operation as a single hospital. Despite this evidence, there are reasons to believe that the types of system acquisitions that we study and that have become increasingly common in the last several decades may not have the same impact on the cost efficiency of acquired hospitals. In fact, Dranove and Lindrooth (2003) separately examine pairs of local hospitals that merged to form systems but continued to operate as separate hospitals and found no evidence of post-merger cost savings.

We can investigate the extent to which observed increases in margins reflect efficiency improvements by using the same difference-in-differences approach to estimate the impact of a system acquisition on treatment costs.[25] The average cost of care per discharge becomes the dependent variable and controls are included for the hospital's average length of stay

[25] Note that we can determine if there is an efficiency increase but cannot determine whether the merger is necessary for any observed efficiency increase as required by the U.S. Horizontal Merger Guidelines (Section 10).

TABLE 7—THE IMPACT OF SYSTEM MEMBERSHIP ON HOSPITAL COST OF CARE

	ln(Cost/Dis.)		ln(CMI)		ln(Utilization)		ln(Avg. LOS)	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
T1	-0.058^{b}	-0.026	-0.004	0.001	-0.056	-0.056	-0.002	0.028
	(0.027)	(0.028)	(0.012)	(0.012)	(0.041)	(0.041)	(0.037)	(0.040)
T2	-0.021	-0.007	0.007	0.007	-0.006	-0.017	0.007	0.010
	(0.013)	(0.014)	(0.006)	(0.007)	(0.018)	(0.019)	(0.016)	(0.017)
Hospital or rival acquired by an...								
Out-of-market system		-0.034^{c}		-0.005		0.001		-0.033
		(0.011)		(0.005)		(0.013)		(0.021)
In-market system		-0.022^{c}		-0.001		0.016		-0.006
		(0.008)		(0.003)		(0.011)		(0.011)
Adj. R^2	0.496	0.498	0.070	0.070	0.080	0.080	0.102	0.102
N	19,009	19,009	16,028	16,028	19,009	19,009	19,009	19,009

Notes: All specifications include hospital and treatment-specific year fixed effects; specifications in the bottom panel additionally include treatment-group-specific year fixed effects; and columns 1 and 2 additionally include bed utilization, bed utilization squared, and the shares of Medicare and Medicaid patients as well as the out-patient share of revenues. Standard errors in parentheses are clustered by MSA. Significance Levels: a =p $< .10$, b =p$< .05$, c =p $< .01$

(which, to some extent, controls for changes in illness severity and a hospital's efficiency in providing treatment) and the proportions of patients covered by Medicare and Medicaid.[26] We estimate the cost regression using both difference-in-differences and triple-differences specifications. Finally, in addition to studying costs directly, we also separately investigate the impact of acquisition on several characteristics likely to impact hospital cost: patient case-mix, hospital utilization rate, and average length of stay. Treatment effect estimates from all specifications are reported in Table 7.

The hospital characteristic regressions in columns 3 through 8 reveal no significant evidence of changes in the case-mix of the patients, the average length of stay, or the utilization rate following system acquisition. With regard to overall cost levels, the estimated difference-in-differences treatment effects in column 1 suggest that average costs fall by nearly 6 percent when acquired by out-of-market systems and by about 2 percent when acquired by in-market systems. However, the triple-differences estimates in column 2 reveal that these acquisitions appear to be more likely to occur in markets where costs are increasing at a slightly lower rate, and after accounting for this, costs at acquired hospitals do not fall significantly relative to other hospitals in the same city following acquisition. Although the coefficient estimates of the impact of acquisition on costs remain negative, they are both statistically insignificant and quite small in magnitude relative to the acquisition-related in-

[26]Measures of payer type can help control for associated differences in illness complexity as well as the possibility that differences in reimbursement rates might influence the resources used to treat particular types of patients.

creases in profit margins identified in our price regressions. For robustness, we also adapt our leads and lags specification from Section III.B to analyze how costs evolve before and after hospitals are acquired.[27] The estimates confirm the findings of the specifications in Table 7, revealing no systematic change in costs relative to other hospitals within the same market around the time of acquisition.

Our results give no indication that acquisitions by multi-hospital systems occurring either within the same market and across markets produce any significant cost savings. The observed increases in profit margins following both in-market and out-of-market system acquisitions appear to be almost entirely due to increases in reimbursement prices. These findings provide an interesting contrast to existing studies that do find evidence of cost efficiencies following mergers of independent hospitals within the same local market.

V. Robustness and Alternative Explanations

Our results reveal important cross-market price effects of system formation that existing studies of hospital competition do not incorporate. In this section we perform several empirical tests to explore and verify the robustness of our findings.

A. Hospital Quality

Acquired hospitals may command higher post-merger reimbursement rates if they increase their quality of care or the quality of their facilities following acquisition. Unfortunately it is difficult to observe measures of perceived hospital quality so we cannot directly control for these potential changes like we can for changes in cost of treatment. Nevertheless, if acquired hospitals exhibit an increase in admissions despite the fact that prices have increased, this would be a fairly clear sign that the hospital improved its quality or became more attractive to patients. In light of this we adopt our standard difference-in-differences specification to examine the effects of system acquisitions on the quantity of care provided. We consider two different measures of hospital output—the total number of discharges and the number of privately insured patient discharges—as well as several measures of the hos-

[27] Difference-in-differences estimates of the evolution of costs before and after acquisition are plotted in Appendix Figures B1(a) and B1(b) and triple-differences estimates are plotted in Figures B1(c) and B1(d).

TABLE 8—THE IMPACT OF SYSTEM MEMBERSHIP ON DEMAND

	45 mile Market Share (1)	15 mile Market Share (2)	# Discharges (3)	# Private Discharges (4)	# Medicare Discharges (5)	Cost/Discharge (6)
T1	0.002	−0.009	−0.033	−0.064	−0.029	
	(0.032)	(0.022)	(0.042)	(0.054)	(0.039)	
T2	−0.010	−0.021	0.021	0.012	0.018	
	(0.015)	(0.015)	(0.019)	(0.027)	(0.023)	
# Rivals within 10 mi. acquired by an...						
out-of-market system						-0.056^{a}
						(0.032)
in-market system						-0.034^{b}
						(0.014)
Adj. R^2	0.027	0.033	0.080	0.059	0.097	0.493
N	19,009	19,009	19,009	19,000	19,006	14,264
	(7)	(8)	(9)	(10)	(11)	(12)
T1	0.002	−0.021	-0.077^{a}	−0.093	-0.075^{a}	
	(0.032)	(0.022)	(0.045)	(0.061)	(0.044)	
T2	−0.002	−0.017	−0.018	0.029	−0.036	
	(0.016)	(0.014)	(0.020)	(0.025)	(0.024)	
# Rivals within 10 mi. acquired by an...						
out-of-market system						−0.041
						(0.038)
in-market system						−0.015
						(0.013)
Adj. R^2	0.027	0.033	0.081	0.059	0.098	0.497
N	19,009	19,009	19,009	19,000	19,006	14,264

Notes: All dependent variables are in natural logs. All specifications include hospital fixed effects; specifications in the top panel additionally includes time fixed effects while the bottom panel includes treatment-group-specific time fixed effects; and columns 6 and 12 do not include treatment hospitals but additionally include bed utilization, bed utilization squared, and the shares of Medicare and Medicaid patients as well as the out-patient share of revenues. Standard errors in parentheses are clustered by MSA. Significance Levels: a =p $< .10$, b =p$< .05$, c =p $< .01$

pital's market share.[28] Market shares are calculated as the hospital's share of all discharges observed at hospitals within a 15-mile or 45-mile radius. Controls for patient care costs or patient mix are no longer necessary here because output measures are being used as the dependent variable. We still consider the same variety of time controls as in earlier specifications to control for other factors affecting hospital usage.

Table 8 reports the estimated treatment effects for each of the different measures of hospital output and market share. For both in-market and out-of-market acquisitions most of the coefficient estimates are statistically indistinguishable from zero and suggest a decrease in the number of privately insured patients. Based on these results, there is no evidence that acquired hospitals exhibited any substantial improvements in quality or became more attractive to patients.

[28]Note that in analyzing the effects of acquisition on costs we already examined whether an acquired hospital's utilization rate increases and, if anything, find that utilization decreases following acquisition.

Quality improvements could have occurred without generating an increase in admissions if the resulting demand increase was counteracted by increases in patient out-of-pocket costs or attempts by MCOs to steer patients away from these now higher priced hospitals. It is important to note, however, that any such increase in quality did not result from an increase in patient care spending. The acquisition-related price increases we observe are estimated conditional on patient care costs and the results of Table 7 reveal that the cost of care at acquired hospitals remained largely unchanged (or decreased slightly) following acquisition. Moreover, our rivals analysis suggests that higher post-acquisition profit margins did not result from hospitals providing higher quality at a lower cost.

Nearby rival hospitals could respond to quality improvements at acquired hospitals by increasing spending on patient care to maintain competitive quality levels, but there is little reason to expect these rival hospitals to become more efficient at producing higher quality after a nearby hospital is acquired by a system. If the main impact of system acquisition was to allow the acquired hospital to provide higher quality care more efficiently, we would expect the profit margins of nearby rivals to decrease (due to the presence of a more efficient competitor) rather than increase as the results of Table 4 show. Moreover, we might expect nearby rival hospitals to respond by spending more on patient care in response to a nearby acquisition. Columns 6 and 12 of Table 8 examine how patient care costs at rival hospitals respond to a nearby acquisition. The sample for these specifications include only the control hospitals, some of which represent rivals to acquired hospitals. The effects of having an acquired rival are fairly precisely estimated and suggest that the rivals to acquired hospitals do not exhibit an increase in their cost of care that would indicate an effort to compete with increased quality. This, together with the lack of a change in quantity or market share for acquired hospitals, suggests that the price increases are likely not driven by quality improvements.

B. Uncompensated Care

Our measure of average reimbursement price has the potential to be biased because hospitals provide some amount of care for which they never receive payment (i.e., uncompensated care). This distorts downward the calculated reimbursement per discharge. More importantly, if hospitals tend to provide less uncompensated care after being acquired this will cause us to overestimate the treatment effects.

Hospitals report the total value of uncompensated care they provide each year to the CMS but collection of this data only began in 2003. Therefore there are only 23 in-market acquisitions and 54 out-of-market acquisitions for which we observe levels of uncompensated care both before and after merger. Using our standard difference-in-differences approach we test whether system acquisitions are accompanied by significant changes in the amount of uncompensated care provided by acquired hospitals. The results from this estimation are reported in Appendix Table B1.

Although the estimates are relatively imprecise (due to the shorter sample period), some specifications suggest that acquired hospitals may reduce their provision of uncompensated care by 16%. However, given that uncompensated care represents a relatively small share of total discharges for the average hospital, these effects would at most translate into an overestimate of the average post-acquisition reimbursement rate by less than one percentage point—a fraction of the estimated treatment effects from our main analysis.

C. For-Profit Status Changes

Another potential source of bias could arise if price effects resulting from a change in the ownership type (i.e., for-profit status) of the hospital become confounded with acquisition effects. Some hospitals change status from non-profit to for-profit after acquisition. Failing to account for any corresponding change in objective could distort our estimates of the price increase directly attributable to system acquisition. We can control for changes in ownership status by including indicators identifying the current for-profit status of the hospital in our standard difference-in-differences specification. The estimates (reported in Appendix Table B2) provide some evidence that prices at hospitals in the out-of-market treatment group are somewhat higher when they operate as for-profit hospitals. Even when controlling for profit-status changes, however, the estimated price increase associated with an out-of-market acquisition falls only slightly from 17% to 15% for out-of-market acquisitions.

We also investigate whether the magnitudes of price increases following system acquisition tend to differ for for-profit and nonprofit hospitals. Our estimates indicate that prices increase about 5 percentage points more at nonprofits following an out-of-market acquisition and about 8 percentage points more following an in-market acquisition, though none of the for-profit results are statistically significant.

D. Medicare Reimbursement Rates and System Acquisition

All of our empirical findings support the assertion that there is a market power effect associated with out-of-market acquisitions. In light of this we implement one final falsification test by examining the effect of these acquisitions on a hospital's average reimbursement rate from Medicare patients. The Medicare reimbursement prices are administratively set and based on the average costs of providing care nationally and adjusted for case severity and geographic factors. As a result, a change in a hospital's bargaining power should not impact revenues from Medicare patients the way it does for privately insured patients.

For each hospital we observe both the total revenues for inpatient-care from Medicare patients and the total number of Medicare patient discharges and inpatient days. With this information we can construct accurate measures of average reimbursement rates for these patients. We estimate our standard difference-in-differences regression using Medicare prices as the dependent variable. The estimates are reported in Appendix Table B3.

The out-of-market treatment effects are all near zero. The estimated treatment effects for Medicare prices are statistically different from zero for in-market acquisitions; however, they are all of opposite sign and much smaller in magnitude than those estimated in the price regressions for privately insured patients. The absence of an increase in average reimbursement rates for Medicare patients suggests that the treatment effects we observe for privately insured patients are not simply the result of unobserved increases in the cost of treatment or the severity of the case mix of a hospital's overall patient population. These findings reinforce the robustness of the treatment effects identified for privately insured patients.

VI. Discussion and Possible Sources of Cross-Market Dependencies

Our empirical results reveal that many out-of-market mergers result in significantly higher prices for merging hospitals and their local rivals, suggesting that the competitiveness of individual patient markets may depend on factors beyond the local market structure. In this section we highlight several mechanisms capable of generating cross-market effects and explore some of the testable implications of these mechanisms.

Like most recent studies of hospital competition (including Capps et al., 2003; Lewis and Pflum, 2015a; Gowrisankaran et al., 2015), our discussion utilizes a Nash bargaining framework to directly model price negotiations between hospitals and MCOs, and an *option demand* approach to specify the gains from contracting.[29] Similar models have been used to study negotiated prices in other industries as well; including television content (Crawford and Yurukoglu, 2012) and coffee (Draganska, Klapper, and Villas-Boas, 2010). Let $\vec{p}_{sm}$ represent the vector of reimbursement prices paid by MCO m to the hospitals in system s; let $\Delta_s \Pi_m(\vec{p}_{sm})$ denote m's additional profit when it includes s in its provider network and pays reimbursements $\vec{p}_{sm}$; and let $\Delta_m \Pi_s(\vec{p}_{sm})$ denote s's additional profit when it is included in m's provider network and receives reimbursement payments $\vec{p}_{sm}$. Together the MCO and hospital system choose the price vector $\vec{p}_{sm}$ that maximizes the Nash product of the respective gains; i.e., the system and MCO together solve the optimization problem

$$(3) \qquad \max_{\vec{p}_{sm}} \left\{ [\Delta_m \Pi_s(\vec{p}_{sm})]^{\beta_{s(m)}} \times [\Delta_s \Pi_m(\vec{p}_{sm})]^{1-\beta_{s(m)}} \right\},$$

where $\beta_{s(m)}$ is the relative bargaining power of the system vis-à-vis the MCO and subject to the MCO and system being no worse off from contracting (i.e., subject to $\Delta_m \Pi_s(\vec{p}_{sm}) \geq 0$ and $\Delta_s \Pi_m(\vec{p}_{sm}) \geq 0$).

The resulting first-order conditions can then be expressed as

$$(4) \qquad \beta_{s(m)} \left[\frac{\partial \Delta_m \Pi_s}{\partial p_{hm}} \frac{p_{hm}}{\Delta_m \Pi_s} \right] = -(1 - \beta_{s(m)}) \left[\frac{\partial \Delta_s \Pi_m}{\partial p_{hm}} \frac{p_{hm}}{\Delta_s \Pi_m} \right]$$

showing that the equilibrium vector of prices equates the elasticity of the system's profit with respect to price to that of the elasticity of the MCO's profit with respect to price, weighted by their relative bargaining power ($\beta_{s(m)}$).[30]

Alternatively, the first order condition from (4) can be rearranged and is more com-

[29]Brooks et al. (1997) utilize a Nash bargaining framework with a monopoly hospital and MCO. Town and Vistnes (2001) utilize an option demand framework but not within the context of a bargaining game.

[30]Note that $\frac{\partial \Delta_s \Pi_m}{\partial p_{hm}} \leq 0$.

monly expressed in the literature in the form

$$\Delta_m \Pi_s(\vec{p}_{sm}) = \beta_{s(m)}[\Delta_s \Pi_m(\vec{p}_{sm}) + \gamma \Delta_m \Pi_s(\vec{p}_{sm})], \tag{5}$$

where $\gamma = \left| \frac{\partial \Delta_m \Pi_s}{\partial p_j} \Big/ \frac{\partial \Delta_s \Pi_m}{\partial p_i} \right|$. In equilibrium, the system's incremental profit from contracting with the MCO is the $\beta_{s(m)}$-share of the total gains from contracting, where the MCO's gains are weighted to account for how a change in reimbursement price differentially impacts the profits of the system compared to the MCO. Studies often assume that patients' hospital choice does not respond to the reimbursement rate, implying that γ is equal to one.[31] In this case the reimbursement is a simple transfer from MCO to hospital, and the ability of a system to negotiate a reimbursement rate above its costs depends on two key components: the bargaining power of the system ($\beta_{s(m)}$) and the total contract gains ($\Delta_h \Pi_m(\vec{p}_{hm}) + \Delta_m \Pi_h(\vec{p}_{sm})$).

Mergers can increase hospital market power whenever the gains generated by including a system of hospitals in the provider network are greater than the sum of the incremental gains that would have been generated by including each individual hospital in the provider network had they negotiated separately as standalone hospitals. In the spirit of Horn and Wolinsky (1988), Stole and Zwiebel (1996) and Inderst and Wey (2003), the potential anticompetitive effects of the merger are largely reflected in the degree of nonlinearity in the relationship between the incremental gains from contracting with each hospital individually and the gain from the contracting with the system as a whole.[32]

Although nearly all studies in the hospital competition literature utilize this same general framework, they often make different assumptions about the factors that impact the gains from contracting. For example, Capps et al. (2003) and many others assume that enrollees will remain with their MCO when hospitals are dropped from the provider network, but Ho (2009) and Ho and Lee (2013) allow enrollees to potentially switch MCOs in response to changes in provider network. In a theoretical study, Lee and Fong (2012) specify a dynamic model of network formation to consider the effects of allowing MCOs to potentially re-optimize the structure of its remaining provider network if it fails to contract with a specific

[31] One notable exception is Gowrisankaran et al. (2015), who explicitly allow patients facing coinsurance payments to respond to differences in out-of-pocket costs across hospitals.

[32] More generally, the incremental profit is concave in the size of the counter party, as Chipty and Snyder (1999) highlighted in their empirical investigation of bargaining outcomes in the cable television industry.

hospital.

Notably, all of these studies retain the assumption that an MCO's incremental profits are additively separable across different patient markets, implying that cross-market mergers cannot generate any market power effects. As Vistnes and Sarafidis (2013) point out, however, this may not be true given that the majority of private health coverage is purchased by employers rather than individuals. Employers arranging provider networks for employees located in different patient markets may place a higher value on having access to a geographically dispersed hospital system even if their individual employees do not. Such linkages could allow mergers of hospitals in different patient markets to significantly increase average prices, but only if a substantial number of residents in both markets share the same employers. We do not have the detailed employment data necessary to test this prediction directly. But this type of overlap is probably most likely to occur around major cities where large employers have employees living in different patient markets within the same broader metropolitan region or, perhaps, the same state. Therefore, we implement a more indirect test, empirically examining the importance of regional mergers by expanding the out-of-market cut-off distance to determine whether strong merger price effects are still present even when hospitals are located in entirely different regions.

Table B4 in the Appendix reports the treatment effects from specifications in which the minimum distance cut-off used to define an out-of-market acquisition is increased from 45 miles to 75 miles, and again to 90 miles. Even when considering only the 37 acquired hospitals in our sample that were more than 90 miles from their nearest system partner, there is strong evidence that prices increase following acquisition. In fact, these hospitals exhibit an average increase in price that is around 40% larger than that estimated using the 45 mile cut-off. As hospitals negotiate with insurers that often operate at a state-wide level, we additionally estimate a specification (in Appendix Table B5) which uses the standard 45 mile cut-off but allows the acquisition price effect to differ when the acquired hospital has no system partners in the same state. These *out-of-state* acquisitions generate price increases that are roughly 50% larger on average than system acquisitions involving a hospital that is in a different market but within the same state as an existing system partner. Both findings suggest that the identified price effects from acquisitions classified as out-of-market are not likely to be the result of regional employers purchasing common coverage for employees or of patient substitution over a broader geographic area.

It is still possible that very large nationwide or regional employers may have employees in entirely different regions or metro areas, but the degree of overlap in employers generally across any two cities is likely fairly small. Moreover, many of the systems in our sample that acquire an out-of-market hospital have members that are spread across the country in seemingly unrelated markets. Some of these (for example: Community Health Systems or Regional Care Hospital Partners) are made up of regional hospitals located in smaller cities where the presence of large national employers might be relatively limited.[33] These observations suggest that increased geographic coverage of large employers' employee base is not a likely driver of the increased market power from out-of-market acquisitions.[34]

Another way system membership can create a cross-market dependency is through a hospital's bargaining power ($\beta_{s(m)}$). There is strong evidence that relative bargaining power can vary, sometimes substantially, across hospital/MCO pairs (Lewis and Pflum, 2015a; Gowrisankaran et al., 2015). Lewis and Pflum (2015a) in particular find significant differences in bargaining power for system members compared to stand-alone hospitals. [35] Although bargaining power is represented by an exogenous parameter in the Nash model, several findings from the theoretical bargaining literature (including Fudenberg and Tirole (1983) and Sobel and Takahashi (1983)) suggest that bargaining power may depend on factors such as the amount of information one party has regarding the value of the contract to the other party. A hospital's bargaining power might therefore increase after joining a system if the system shares the costs of creating a larger and more skilled team of contract negotiators or if it pools information from previous contract negotiations giving the member hospital more information to use in the bargaining process.[36] As a result, system membership has

[33]As of 2015, the Regional Care Hospital Partners system included 8 hospitals spread across 7 states (AL, AZ, CT, IA, MT, OH, TX) and all located in smaller cities.

[34]It may, however, be an important source of market power arising from in-market acquisitions. Such acquisitions often involve hospitals that are not particularly close substitutes for individual patients but that may be close to employees working for the same large employer throughout the metropolitan area.

[35]Like much of the literature, however, Lewis and Pflum (2015a) assume that the incremental profit of contracting with a system is separable across patient markets so their estimated bargaining power differences could be capturing other cross-market linkages.

[36]For example, Tenet Healthcare—a national hospital system having 80 members in 2015—reportedly adopted a "national negotiating template and new technology to analyze payer-specific profit and loss data, giving negotiators ammunition during contract talks" (Colias, 2006).

TABLE 9—HETEROGENEOUS IMPACT OF OUT-OF-MARKET ACQUISITIONS ON PRICE/DISCHARGE

ln(Price/Discharge)	(1)	(2)
Size of acquiring system		
Bottom tercile (≤ 4 members)	0.166^b	
	(0.078)	
Middle tercile (> 4 and ≤ 26 members)	0.247^c	
	(0.094)	
Upper tercile (> 26 members)	0.286^c	
	(0.107)	
Size (# of beds) of acquired hospital		
Bottom tercile (≤ 58 beds)		0.332^c
		(0.076)
Middle tercile (> 97 and ≤ 184 beds)		0.226^b
		(0.101)
Upper tercile (> 184 miles)		0.211^b
		(0.101)
Adj. R^2	0.476	0.476
N	19,009	19,009

Notes: Each specification includes the same control variables reported in column 6 of Table 2 as well as hospital and year fixed effects. The cut-off for defining an out-of-market acquisition is 90 miles. Standard errors in parentheses are clustered by MSA. Significance Levels: a =p $< .10$, b =p$< .05$, c =p $< .01$

the potential to increase an acquired hospital's bargaining power and, consequently, its reimbursement price, regardless of whether its acquisition alters the concentration of hospitals in a local patient market.

We do not observe direct measures of bargaining power but larger systems are likely to have more experience and a greater informational advantage whereas smaller hospitals may be in a position to benefit most from joining a system given their relatively limited resources. To examine these potential relationships, we alter column 6 from Table 2 to allow the out-of-market treatment effect to vary based on either the size of the acquiring system or the size of the acquired hospital. We utilize the 90-mile cut-off to define out-of-market acquisitions in this case to ensure that these represent acquisitions in completely separate metropolitan areas. The estimates, reported in Table 9, reveal that acquisitions by smaller systems (of 4 or fewer hospitals) generate substantially smaller price increases than acquisitions by larger systems. Additionally, prices increase substantially more when the acquired hospital is relatively small.

These results are consistent with the notion that observed acquisition effects at least partially reflect improvements in bargaining power. In contrast, we found no evidence that prices were influenced by employers operating in multiple markets. Despite these suggestive results, we hesitate to draw strong conclusions at this point given the relative infancy of the theoretical literature and the possibility that there may be other unidentified mechanisms through which contract negotiations in one patient market could influence negotiations in other markets. Future theoretical and empirical work identifying and testing the relative importance of these mechanisms is likely to be particularly valuable, especially because the antitrust implications depend heavily on the mechanisms responsible for these cross-market dependencies. For example, a price increase generated by additional bargaining power more closely resembles a transfer of profits from MCOs to hospitals than a true restriction of competition generated by increased market power. Mergers that substantially increase hospital bargaining power can ultimately result in higher overall medical costs and insurance premiums as well as increased market power for nearby rivals, but it is unclear whether this would represent a violation of antitrust law.[37] In contrast, cross-market dependencies that influence the value of contracting with the hospital system would generate additional market power in much the same way as a within-market merger and would represent a more standard antitrust violation.

Regardless of the underlying mechanism, evidence of cross-market acquisition price effects suggests that existing models of competition, such as the option demand approach developed by Town and Vistnes (2001) and Capps et al. (2003) and used by antitrust authorities,[38] may need to be re-evaluated. These models could significantly underestimate the potential price impacts of mergers that do not involve close competitors. In addition, adjustments might be necessary to assure that assessments of local mergers are not biased in some way by excluding other relevant factors. For example, if some of these additional mechanisms tend to operate more strongly at hospitals in areas with high local market concentration, the current approach to merger simulation may incorrectly attribute higher prices to reductions in local concentration.

[37]To our knowledge this issue has not been legally challenged or faced careful debate within the scholarly literature.

[38]Dranove and Sfekas (2009) provide an overview of how these methods have been used in antitrust cases and Farrell et al. (2011) describes how the method is used in hospital cases specifically.

VII. Conclusions

Whereas previous studies have shown that merging hospitals can increase their reimbursement prices by reducing competition over patients, our results indicate that mergers can increase hospital market power even when they do not reduce competition within a patient market. We find that hospitals chosen for acquisition by out-of-market systems throughout the U.S. between 2000 and 2010 exhibited significant post-merger increases in reimbursement rates. Using a variety of specifications and auxiliary regressions we confirm that these price increases are not driven by changes in patient case-mix, quality of care, or the cost of providing care more generally. Discrete jumps and changes in the trends of prices at acquired hospitals around the year of the merger and associated price increases observed at nearby rival hospitals provide additional evidence that the identified price changes reflect the direct causal impact of acquisition. Overall, our findings reveal the existence of important cross-market dependencies that can allow some hospitals to strengthen their market power by affiliating with out-of-market systems. Further development and incorporation of these insights into existing models will be necessary to fully understand the competitive impacts of the wave of cross-market consolidation that has reshaped the hospital industry over the last several decades.

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APPENDIX A: HOSPITAL PRICE CALCULATION

The average price per discharge for a given year is calculated as follows. As HCRIS reports gross revenues for both inpatient and outpatient services, but reports only total contractual deductions, the net inpatient revenues are found by discounting the gross inpatient revenues by the amount implied by the contractual adjustments. That is, gross inpatient revenues are multiplied by $1-(\text{total contractual adjustments})/(\text{gross inpatient revenues} + \text{gross outpatient revenues})$. This generates an estimate for total net inpatient revenue, from which we subtract the total payments received from CMS for discharges from patients enrolled in Original Medicare (OM).[39] Lastly, this non-OM net-revenue is divided by the number of non-OM discharges to generate an average price per discharge. The following equation reports the calculation used to estimate a hospital average price per discharge.

$$\text{Discharge Price} = \frac{\big[[\text{Gross Inpatient Revenue} \times (1-\text{discount})] - \text{OM Payments}\big]}{\text{Non-OM Discharges}},$$

where

$$\text{discount} = \frac{\text{Total contractual adjustments}}{\text{Gross Inpatient Revenue} + \text{Gross Outpatient Revenue}}.$$

Net patient revenues come from line 3 of Worksheet G-2 for both Forms CMS-2552-96 and CMS-2552-10 and gross inpatient revenues are the difference between line 25 with lines 22 and 23 of Worksheet G-2, Form CMS-2552-96 for data before 2010; and the difference between line 28 with lines 25 and 26 of Worksheet G-2, Form CMS-2552-10 for 2010 data.

The OM payments represent all payments received from inpatient care provided to Original Medicare patients including patients' out-of-pocket costs, adjustments given for graduate medical education, cost of teaching physicians, special add-on payments for new technologies, and other pass-through costs. The Medicare payments come from line 16 of Worksheet E, Part A, line 17.1 of Worksheet E, Part B, line 17 of Worksheet E-3 Part I, and lines 23 and 30 of Worksheet E-3 Part II from Form CMS-2552-96 for data before 2010 and from line 59 of Worksheet E, Part A, line 24 of Worksheet E, Part B, line 18 of Worksheet E-3 Part I, and lines 17 and 31 of Worksheet E-3 Part II from Form CMS-2552-10.

Total contractual adjustments come from line 2 Worksheet G-3 for both forms CMS-2552-96 and CMS-2552-10. The non-OM discharges is the difference between line 5, column 6 with columns 3, 4, and 5 of Worksheet S-3, Form CMS-2552-96 for data before 2010; and the difference between line 7, column 8 with columns 5, 6, and 7 of Worksheet S-3,.

[39]Patients enrolled in a Medicare Advantage plan are members of a private MCO, which makes payments to the hospital on behalf of its members. Although these MCOs receive captiated payments from CMS, the MCOs negotiate rates with hospitals similar to how is done with a commercial insurance plan. These patients (and their associated payments) are treated the same as privately insured patients in our analysis.

Appendix B: Additional Specifications

Table B1—The Impact of System Membership on Uncompensated Care

$\ln\left(\frac{\text{Uncompensated Costs}}{\text{Total Costs}}\right)$	(1)	(2)	(3)	(4)
T1	0.202	−0.135	0.204	−0.163
	(0.336)	(0.500)	(0.342)	(0.498)
T2	0.026	0.010	0.029	0.006
	(0.102)	(0.150)	(0.102)	(0.150)
Additional patient shares	No	No	Yes	Yes
Year Fixed Effects				
Year Only	X		X	
Treatment×Year		X		X
Adj. R^2	0.077	0.080	0.078	0.082
N	5832	5832	5832	5832

Notes: Hospitals report to the CMS an annual measure of uncompensated care (gross charges for which they are not reimbursed), but the data is only available from 2003 to 2010. Due to this shorter sample period there are only 23 in-market acquisitions and 36 out-of-market acquisitions for which we observe levels of uncompensated care both before and after merger, however, we can still employ our standard difference-in-differences specification to gain some insight into how hospitals alter their level of uncompensated care following an acquisition. On average, non-Medicare patients represent around 50% of a hospital's discharges and for both treatment and control hospitals uncompensated care charges represent about 4% of the total charges. A 16% decrease in the share of uncompensated care represents about a 0.75 percentage point decrease in uncompensated care which would bias the estimated price for non-Medicare patients by $0.75/.5 = 1.5$ percentage points. All specifications also include hospital fixed effects. Standard errors in parentheses are clustered by MSA.

Significance Levels: a =p $< .10$, b =p$< .05$, c =p $< .01$

TABLE B2—THE IMPACT OF HOSPITAL OWNERSHIP STATUS ON PRICES

ln(Price/Discharge)	(1)	(2)	(3)	(4)	(5)
T1× Private Share	0.173^{c}	0.153^{c}	0.161^{c}		
	(0.058)	(0.056)	(0.056)		
T2× Private Share	0.109^{c}	0.104^{c}	0.106^{c}		
	(0.026)	(0.025)	(0.025)		
FP× Treatment Hospital		0.087^{b}	0.050	0.138^{b}	0.138^{b}
		(0.039)	(0.077)	(0.056)	(0.056)
FP× Control Hospital		−0.037	−0.074	−0.037	−0.037
		(0.033)	(0.071)	(0.033)	(0.033)
Nonprofit× T1				0.163^{c}	0.163^{c}
				(0.057)	(0.057)
Nonprofit× T2				0.112^{c}	0.112^{c}
				(0.025)	(0.025)
For-Profit× T1				0.106	0.106
				(0.083)	(0.083)
For-Profit× T2				0.025	0.025
				(0.059)	(0.059)
For-Profit Status×Year			X		X
Adj. R^2	0.475	0.476	0.476	0.476	0.476
N	19,009	19,009	19,009	19,009	19,009

Notes: Of the hospitals in our sample, 13 acquired by an out-of-market system and 23 acquired by an in-market system change to for-profit status after acquisition and 50 control group hospitals change to for-profit status. These specifications provide more insight into how non-profit and for-profit hospitals differ. All reported variables are interacted with the private share of non-Medicare patients. Each specification includes the same control variables reported in column 6 of Table 2 as well as hospital and year×treatment fixed effects. Standard errors in parentheses are clustered by MSA.
Significance Levels: a =p $<$ .10, b =p$<$ .05, c =p $<$ .01

TABLE B3—THE IMPACT OF SYSTEM MEMBERSHIP ON THE MEDICARE PRICE/DISCHARGE

ln(Medicare Price/Discharge)	(1)	(2)	(3)	(4)
T1	−0.003	−0.004	0.004	−0.006
	(0.018)	(0.019)	(0.018)	(0.020)
T2	-0.029^{b}	-0.028^{b}	-0.028^{b}	-0.029^{c}
	(0.014)	(0.011)	(0.012)	(0.011)
ln(Avg. Medicare Length of Stay)			0.409^{c}	0.409^{c}
			(0.069)	(0.069)
ln(Cost/Day)			0.207^{c}	0.207^{c}
			(0.024)	(0.024)
Year Fixed Effects				
Year Only	X		X	
Treatment×Year		X		X
Adj. R^2	0.513	0.513	0.578	0.578
N	18,857	18,857	18,856	18,856

Notes: The dependent variable is the log average revenue per discharge for Medicare patients. All specifications include hospital fixed effects and the indicated level of year fixed effects. Standard errors in parentheses are clustered by MSA.
Significance Levels: a =p $<$ .10, b =p$<$ .05, c =p $<$ .01

TABLE B4—THE EFFECT OF SYSTEM MEMBERSHIP ON THE PRICE/DISCHARGE USING ALTERNATIVE DISTANCE CUTOFFS

ln(Price/Discharge)	(1)	(2)	(3)	(4)	(5)	(6)
45 Mile Cutoff (81 out-of-market acquisitions)						
T1× Private Share	0.037	0.096^b	0.096^b	0.128^a	0.170^c	0.173^c
	(0.057)	(0.041)	(0.042)	(0.075)	(0.055)	(0.058)
T2× Private Share	0.067^b	0.099^c	0.096^c	0.060^a	0.102^c	0.109^c
	(0.026)	(0.020)	(0.019)	(0.036)	(0.025)	(0.026)
75 Mile Cutoff (51 out-of-market acquisitions)						
T1× Private Share	−0.022	0.064	0.060	0.084	0.152^b	0.160^b
	(0.064)	(0.052)	(0.052)	(0.090)	(0.067)	(0.071)
T2× Private Share	0.074^c	0.103^c	0.101^c	0.070^b	0.108^c	0.115^c
	(0.027)	(0.019)	(0.018)	(0.034)	(0.024)	(0.025)
90 Mile Cutoff (37 out-of-market acquisitions)						
T1× Private Share	−0.041	0.076	0.074	0.169^a	0.235^c	0.244^c
	(0.079)	(0.052)	(0.053)	(0.092)	(0.077)	(0.079)
T2× Private Share	0.072^c	0.100^c	0.098^c	0.065^a	0.104^c	0.111^c
	(0.026)	(0.018)	(0.018)	(0.034)	(0.023)	(0.024)
Year Fixed Effects						
Year Only	X	X	X			
Trmt×Year				X	X	X
N	19,009	19,009	19,009	19,009	19,009	19,009

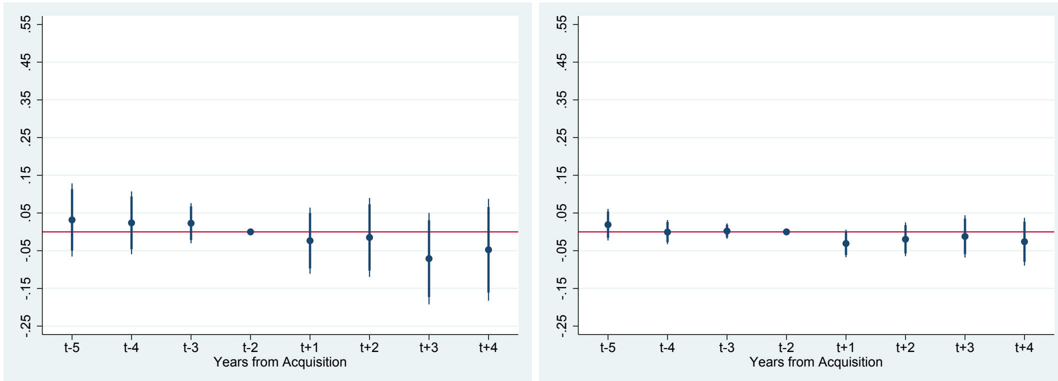
Notes: All specifications include hospital and year fixed effects and the other control variables included in the same numbered specifications from Table 2. Standard errors in parentheses are clustered by MSA. Significance Levels: a =p < .10, b =p< .05, c =p < .01

TABLE B5—OUT-OF-MARKET ACQUISITIONS BY IN- AND OUT-OF-STATE SYSTEMS

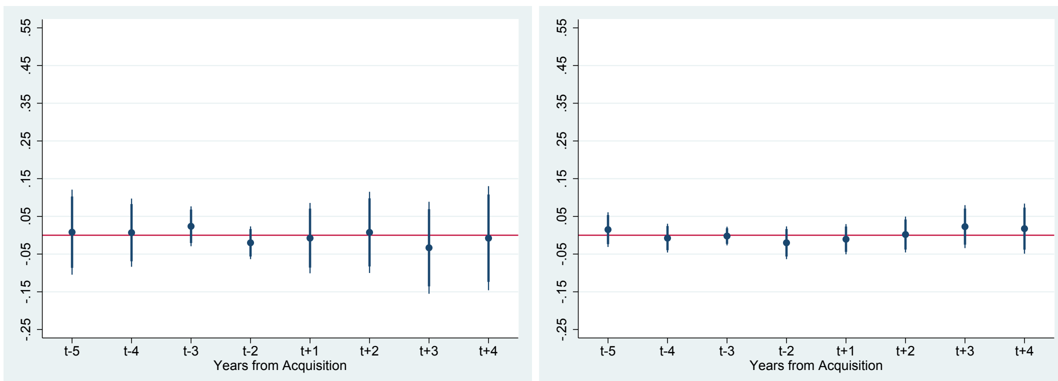
	Pure Treatment Effect		Interacted with Private share	
ln(Price/Discharge)	(1)	(2)	(3)	(4)
T1×...				
has no in-state partners	0.214^c	0.189^c	0.278^c	0.245^c
	(0.044)	(0.044)	(0.056)	(0.056)
has in-state partners	0.122^c	0.129^c	0.167^c	0.173^c
	(0.047)	(0.047)	(0.057)	(0.057)
Additional Controls	No	Yes	No	Yes
Adj. R^2	0.471	0.475	0.471	0.475
N	19,009	19,009	19,009	19,009

Notes: Each specification includes hospital and treatment-specific year fixed effects. Additional Controls indicates whether the specification additionally includes the controls included in specifications (3) and (6) from Table 2. Significance Levels: a =p < .10, b =p< .05, c =p < .01

FIGURE B1. LEAD AND LAG TREATMENT EFFECTS ON THE COST OF CARE FOR ACQUIRED HOSPITALS



(a) Out-of-market Acquisitions (b) In-market Acquisitions



(c) Out-of-market Acquisitions (d) In-market Acquisitions

Notes: All figures are based on specifications that include hospital and year fixed effects and the other control variables included in specification (6) from Table 2. The bottom two figures additionally include rival lead and lag treatment effects.

TABLE B6—CLOSENESS OF SUBSTITUTES AND MARKET POWER FROM IN-MARKET ACQUISITIONS

	Pure Treatment Effect		Interacted with Private share	
ln(Price/Discharge)	(1)	(2)	(3)	(4)
T2×closest partner...				
$\leq$ 2 miles away	0.107^c	0.115^c	0.159^c	0.169^c
	(0.033)	(0.031)	(0.044)	(0.040)
between 2 and 5 miles away	0.065	0.068	0.107^a	0.110^a
	(0.055)	(0.057)	(0.063)	(0.065)
$>$ 5 miles away	0.076^c	0.076^c	0.104^c	0.102^c
	(0.022)	(0.022)	(0.026)	(0.026)
Additional Controls	No	Yes	No	Yes
Adj. R^2	0.471	0.475	0.471	0.475
N	19,009	19,009	19,009	19,009

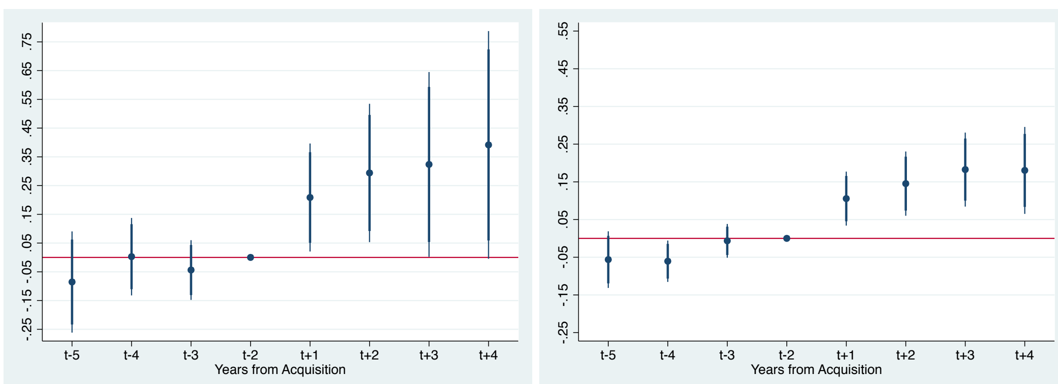
Notes: Each specification includes hospital and treatment-specific year fixed effects. Additional Controls indicates whether the specification includes the additional controls included in specifications (3) and (6) from Table 2. Significance Levels: a =p $<$ .10, b =p$<$ .05, c =p $<$ .01

TABLE B7—THE EFFECT OF SYSTEM MEMBERSHIP ON THE PRICE/DISCHARGE CONTROLLING FOR HOSPITAL-SPECIFIC TIME TRENDS

	Pure Treatment Effect			Interacted with Private Share		
ln(Price/Discharge)	(1)	(2)	(3)	(4)	(5)	(6)
T1	0.071	0.090^{b}	0.083^{a}	0.090	0.113^{b}	0.114^{a}
	(0.064)	(0.046)	(0.047)	(0.077)	(0.057)	(0.060)
T2	0.006	0.043^{b}	0.045^{b}	0.013	0.056^{b}	0.064^{c}
	(0.027)	(0.020)	(0.020)	(0.033)	(0.024)	(0.024)
ln(Cost/Discharge)		1.254^{c}	1.167^{c}		1.254^{c}	1.166^{c}
		(0.043)	(0.045)		(0.043)	(0.045)
For-Profit			0.022			−0.010
			(0.022)			(0.027)
% Medicare			0.384^{b}			0.386^{b}
			(0.186)			(0.186)
% Medicaid			−0.048			−0.046
			(0.073)			(0.073)
% OP Revenue			-0.879^{c}			-0.880^{c}
			(0.213)			(0.213)
ln(# Beds)			−0.006			−0.006
			(0.030)			(0.030)
Bed Utilization			0.218^{c}			0.218^{c}
			(0.067)			(0.067)
Bed Utilization Sqrd.			−0.041			−0.041
			(0.031)			(0.031)
Unemployment			−0.002			−0.002
			(0.003)			(0.003)
Poverty rate			0.001			0.001
			(0.002)			(0.002)
Median income (/1000)			0.016			0.016
			(0.016)			(0.016)
% Population over 65			0.666^{a}			0.660^{a}
			(0.389)			(0.388)
Adj. R^2	0.397	0.607	0.618	0.397	0.607	0.618
N	19,009	19,009	19,009	19,009	19,009	19,009

Notes: All specifications include hospital and year fixed effects as well as hospital-specific time trends. Standard errors in parentheses are clustered by MSA. Significance Levels: a =p $<$ .10, b =p$<$ .05, c =p $<$ .01

FIGURE B2. LEAD AND LAG TREATMENT EFFECTS OF ACQUIRED HOSPITALS CONTROLLING FOR HOSPITAL-SPECIFIC TIME TRENDS



(a) Out-of-market Acquisitions (b) In-market Acquisitions

Notes: All figures are based on specifications that include hospital-specific time trends and the other control variables included specification (6) from Table 2.