Hospital Market Competition and Participation in Value-Based Payment Programs

Yucheng Hou (UTHealth), Kun Li (Amazon), M. Kate Bundorf (Duke)

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The US hospital market has become highly concentrated through extensive vertical and horizontal consolidation over recent decades. Meanwhile, public and private payers also increasingly rely on value-based payment (VBP) models to contain costs while improving the quality of care. While extensive evidence has documented that provider consolidation raises the *level* of health care prices, little is known about whether it may also affect the *form* of prices (VBP vs. fee-for-service). Hospitals with market power may resist contracts that expose them to more financial risk, potentially deterring VBP participation. Alternatively, hospitals with market power may have more favorable cost structures or better control over patient care flow, which allows them to perform well under VBP models. In this paper, we consider the relationship between hospital market concentration and the adoption and performance of value-based payments in the context of Medicare's Comprehensive Care for Joint Replacement Model. Study findings may reveal a new channel for how increasing horizontal and vertical provider consolidation is affecting health care markets.

I. Introduction

Policy makers have focused on changing how insurers pay providers as a mechanism to control rising health expenditures while maintaining or improving health care quality and outcomes. Under fee-for-service (FFS) systems, the dominant form of payment in the U.S., health care providers are paid based on how many services they provide to patients. Value-based systems, in contrast, are intended to link payments to service quality and patient outcomes with the objective of creating incentives more directly aligned with value than volume. While many forms of value-based payment exist, they generally incorporate three features: 1) incorporating metrics of quality and outcome in payment systems to financially reward quality, 2) broadening the scope of providers included under a single payment to create incentives for greater coordination of care, and 3) shifting the financial risk for health care use from insurers to providers to promote more efficient use of services. While data from the Health Care Payment Learning and Action Network (HCPLAN) suggests that the use of alternative payment models is growing, with a shift toward models which incorporate broader population-based payments with significant downside risk, adoption is far from complete with many providers participating in contracts covering a narrow scope of services and exposing providers to limited financial risk (HCPLAN, 2024).

At the same time, markets for health care delivery in the U.S. have become highly concentrated through extensive vertical and horizontal consolidation over recent decades. In the case of hospitals, horizontal consolidation has occurred through mergers between local hospitals within markets, the acquisition of independent hospitals by large systems, and cross-market mergers between systems operating in different markets. Similar dynamics have affected the market for physician services – physician organizations have merged with each other and have been acquired by hospitals and health systems, generating both horizontal and vertical consolidation. KFF estimates that one or two health systems control the entire market for inpatient hospital care in nearly half of metropolitan areas (KFF, 2024), and the GAO reports that the percentage of physicians consolidated with health care systems rose from less than 30% in 2012 to at least 47% in 2024. In addition, technological change has shifted the site of many services from inpatient to outpatient care, increasing the importance of rising vertical integration between inpatient and outpatient care for health care delivery.

Research on the effects of hospital market consolidation has largely focused on the effects on price levels and, to a lesser degree, on patterns of health care use and outcomes. A large body of literature documents that horizontal consolidation among hospitals increases the prices of inpatient services (Liu et al., 2023). Cooper et al. (2019) report that prices at monopoly hospitals are 12% higher than those in markets with four or more competitors. Similarly, evidence indicates that vertical integration between hospitals and physicians leads to increased prices and higher spending (GAO, 2025), Studies have also documented that vertically integrated organization can also change the site of both inpatient and outpatient care, leading to higher spending with no evidence of improved quality (L. C. Baker et al., 2016; Richards et al., 2022).

Provider market consolidation may affect the form of pricing through two possible mechanisms. First, larger and more integrated organizations may respond more effectively to the incentives created by value-based payments for health care delivery. Horizontal consolidation may lead to

higher volume which may generate economies of scale in production and allow organizations to more effectively bear risk. Vertical consolidation may generate economies of scope and greater coordination across care settings. In both cases, larger, more integrated organizations may more effectively reduce costs and provide care more efficiently in response to payment incentives and also be more effectively positioned to bear financial risk. On the other hand, larger, more integrated organizations may have greater market power relative to insurers than smaller organizations, providing them with greater leverage in negotiations that determine both the level and form of pricing. Depending on their preference for bearing risk and their likely performance on value-based contracts, they may more effectively tailor their participation in value-based programs based on their organizational characteristics. In addition, they may use their negotiating power to obtain more favorable compensation overall, such as more generous population-based payment levels.

Research examining the relationship between hospital market structure and hospital participation in prospective payment in the U.S. provides evidence that hospital market structure can affect payment form. While Medicare pays nearly exclusively by admission (DRG), payment form for inpatient services in the commercial market is determined through negotiations between insurers and providers. In this setting, evidence indicates that hospitals with greater market power are less likely to participate in contracts which involve greater financial risk. Baker et al. (2019) find that the extent of prospective payment in commercial insurance is lower and more variable across hospitals and geographic areas than in Medicare and is positively associated with hospital market competitiveness as well as the share of admissions paid by commercial managed care plans, and the share paid by the Medicare Prospective Payment System. Similarly, Cooper et al. (2019) find that hospitals prefer contracts that involve less risk and place less pressure on them to reduce costs. Monopoly hospitals are more likely to have contracts that transfer financial risk of utilization to insurers.

Little evidence exists on whether changes in market structure have influenced the adoption of alternative payments models by providers. Research on the effects of consolidation has focused more on the level than the form of prices paid to providers. And, while providing important evidence that payment form is related to provider market competition, studies of hospital prospective payment in commercial markets may not necessarily reflect the experience of broader VBP initiatives. Inpatient prospective payment is limited in scope – only affecting payments for inpatient services. In addition, prospective payment does not incorporate quality metrics. In addition, this research highlights important difference in the institutional settings. While one study documents that physician group size and organizational complexity are positive associated with participation in APMs, a physician organization's self-reported intensity of competition in the outpatient setting is not significant associated with APM participation (Ouayogodé et al., 2020). Most closely related to your paper, Chen and Delana (2026) examines how physicians are organized, both horizontally joining large practices and vertically being employed by hospitals or health systems, drives heterogeneity of the effect of the Medicare's Comprehensive Care for Joint Replacement (CJR) Model. They find that CJR hospitals with fully horizontally and vertically integrated orthopedic surgeons lead to a significant increase in Medicare payments to hospitals and greater CJR reward payments. However, whether and how

hospital market concentration interacts with APM participation and performance remains unanswered.

In this paper, we examine the net effect of hospital market concentration on APM participation and performance. We focus on the case of Medicare's Comprehensive Care for Joint Replacement (CJR) Model. Initiated in 2016, CJR is a national episode-based payment model for lower-extremity joint replacement where hospitals in randomly assigned 67 Metropolitan Statistical Areas (MSAs) were required to participate. In 2018, Medicare made hospital participation voluntary in half of the MSAs. This has two advantages for our work. First, we are able to access comprehensive, publicly available data on hospital participation in this program from CMS. Although this is only one of many models, we feel confident in our ability to measure participation in this model. Second, the design of the program has some attractive features for our study design. In particular, we leverage the exogenous shift from mandatory to voluntary participation to examine how hospital market competition is associated with participation decisions.

We link hospital-level CJR reconciliation files to the American Hospital Association Annual Survey and other Centers for Medicare & Medicaid Services (CMS) hospital public use files from 2016-2022. Hospital market concentration is measured using a hospital-specific Herfindahl Hirschman Index (HHI) based on Medicare patient origins (zip codes from which the hospital admits Medicare patients). A notable empirical challenge is that observed hospital HHI is likely endogenous. Hospitals in more concentrated markets may differ from hospitals in less concentrated markets in unobserved ways that also affect VBP performance. Moreover, participation in VBP contracts may itself affect consolidation, referral patterns, and market structure. Following the logic of Kessler & McClellan (2000), we construct geography-predicted hospital HHIs based on hospital choice driven by distance and use them as an instrument for observed hospital HHIs. Specifically, we estimate a set of conditional logit models of hospital choice with hospital-level fixed effects in each hospital referral region and year, and construct predicted choice probabilities using distance-related coefficients. The predicted HHI, therefore, captures plausibly exogenous variation in market concentration generated by patient geography and the spatial distribution of hospitals, rather than by hospital unobserved quality or attractiveness.

We first examine the cross-sectional association between baseline market structure (2016-2017) and hospitals' decisions to remain in CJR in 2018, among MSAs where participation became voluntary, adjusting for hospitals' CJR volumes, Medicare patient mix, ownership, rural status (per the CJR definition), and MSA fixed effects. We did not find evidence that hospitals with higher baseline HHI have differentially remained in CJR than those with lower HHI. In contrast, baseline CJR performance, measured by receipt of reconciliation bonuses during the mandatory period, was a strong predictor of staying in the program (15 pp.; $p<0.01$), suggesting that realized performance, rather than market structure, primarily drives voluntary participation.

Since voluntary participation is strongly driven by realized performance, we next examine whether hospital market structure predicts CJR performance for hospitals in always mandatory MSAs where selection is not a concern. We estimate the longitudinal relationship between

lagged observed hospital HHI and reconciliation outcomes from 2016-2022; lagged observed hospital HHI each year was instrumented with corresponding lagged geography-predicted HHI. Our instrumental variable (IV) estimates show that a one standard deviation increase in lagged hospital HHI leads to a 5.7 percentage point increase in the probability of receiving a CJR bonus ($p<0.05$) but do not significantly differ in CJR quality measures.

To shed light on potential mechanisms through which hospitals in more concentrated markets achieve better financial performance under CJR, we examine target prices and episode complexity, measured by the share of episodes with major complications or comorbidities (MCC). Our 2SLS results suggest that higher target prices may contribute to this financial advantage during the period when hospital-specific historical spending was incorporated into the benchmark. We did not find evidence that market concentration affects composition of episode complexity.

Our preliminary results contribute to the literature on APM performance and hospital market competition. We extend the evidence that hospital market power may affect not only the level of prices but also a hospital's performance when it is subject to an alternative form of price, namely episode-based reconciliation payments under administratively set Medicare prices. These results are relevant to policymakers designing VBP programs as provider integration continues to grow. VBP, while designed to reduce costs and improve quality, may not offset the consequences of provider market power. To the extent that market power shapes hospitals' ability to benefit from risk-based payment, the shift toward mandatory VBP may reduce participation selection while still amplifying differences in financial performance across hospitals with different market power.

II. Policy Environment

The CJR model is one of the most high-profile value-based payment initiatives implemented by the Centers for Medicaid and Medicare Services (CMS). Launched in 2016, the CJR defines an episode of care for a joint replacement, beyond the hospital admission, and bundles the payments to the various providers involved in care during the episode into a single retrospective payment. The episode begins with the joint replacement procedure and extends for 90 days after discharge. All Medicare spending (Parts A and B) within those 90 days is included in the bundle. The hospital is ultimately responsible for the total cost and quality of care during the episode. The model initially focused on inpatient hip and knee replacements and was subsequently extended to ankle replacements and some outpatient joint replacements, in response to the removal of these procedures from the inpatient only list in 2018 (knee replacements) and 2020 (hip replacement). In a retrospective payment model, providers are initially paid under the traditional Medicare fee-for-service system. Medicare retrospectively calculates total episode spending and then compares it to a hospital-specific target payment. If spending is below target and the hospital meets quality standards, the hospital receives a bonus (reconciliation payment). If spending exceeds the target, the hospital may have to make payments to CMS (reconciliation penalty). The target payment initially combined historical hospital-specific and regional spending, with increasing weight

placed on regional spending until the historical spending component became entirely regional beginning in performance year 4 (2019).

CJR was initially implemented as a mandatory model with 67 metropolitan areas randomly selected for mandatory CJR participation. Finkelstein et al (2018), evaluating the first year of the program, found that hospital reduced use of institutional post-acute care and the Medicare spending on post-acute care fell for approximately $307 per episode. While overall spending was about $453 lower in CJR markets, the reduction was not statistically significant. Similarly, Barnett et al (2019), who examine the first two years of the program, similarly document that the reduction in spending was generated primarily by the use of institutional post-acute care and estimate a reduction of approximately $812 per episode in institutional spending relative to control areas with no evidence of a reduction in quality. Overall, the literature indicates that the CJR generated modest reductions in spending primarily through reductions in post-acute care which were partially offset by bonus payments with no evidence of a reduction in quality.

Starting in February 2018, CMS substantially scaled back mandatory participation. Among the original 67 participating markets, only 34 remained mandatory with participation voluntary in the other 33 markets. Many low-volume and rural hospitals were also exempted in mandatory markets. Examining the shift to voluntary participation among some but not all hospitals, Einav et al (2022) demonstrate that participation under a voluntary regime is more likely for hospitals that can increase revenue without changing behavior. They differentiate between "selection on levels," in which hospitals that are likely to perform well under the VBP program are more likely to participate in a voluntary regime, and "selection on slopes," which describes changes in behavior that generated favorable performance based on behavior change under a mandatory regime. As of October 1, 2021, the final rule that extends the CJR model for an additional three performance years only applies to hospitals in one of the 34 required MSAs and not designated as low volume or rural.

III. Data and Methods

Data

We use hospital-level CJR reconciliation files linked to the AHA Annual Survey and hospital public use files for 2016-2022, including hospital cost reports from the Healthcare Cost Report Information System (HCRIS), hospital Medicare patient origins from the Hospital Service Area files, and Medicare inpatient Public Use Files (PUF).

The Hospital Service Area files report annual Medicare fee-for-service and Medicare Advantage (MA) discharge counts by patient residence ZIP code and hospital (identified by CMS certification number [CCN]), allowing us to observe realized patient flows from residence ZIP codes to hospitals. We use these files to construct ZIP-to-hospital flows for each year from 2016 through 2022 and estimate hospital choice models. The hospital cost report, AHA, and Medicare

inpatient PUFs are used to identify the universe of Medicare-certified hospitals and to obtain hospital ZIP codes and other hospital-level characteristics.

Study Sample

Our analysis focuses on general acute care hospitals located in the selected MSAs (voluntary and mandatory) under the CJR model. We restricted the analyses to hospitals with at least 11 CJR *final* episodes. Health system affiliation was obtained from the AHA data each year. We link the hospital-specific market concentration measures (described in detail below) with the CJR hospital-year analytic file by hospital CCN and year. Two samples were developed for different parts of our analysis: (1) hospitals in MSAs that switched from mandatory to voluntary in 2018; (2) hospitals in always-mandatory MSAs.

Measures

Measuring Hospital Market Concentrations

Hospital market concentration is measured using HHIs constructed from longitudinal Medicare patient origin flows. We use the annual ZIP-to-hospital discharge flows from CMS Hospital Service Area files to identify the ZIP codes of residence from which each hospital admits Medicare patients. We convert both patient residence ZIP codes and hospital ZIP codes to ZIP Code Tabulation Areas (ZCTAs) and measure distance as the distance between the centroids of the patient residence ZCTA and the hospital ZCTA.

Following prior research, we do not define hospital markets by imposing fixed administrative boundaries. Instead, for each patient residence ZCTA, we define the potentially relevant inpatient hospital market as the set of all eligible general acute care hospitals (with at least 11 CJR inpatient discharges; of note, this is broader than 11 CJR final episodes), including critical access hospitals, located within 35 miles of the residence. This choice set includes eligible hospitals with zero observed flows from the residence ZCTA, so that the market reflects feasible hospital alternatives rather than only realized admission patterns. Residence ZCTAs for which all observed discharges go to hospitals outside the 35-mile choice set are excluded. We also test an alternative market boundary that includes hospitals within a 50-mile radius of the residence.

Let D_{kjt} denote observed discharges from residence ZCTA k to hospital j in year t, and let $D_{kt} = \sum_{\ell \in C_{kt}} D_{k\ell t}$ denote total observed discharges from residence ZCTA k across hospitals in the choice set. Observed hospital shares within each residence ZCTA are $s_{kjt}^{obs} = \frac{D_{kjt}}{D_{kt}}$.

To account for hospital system affiliation, we combine discharges to hospitals affiliated with the same system within each residence ZCTA. Let D_{kgt} denote observed discharges from residence ZCTA k to hospitals in system g in year t among hospitals in the choice set. The observed

ZCTA-level HHI is the sum of squared shares across these systems; we treat each non-system affiliated hospital as a separate entity: $HHI_{kt}^{obs} = \sum_g \left(\frac{D_{kgt}}{D_{kt}}\right)^2$. We then aggregate this system-adjusted ZCTA-level HHI to the hospital-year level by weighting each residence ZCTA by its contribution to hospital j's observed discharges. Hospital j's observed HHI is:

$$HHI_{jt}^{obs} = \sum_k \left(\frac{D_{kjt}}{D_{jt}}\right) HHI_{kt}^{obs}$$

where $D_{jt} = \sum_k D_{kjt}$ is hospital j's total observed discharges.

Dependent Variables

Key outcome variables include whether the hospital remained in the CJR program after half of MSAs switched from mandatory to voluntary in 2018, and whether the hospital received a positive reconciliation payment (CJR bonus) for 2016-2022. For quality of care, we use the CJR-required quality measures, including risk-standardized complication rates for elective primary total hip arthroplasty and total knee arthroplasty as well as a hospital patient experience measure in the Hospital Consumer Assessment of Healthcare Providers and Systems (HCAHPS) survey. The complication rate is aggregated over a three-year rolling window; The HCAHPS data is aggregated over a year.

To examine potential mechanisms, we also consider hospital target prices and episode complexity. We measure this outcome as the mean target price across CJR defined DRGs within each performance year and restrict the analysis to the earlier years (PY1-PY4) when benchmarks incorporated hospital-specific historical spending prior the transition to fully regional benchmarks. We measure episode complexity as the share of CJR discharges with MCC.

Control Variables

The CJR reconciliation file provides information on total CJR episode counts and rural status (the CJR definition). We also obtain additional patient mix variables from the Medicare inpatient PUFs, including average clinical risk (Hierarchical Condition Category [HCC] score), average patient age, and the share of dually eligible beneficiaries. Hospital ownership information is obtained from the hospital cost report.

Methods and Models

Remain in the CJR program when participation becomes voluntary

For this analysis, we restrict our sample to hospitals located in voluntary MSAs. Our basic specification is a linear probability model comparing the predicted probability of staying in the CJR program across hospitals with varying baseline market concentration when CJR participation became suddenly voluntary (2018).

$$Y_{j,m,t=2018} = \beta_0 + \beta_1 HHI_{j,t=2016-2017} + \mu_m + X_j\beta + \varepsilon_{jm}$$

Where $Y_{j,m,t=2018}$ is hospital j's participation in CJR in voluntary MSA_m and year t. μ_m is MSA fixed effects. X_i includes hospitals' Medicare patient mix, CJR rural status, and receiving positive reconciliation payment during 2016-2017. Standard errors are clustered at the MSA level.

CJR performance when participation is always mandatory

For this analysis, we restrict our sample to hospitals located in mandatory MSAs so that hospital self-selection into the CJR program is no longer a concern. A naïve model of the effect of hospital HHI on CJR performance faces the problem that observed hospital concentration may be endogenous. Hospitals with higher observed HHI may differ systematically in unobserved quality, referral relationships, system affiliation, patient mix, or other market characteristics that also affect CJR performance.

To address this endogeneity, we instead use the Kessler-McClellan style geography-predicted HHI as an instrument for the observed hospital-specific HHI in the CJR mandatory MSA sample. The geography-predicted HHI isolates the component of concentration generated by patient location and the spatial distribution of hospitals; we describe in detail how we constructed the longitudinal geography-predicted HHI in the Instrument section below. The identifying assumption is that geography-predicted HHI affects CJR performance only through its effect on observed hospital-specific market concentration, conditional on CJR volume, hospital patient mix, MSA fixed effects, and secular trend.

Our first-stage equation is

$$HHI^{obs}_{j,t-1} = \pi HHI^{geo}_{j,t-1} + X'_{jt}\delta + \mu_m + \tau_t + \nu_{jt}$$

where j, t, m index hospital, year, and MSA, respectively. X_{jt} is a vector of hospital and patient-mix controls, $\mu_m$are MSA fixed effects, and $\tau_t$are year fixed effects.

The second-stage equation is

$$Y_{jt} = \theta\widehat{HHI}^{obs}_{j,t-1} + X'_{jt}\gamma + \mu_m + \tau_t + \varepsilon_{jt}$$

The dependent variable $Y_{jt}$is an indicator for whether hospital $j$received a positive reconciliation payment in year t. The endogenous variable is lagged observed hospital-specific HHI. The

instrument is lagged geography-predicted hospital-specific HHI. In the main specification, both HHI variables are standardized, so the coefficient on observed HHI can be interpreted as the effect of a one-standard-deviation increase in observed hospital-specific concentration, using only geography-induced variation in concentration.

Instruments

Hospital Choice Model

Using the CMS hospital service area file linked to the hospital cost report, we estimate a set of conditional logit models of hospital choice for the patient's residence ZCTA, separately by year and by the patient's residence HRR. Let k index patient residence ZCTAs, j index hospitals, t index years, and $h(k)$ denote the HRR containing residence ZCTA k. We note that the HRR is defined by the patient's residence, not by the hospital's location; hospitals in a residence ZCTA's choice set may lie outside the patient's HRR.

For patient $i$residing in ZCTA k, utility from choosing hospital $j \in C_{kt}$ is

$$U_{ikjt} = \alpha_{jht} + \beta_{1ht} Distance_{kjt} + \beta_{2ht}\left(Distance_{kjt} \times Nonmetro_k\right) + \varepsilon_{ikjt}$$

Here, α_{jht} is a hospital fixed effect in the HRR-year estimation sample, $Distance_{kjt}$ is the distance from the centroids of residence ZCTA k to hospital j, and $Nonmetro_k$is an indicator for whether residence ZCTA $k$is in a nonmetropolitan area. The interaction term allows us to capture systematic differences in patient travel patterns between urban and rural areas.

Assuming that ε_{ikjt}follows a type 1 extreme value distribution, the probability that a patient in residence ZCTA k chooses hospital j is

$$P_{kjt} = \frac{\exp\left[\alpha_{jht} + \beta_{1ht} Distance_{kjt} + \beta_{2ht}\left(Distance_{kjt} \times Nonmetro_k\right)\right]}{\sum_{\ell \in C_{kt}} \exp\left[\alpha_{\ell ht} + \beta_{1ht} Distance_{k\ell t} + \beta_{2ht}\left(Distance_{k\ell t} \times Nonmetro_k\right)\right]}$$

Because the Hospital Service Area data are aggregated to ZCTA-by-hospital counts, we estimate the model using a conditional logit model adapted to grouped choice counts. In practice, we implement the model in R using the 'mclogit' package, treating each residence ZCTA as a choice case and observed discharge counts as grouped choices (*The IWLS Algorithm Used to Fit Conditional Logit Models*, n.d.).

Hospital fixed effects are included in the estimation step to absorb persistent differences in hospital attractiveness, such as reputation, service mix, referral relationships, and other hospital-level time-invariant unobserved factors. The coefficients on distance and the distance-by-nonmetro interaction are then used to construct geography-based predicted choice probabilities.

In HRR-year models where the distance-by-nonmetro interaction is not identified because of insufficient within-market variation, the interaction coefficient is set to zero when constructing predicted probabilities.

The purpose of the hospital choice model is to identify the relationship between distance and hospital choice while netting out persistent hospital attractiveness. After estimating the model, we then construct predicted probabilities using only the distance-related component of utility. Specifically, we define:

$$\widetilde{U}_{kjt} = \hat{\beta}_{1ht} Distance_{kjt} + \hat{\beta}_{2ht}\left(Distance_{kjt} \times Nonmetro_{k}\right)$$

We then compute geography-predicted hospital choice probabilities as

$$\tilde{P}_{kjt} = \frac{\exp\left(\widetilde{U}_{kjt}\right)}{\sum_{\ell \in C_{kt}} \exp\left(\widetilde{U}_{k\ell t}\right)}$$

Thus, $\tilde{P}_{kjt}$reflects predicted hospital choice based only on distance and the interaction between distance and nonmetropolitan residence. Hospital fixed effects are omitted from this prediction step. This ensures that the predicted probabilities reflect geographic access rather than realized hospital attractiveness or endogenous patient sorting. We verify that the predicted probabilities sum to one within each ZCTA's hospital choice set.

Geography-predicted hospital-level HHI

Analogous to observed hospital-level HHI, for the geography-predicted measure, we define predicted discharges from ZCTA k to hospital j as $\widetilde{D}_{kjt} = D_{kt}\tilde{P}_{kjt}$. Similarly, to adjust for hospital system affiliation, we let $\widetilde{D}_{kgt}$ denote predicted discharges from residence ZCTA k to hospitals in system g in year t, restricted to hospitals in the ZCTA k's choice set. The geography-predicted ZCTA-level HHI is $HHI_{kt}^{geo} = \sum_{g}\left(\frac{\widetilde{D}_{kgt}}{D_{kt}}\right)^{2}$, and hospital j's geography-predicted HHI is then:

$$HHI_{jt}^{geo} = \sum_{k}\left(\frac{\widetilde{D}_{kjt}}{\widetilde{D}_{jt}}\right) HHI_{kt}^{geo}$$

where $\widetilde{D}_{jt} = \sum_{k}\widetilde{D}_{kjt}$. Thus, both observed and geography-predicted hospital-specific HHI measures summarize the average concentration of the residence markets from which a hospital draws patients, weighted by observed discharge weights and geography-predicted discharge weights, respectively.

IV. Results

Characteristics of hospitals by CJR Mandatory and Voluntary MSAs

Table 1 describes hospitals in mandatory and voluntary MSAs by baseline market concentration. Mandatory MSAs included more hospitals and larger hospitals, but they had fewer CJR episodes (mean [SD]: 154 [253]) than voluntary MSAs (164 [172]). High-HHI hospitals are similar in hospital size, but more likely to be nonprofit, and less likely to be for-profit, as compared to low-HHI hospitals; this pattern is consistent with both mandatory and voluntary MSAs.

Hospitals in high-HHI markets had more eligible CJR episodes and higher reconciliation amounts in mandatory MSAs. Reconciliation payment per episode was also higher among high-HHI hospitals (mean [SD]: $705 [1,077] vs. $502 [1,207] in mandatory MSAs; $696 [908] vs. $656 [1,025] in voluntary MSAs). Quality performance was generally similar across HHI levels and MSA types.

Does market structure predict staying in CJR when participation becomes voluntary?

Among hospitals in markets in which participation in the Medicare CJR program was mandatory, those in less competitive markets were more likely than those in more competitive markets to remain in the CJR program when participation become voluntary.

We find that, among hospitals located in MSAs that switched from mandatory to voluntary participation, the decline in participation was larger among those in high HHI (less competitive) markets, although the difference was relatively small (**Figure 1**).

In regression models, we did not find evidence that hospitals with above-median baseline observed HHI have differentially remained in CJR than those with lower HHI (**Table 2**). In contrast, baseline CJR performance, measured by receipt of reconciliation bonuses during the mandatory period, was a strong predictor of staying in the program (15 pp.; $p<0.01$), suggesting that realized performance, rather than market structure, primarily drives voluntary participation.

Does market structure predict performance under CJR in always mandatory markets?

Among hospitals in markets in which participation in the Medicare CJR program is always mandatory, those in less competitive markets (with a higher lagged HHI) is positively associated with receiving a bonus under CJR ($p<0.01$) but do not significantly differ in CJR quality measures, consistent with the interpretation that hospitals in more concentrated markets have a structural advantage in achieving greater episode savings.

Table 3 reports OLS and 2SLS estimates of the relationship between lagged hospital HHI and CJR financial and quality performance in mandatory MSAs. Both OLS and 2SLS estimates are positive and statistically significant, with the 2SLS estimate being larger in magnitude: a one-standard-deviation increase in hospital lagged HHI is associated with a 5.7 percentage point ($p<0.05$) increase in the probability of receiving a bonus. The first-stage diagnostics indicate a strong relationship between predicted and observed HHI, with a Kleibergen-Paap Wald F statistic of 81. The difference between the OLS and IV estimates suggests that the naive association may understate the relationship between market concentration and CJR financial performance. We did not find evidence that higher market concentration leads to significant differences in CJR quality measures. If anything, both the 2SLS estimates on both quality measures are negative. A one-standard-deviation increase in hospital lagged HHI leads to a 0.071 standard deviation decrease in patient experience score, although the estimated effect is imprecise ($p<0.1$).

Does market structure predict target prices and episode complexity?

Among hospitals in mandatory MSAs, higher market concentration was associated with higher target prices. We did not indicate a statistically significant association between market concentration and the share of episodes with major complications or comorbidities.

Table 4 examines target prices and episode complexity as potential mechanisms of why higher market concentration may lead to better CJR financial performance. Higher lagged hospital HHI was associated with higher target prices in both OLS and 2SLS models. In the 2SLS model, a one-standard-deviation increase in lagged hospital HHI was associated with a 0.120 standard deviation increase in target price ($p<0.05$).

For episode complexity, higher lagged HHI was associated with a greater share of episodes with MCC in the OLS model (1.1 percentage points; $p<0.01$). However, the 2SLS estimate was not statistically significant. These findings are consistent with higher payment benchmark potentially contributing to the financial advantage of hospitals in more concentrated markets. In robustness analysis, we used a different definition of market boundary from 35 to 50 miles to calculate observed and predicted hospital HHIs. All of our findings remained robust (**Tables 5a and 5b**).

V. Conclusion

In this paper, we examine how hospital market concentration affects participation and performance under APMs, leveraging the unique design of the CJR program. Hospitals in more concentrated markets are more likely to receive reconciliation bonuses under mandatory participation; we did not find clear effects on quality performance. Our preliminary analysis on mechanisms shows obtaining higher target prices may contribute to this financial advantage. We find little evidence that market concentration predicts remaining in CJR when participation becomes voluntary.

These findings suggest that market power may shape hospitals' financial returns under alternative forms of price. As CMS shifts toward testing more mandatory models, mandatory participation may reduce selection into these programs, but financial advantages associated with market concentration appear to be intact. Our results have implications for policymakers designing VBP incentives and highlight the importance of considering how market structure shapes hospitals' responses to these incentives. Our results also inform regulatory concerns about provider consolidation by pointing to a potential new channel through which rising concentration may increase health care spending. Market power may affect spending not only through the level of negotiated prices but also through alternative forms of price, including episode-based reconciliation payments under administratively set Medicare prices.

Our ongoing work will consider vertical integration, examine additional channels through which hospitals in higher concentrated market achieve better performance under CJR, and examine whether similar patterns exist in other Medicare VBP models.

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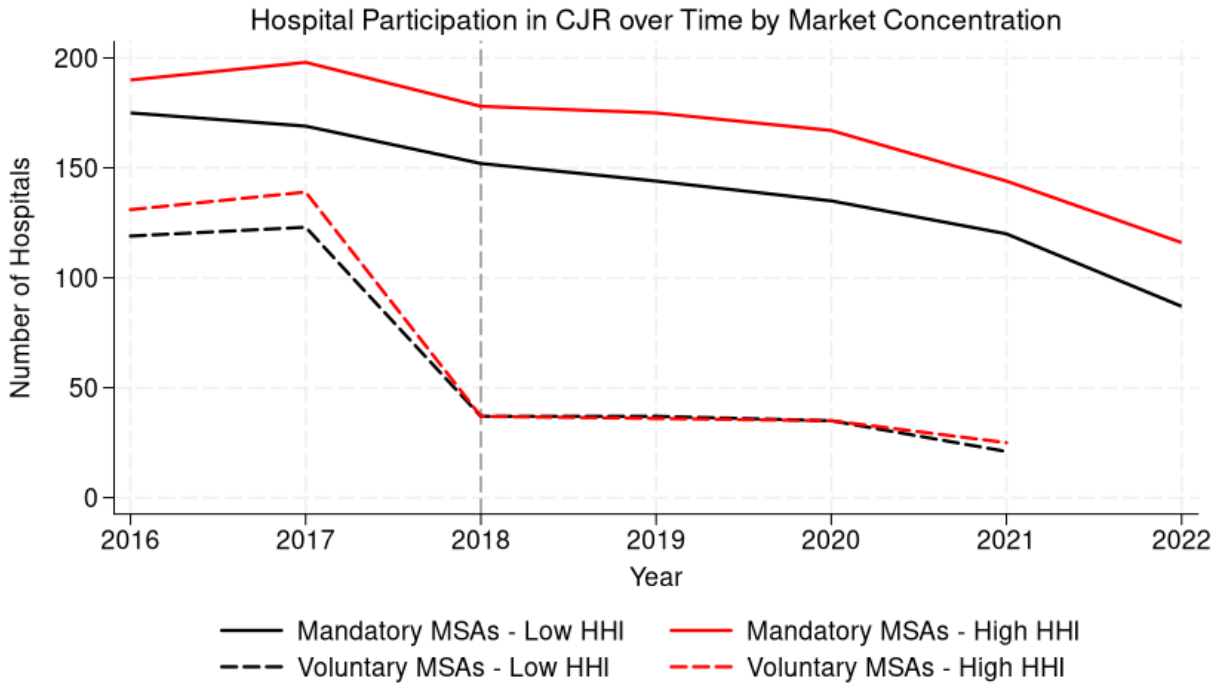
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Figure 1: Hospital Participation in CJR over time by Market Concentration



Note: High HHI is defined as average baseline HHI (2016-2017) above the MSA-specific median; low HHI is defined as average 2016-2017 HHI at or below the MSA-specific median.

Table 1. Baseline Characteristics of Hospitals by CJR Mandatory and Voluntary MSAs in Analytic Sample

Hospital Market Concentration	Mandatory MSAs (n=34)			Voluntary MSAs (n=32)		
	All	Low HHI	High HHI	All	Low HHI	High HHI
Average HHI (2016-2017)	0.453 (0.137)	0.376 (0.101)	0.519 (0.129)	0.461 (0.116)	0.398 (0.083)	0.518 (0.111)
Number of hospitals, 2017	367	169	198	262	123	139
Hospital size, number of beds, mean (SD) 2017	303 (257)	302 (261)	303 (254)	233 (190)	236 (194)	230 (187)
Ownership, n (%) 2017						
Nonprofit	232 (63)	97 (57)	135 (68)	167 (64)	72 (59)	95 (68)
For-profit	100 (27)	59 (35)	41 (21)	56 (21)	33 (27)	23 (17)
Government	35 (10)	13 (8)	22 (11)	39 (15)	18 (15)	21 (15)
Participation in CJR (2018)	330	152	178	74	37	37
Number of eligible CJR episodes, mean (SD) 2017	154 (253)	131 (296)	174 (209)	164 (172)	164 (175)	163 (170)
Target price, mean (SD) 2017	39,734 (5,056)	39,693 (5,337)	39,769 (4,817)	35,371 (5,095)	35,528 (4,765)	35,233 (5,381)
Share of CJR episodes with MCC, mean (SD) 2017	0.029 (0.054)	0.024 (0.055)	0.033 (0.054)	0.036 (0.055)	0.029 (0.051)	0.042 (0.058)
Reconciliation amount, mean (SD) 2017	144,672 (369,274)	123,077 (458,360)	163,105 (270,992)	143,199 (210,196)	145,918 (207,066)	140,793 (213,646)
Reconciliation per episode, mean (SD) 2017	611 (1,142)	502 (1,207)	705 (1,077)	677 (963)	656 (1,025)	696 (908)
Quality performance, mean (SD) 2017						
Complication rate	2.54 (0.50)	2.57 (0.51)	2.52 (0.50)	2.59 (0.57)	2.57 (0.53)	2.61 (0.61)
Patient experience score	84.97 (2.96)	84.34 (3.22)	85.50 (2.60)	86.47 (2.25)	86.34 (2.38)	86.59 (2.14)

Note: High HHI is defined as average 2016-2017 observed HHI greater than or equal to the MSA-specific median; low HHI is defined as average 2015–2017 HHI below the MSA-specific median. Abbreviations: MCC, major complication or comorbidity.

Table 2. Association between Hospital Market Concentration and Staying in CJR when Participation Became Voluntary in 2018, Among Hospitals in Voluntary MSAs.

Baseline characteristics[1]	**(1)**	**(2)**
Reconciliation payment	0.15**	0.15**
	(0.06)	(0.06)
Hospital HHI[2]		0.03
		(0.05)
Number of CJR episodes[2]	0.13***	0.13***
	(0.03)	(0.03)
Average HCC scores[2,3]	-0.07*	-0.07
	(0.04)	(0.04)
Average patient age[2,3]	-0.01	-0.01
	(0.04)	(0.04)
Share of duals[2,3]	-0.03	-0.03
	(0.04)	(0.04)
CJR rural	0.05	0.05
	(0.10)	(0.10)
For-profit	0.07	0.07
	(0.08)	(0.08)
Constant	0.17***	0.18**
	(0.06)	(0.06)
MSA FE	X	X
Observations	250	250
R-squared	0.38	0.38

Note: Each column is a separate regression. Dependent variable is a binary indicator of whether hospital stayed in CJR in 2018. [1]Baseline characteristics are averaged across 2016-2017 prior to voluntary switch in 2018. [2]Continuous variables are standardized. [3]Among Medicare fee-for-service beneficiaries. Robust standard errors in parentheses and clustered at the MSA level. *** p<0.01, ** p<0.05, * p<0.1. Abbreviations: CJR, Comprehensive Care for Joint Replacement Model; MSA, Metropolitan Statistical Areas; HHI, Herfindahl-Hirschman Index; HCC, Hierarchical Condition Categories; FE, fixed effects.

Table 3. Effects of Hospital Market Concentration on CJR Financial and Quality Performance 2017-2022, Among Hospitals in Mandatory MSAs.

	CJR bonus		Complication Rate		Patient Experience	
Variables	OLS	2SLS	OLS	2SLS	OLS	2SLS
Lagged Hospital HHI	0.033**	0.057**	-0.065	-0.139	0.002	-0.071*
	(0.012)	(0.023)	(0.071)	(0.102)	(0.041)	(0.040)
Control	X	X	X	X	X	X
MSA FE	X	X	X	X	X	X
Year FE	X	X	X	X	X	X
Observations	1,693	1,693	1,611	1,611	1,688	1,688
R-squared	0.284	0.129	0.283	0.069	0.546	0.331
First-stage diagnostic statistics						
KP Wald F		81.28		72.34		85.34
Cragg-Donald F		831.9		816.1		844.6
KP LM statistic		3.394		3.347		3.411
KP LM p-value		0.0654		0.0673		0.0648

Note: Each column reports a separate regression. The dependent variables include (1) receipt of CJR bonus as an indicator equal to one if the hospital received a positive reconciliation payment; (2) standardized CJR complication rate; (3) standardized patient experience score. Lagged hospital HHI is standardized. In the IV specification, lagged observed hospital HHI is instrumented with lagged geography-predicted hospital HHI (35 mile). First-stage diagnostic statistics are reported for the IV model. All regressions include the same set of covariates reported in Table 2. Robust standard errors in parentheses and clustered at the MSA level. *** p<0.01, ** p<0.05, * p<0.1.

Table 4. Effects of Hospital Market Concentration on Target Price and Share of Episodes with Major Complication or Comorbidity, Among Hospitals in Mandatory MSAs.

	Target Price		MCC Share	
Variables	OLS	2SLS	OLS	2SLS
Lagged Hospital HHI	0.107***	0.120**	0.011***	-0.004
	(0.030)	(0.047)	(0.003)	(0.005)
Control	X	X	X	X
MSA FE	X	X	X	X
Year FE	X	X	X	X
Observations	907	907	1,693	1,693
R-squared	0.445	0.229	0.177	0.078
First-stage diagnostic statistics				
KP Wald F		64.61		81.28
Cragg-Donald F		405.8		831.9
KP LM statistic		3.590		3.394
KP LM p-value		0.0581		0.0654

Note: Each column reports a separate regression. The dependent variables include (1) standardized mean target price across CJR diagnosis-related groups; (2) share of CJR episodes with MCC. Lagged hospital HHI is standardized. In the IV specification, lagged observed hospital HHI is instrumented with lagged geography-predicted hospital HHI (35 mile). Target price analysis were restricted to years when benchmarks blended hospital-specific and regional historical spending; later years with fully regional benchmarks were excluded. First-stage diagnostic statistics are reported for the IV model. All regressions include the same set of covariates reported in Table 2. Robust standard errors in parentheses and clustered at the MSA level. *** p<0.01, ** p<0.05, * p<0.1.

Table 5 (a). Different Market Boundary: Effects of Hospital Market Concentration on CJR Financial and Quality Performance 2017-2022, Among Hospitals in Mandatory MSAs.

	CJR bonus		Complication Rate		Patient Experience	
Variables	OLS	2SLS	OLS	2SLS	OLS	2SLS
Lagged Hospital HHI	0.036***	0.073***	-0.072	-0.158	0.008	-0.041
	(0.010)	(0.024)	(0.069)	(0.115)	(0.042)	(0.044)
Control	X	X	X	X	X	X
MSA FE	X	X	X	X	X	X
Year FE	X	X	X	X	X	X
Observations	1,693	1,693	1,611	1,611	1,688	1,688
R-squared	0.284	0.127	0.283	0.068	0.546	0.333
First-stage diagnostic statistics						
KP Wald F		89.17		79.89		93.30
Cragg-Donald F		621.5		615.9		634.3
KP LM statistic		3.430		3.405		3.454
KP LM p-value		0.0640		0.0650		0.0631

Note: Each column reports a separate regression. The dependent variables include (1) receipt of CJR bonus as an indicator equal to one if the hospital received a positive reconciliation payment; (2) standardized CJR complication rate; (3) standardized patient experience score. Lagged hospital HHI is standardized. In the IV specification, lagged observed hospital HHI is instrumented with lagged geography-predicted hospital HHI (50 mile). First-stage diagnostic statistics are reported for the IV model. All regressions include the same set of covariates reported in Table 2. Robust standard errors in parentheses and clustered at the MSA level. *** p<0.01, ** p<0.05, * p<0.1.

Table 5 (b). Different Market Boundary: Effects of Hospital Market Concentration on Target Price and Share of Episodes with Major Complication or Comorbidity, Among Hospitals in Mandatory MSAs.

	Target Price		MCC Share	
Variables	OLS	2SLS	OLS	2SLS
Lagged Hospital HHI	0.122***	0.110***	0.010**	-0.005
	(0.028)	(0.039)	(0.004)	(0.005)
Control	X	X	X	X
MSA FE	X	X	X	X
Year FE	X	X	X	X
Observations	907	907	1,693	1,693
R-squared	0.447	0.231	0.174	0.077
First-stage diagnostic statistics				
KP Wald F		71.79		89.17
Cragg-Donald F		299.4		621.5
KP LM statistic		3.598		3.430
KP LM p-value		0.0578		0.0640

Note: Each column reports a separate regression. The dependent variables include (1) standardized mean target price across CJR diagnosis-related groups; (2) share of CJR episodes with MCC. Lagged hospital HHI is standardized. In the IV specification, lagged observed hospital HHI is instrumented with lagged geography-predicted hospital HHI (50 mile). Target price analysis were restricted to years when benchmarks blended hospital-specific and regional historical spending. First-stage diagnostic statistics are reported for the IV model. All regressions include the same set of covariates reported in Table 2. Robust standard errors in parentheses and clustered at the MSA level. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.