# Evaluating Substitutes for Federal Antitrust: The Case of COPAs*

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**Abstract**

Are state-imposed behavioral remedies effective substitutes for federal antitrust enforcement? We evaluate state regulation of hospital mergers under Certificates of Public Advantage (COPAs). Using hospital data from 1996-2022, we compare COPA-regulated mergers to unregulated mergers with similar anticompetitive potential. In highly concentrated markets, COPA mergers result in 11.1 p.p. lower price growth but 0.5 p.p. greater increases in 30-day mortality rates. We find a negative correlation between price and mortality effects for COPA mergers, consistent with theoretical predictions that binding price caps exacerbate quality deterioration. Our findings suggest that COPA contracts are poor substitutes for traditional antitrust enforcement.

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# 1 Introduction

Conflicts between local and federal regulators are common in the regulation of healthcare markets.[1] This is also true for the oversight of hospital markets, which greatly influence the health and economic well-being of local communities while being subject to numerous federal regulations related to patient safety, privacy, and consolidation (Alexander and Richards, 2023; Hoffman, Ha, Fan, and Li, 2024). With respect to consolidation, misalignment between the priorities of the Federal Trade Commission (FTC) and local regulators, including state attorneys general and health departments, has resulted in intervention by the latter in the enforcement of antitrust law.

This paper analyzes the extent to which state oversight constrains the anticompetitive effects of hospital mergers. State oversight typically takes one of two forms: conditional approval via consent decrees or Certificates of Public Advantage (COPAs). State-issued consent decrees are negotiated settlements between the state attorney general and merging parties following a challenge to the merger based on antitrust law. COPAs, on the other hand, provide *antitrust immunity* to hospital mergers for which the "benefits outweigh any anticompetitive effects of joining together competitors" (Texas Department of State Health Services, 2024). COPAs are more homogeneous than consent decrees in their use of behavioral (rather than structural) remedies, in the types of behavioral remedies employed, and in the completeness of available documentation on the terms of the deal. For these reasons, our analysis of state oversight of mergers focuses on COPAs.[2]

The impact of COPAs has become a matter of policy debate. The majority of COPA mergers typically involve two hospital networks located in sparsely populated, highly concentrated healthcare markets. Merging parties typically argue that antitrust officials underestimate the potential for their transaction to produce cost efficiencies and minimize the risk of closure. COPA mergers are subject to regulations, adherence to which is typically moni-

[1] For example, states may design their Medicaid programs that do not align with the preferences of a given Presidential administration.

[2] We return to state-imposed consent decrees in Section 6.

tored by local agencies. COPA proponents contend that such regulatory oversight can help preserve access to care while mitigating the adverse effects of enhanced market power, such as higher prices, suppressed wages, and diminished quality of care. Antitrust entities such as the FTC emphasize that these local agencies are incapable of writing complete contracts to regulate the behavior of merging parties or of effectively enforcing their provisions (Khan, Phillips, Slaughter, Wilson, and Bedoya, 2022). Since the inception of COPA legislation in Wisconsin in 1991, ten hospital mergers that would otherwise have required FTC review have been consummated under COPA protection. As of the time of writing, 19 states have active COPA laws, with Indiana being the latest state to enact such legislation in 2021.

This paper quantifies the effects of behavioral remedies imposed under COPAs by estimating the effects of hospital mergers consummated under different regulatory regimes. We assemble a panel of hospital data from 1996 to 2022 from the Centers for Medicare and Medicaid Services' hospital cost reports and Hospital Compare Database to construct inpatient prices and 30-day mortality rates. We supplement these hospital data with system affiliation data from the American Hospital Association annual survey to construct measures of hospital ownership and changes in ownership resulting from mergers. We identify which of these mergers are subject to conditional approval under COPAs using state and federal records. Using information on COPA terms, we note that while there is variation across COPAs in the breadth of outcomes governed, they all regulate post-merger prices by imposing caps on price growth or operating margins.

To estimate the effect of mergers, we employ the imputation-based two-way fixed effects estimator of Borusyak, Jaravel, and Spiess (2024), which allows for unrestricted heterogeneity in treatment effects. Our primary identification assumption is that, conditional on time-varying hospital and market characteristics, outcomes for non-merging hospitals in our control group are informative about how outcomes would have evolved for merging hospitals had they not merged. Our two treatment groups include hospitals that merge under COPAs and those that merge in the absence of structural or behavioral remedies. However, the

presence of behavioral remedies may not be the only difference between the COPA and non-COPA treatment groups, and these differences may bias our treatment effects. For example, COPA mergers are more anticompetitive than non-COPA ones, as measured by the change in and resultant level of concentration they generate. To account for this difference, we use the Coarsened Exact Matching algorithm of Iacus, King, and Porro (2012) to construct weights that minimize imbalance between the two groups in their anticompetitiveness. We use these weights to compute the weighted average of treatment effects across hospitals in each of our treatment groups, enabling us to attribute differences in merger effects to the presence of behavioral remedies.

In our main results, we find that COPA mergers reduce price growth by an average of 7.3% and increase growth in 30-day mortality rates by 0.8 p.p. over a baseline average of 13.8%. For non-COPA mergers, we estimate a price increase of 6%, implying a 13.3 p.p. lower price effect for mergers subject to behavioral remedies. While the difference in overall mortality rates between the two groups is not significant, we find that COPA mergers result in increases in mortality from heart attack and heart failure that are 0.9 and 0.7 p.p. higher, respectively, than those induced by non-COPA mergers. When focusing on hospitals experiencing relatively large increases in concentration, we find that effects on price and mortality both increase, and that COPA mergers result in 11.1 p.p. lower price growth and 0.5 p.p. greater increases in 30-day mortality rates.

These findings suggest that COPAs successfully constrain the exercise of newfound market power along some dimensions (prices), but exacerbate its effect on others (quality provision as measured by mortality rates). We illustrate this mechanism with a simple theoretical model of price and quality choice by profit-maximizing hospitals from Gaynor (2007). When prices are unregulated, the effects of consolidation on quality are ambiguous and depend on the relative size of changes in the price and quality elasticities of demand. However, under binding post-merger price caps, the predictions are much clearer: lower price caps reduce post-merger margins and weaken incentives to provide quality. These effects are increasing

in the anticompetitiveness of the transaction as post-merger price caps are more likely to bind. Our finding that price growth is constrained and mortality effects are heightened for COPA mergers relative to non-COPA ones is consistent with these predictions.

The theory laid out here implies not just differences in average treatment effects based on the presence of behavioral remedies, but also differences in the *correlation* between individual treatment effects on price and quality across the two treatment groups. Among hospitals undergoing COPA mergers, we estimate that a 1 p.p. reduction in price effects is associated with a 1 p.p. increase in mortality rates, and find no significant correlation between price and mortality effects for hospitals participating in non-COPA mergers. These findings suggest that post-merger price regulation may inadvertently intensify the exercise of market power along harder-to-regulate dimensions of hospital behavior, highlighting the inherent challenge of designing comprehensive regulatory contracts.

This paper's primary contribution is to evaluate the efficacy of local oversight of hospital mergers in constraining the effects of market power and to assess whether local oversight can effectively substitute for federal antitrust regulation. This evaluation is timely given recent evidence of underenforcement of antitrust law in the hospital sector (Brot, Cooper, Craig, and Klarnet, 2024). While the number of COPA mergers is small, the number of hospital mergers remains high, 72 being announced in 2024 alone (Singh, 2025). These facts suggest that the impacts of COPA mergers, while few in number, speak to a larger policy issue of how to regulate ongoing consolidation in hospital markets given finite regulatory resources.

The existing literature on COPAs largely consists of case reports documenting institutional and legislative details of individual COPA mergers (Bovbjerg and Berenson, 2015; Brown, 2018; Fuse Brown, 2019; Amin, 2024; Khan et al., 2022). To our knowledge, Garmon and Kmitch (2018) and Garmon and Bhatt (2022) are the first and only empirical papers to rigorously evaluate the effect of COPAs. Looking at four different COPA mergers, Garmon and Bhatt (2022) construct a synthetic control for each case using hospital systems from the same state to evaluate long-run effects on price and quality. By focusing on COPAs

consummated before 2015, the authors are able to estimate merger effects for years during which the COPA contracts are active as well as the years following their repeal, during which no provisions are enforced. Although results are fairly heterogeneous across mergers, the authors do not find much evidence of price increases while contracts are active, but do find a significant increase in prices for some mergers after contracts are repealed. In some cases, these price increases are as high as 45% relative to controls.

Our work differs in several key ways. We quantify the effects of behavioral remedies by comparing COPA mergers to similar ones free from such regulations, focusing on their short-term effectiveness while abstracting from the possibility of repeal due to regulatory fatigue or lobbying. We also provide and test a theory of spillovers from regulated dimensions of hospital behavior to unregulated ones that explains divergence in the price and quality effects of COPA mergers relative to non-COPA ones. Additionally, we extend the evaluation to include additional outcomes of interest, including wages and employment, and cover seven of the ten total COPA cases, including the Texas mergers consummated in 2020.

We also contribute to a literature studying the effect of hospital mergers, adding to research that leverages the high frequency of mergers in the healthcare sector to quantify the consequences of market power on prices (Dafny, 2009; Cooper, Craig, Gaynor, and Van Reenen, 2019; Brand, Garmon, and Rosenbaum, 2023), quality (Ho and Hamilton, 2000; Wollmann, 2024), and wages (Prager and Schmitt, 2021). Wollmann (2024) is one of the few papers that studies a healthcare market in which prices are regulated. He finds that mergers between dialysis clinics increased mortality rates, arguing that strict price regulation forced firms to exercise market power through quality reductions. More broadly, our work contributes to recent research examining merger effects beyond price, including studies of firm coordination (Miller and Weinberg, 2017), monopsony power (Arnold, 2019; Prager and Schmitt, 2021), and management strategy exchange (Eliason, Heebsh, McDevitt, and Roberts, 2020). To our knowledge, this is the first empirical assessment of how state-imposed behavioral remedies affect both the levels of regulated and unregulated outcomes and the

relationship between them.

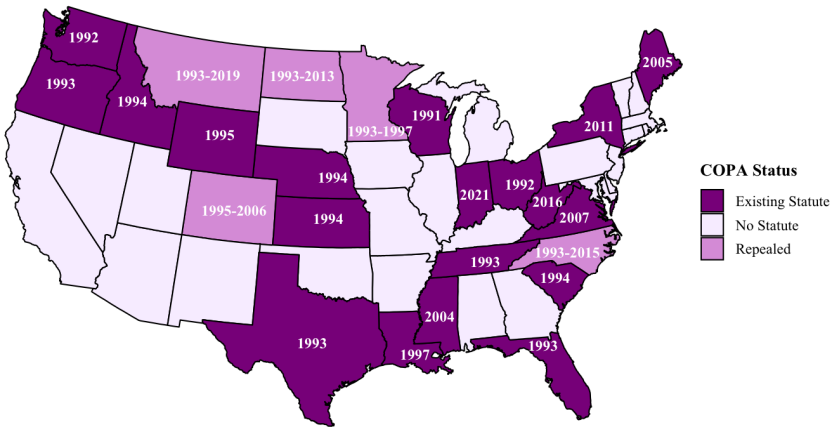
Beyond healthcare markets, this work connects to broader questions about the economics of federalism and tensions between state and federal regulatory authority. The state action doctrine, established in 1943 with *Parker v. Brown*, allows states to grant the conduct of private parties immunity from federal antitrust laws, raising critical questions about public interest and market efficiency (Parker, 1943; Lopatka, 1984). While state regulations may use this power to address market failures, they are also more susceptible to regulatory capture by local interests than federal oversight (Easterbrook, 1983). Our evaluation of state-imposed behavioral remedies as substitutes for federal antitrust scrutiny is particularly timely given proposed budget cuts to the FTC and freezes to Department of Justice Antitrust Division spending (Office of Management and Budget, 2025). Our findings suggest that COPA contracts are incomplete and serve as poor substitutes for traditional antitrust enforcement.

## 2 Certificates of Public Advantage

Federal antitrust law provides a limited exemption for state action under the Parker doctrine, which requires that anticompetitive state conduct be clearly articulated as state policy and actively supervised (Parker, 1943). COPA laws are state legislative regimes that create a process by which hospitals apply for federal antitrust immunity and are evaluated by states on the basis of whether their merger would generate positive externalities. Wisconsin passed the first COPA law in 1991, under which the state would issue the parties a COPA if "the benefits likely to result from the agreement substantially outweigh any disadvantages attributable to a reduction in competition likely to result" (State of Wisconsin, 2024). Since then, 24 other states have passed COPA laws, and, as of 2024, 19 states have active COPA statutes. Indiana enacted its legislation in 2021, and its Department of Health is currently considering the COPA application of Union Health and Terre Haute Regional Hospital in

Vigo County (Cooper, 2025). Six states that at one point had COPA laws have repealed them as of 2024. An illustration of the timing of each COPA law can be found in Figure 1.

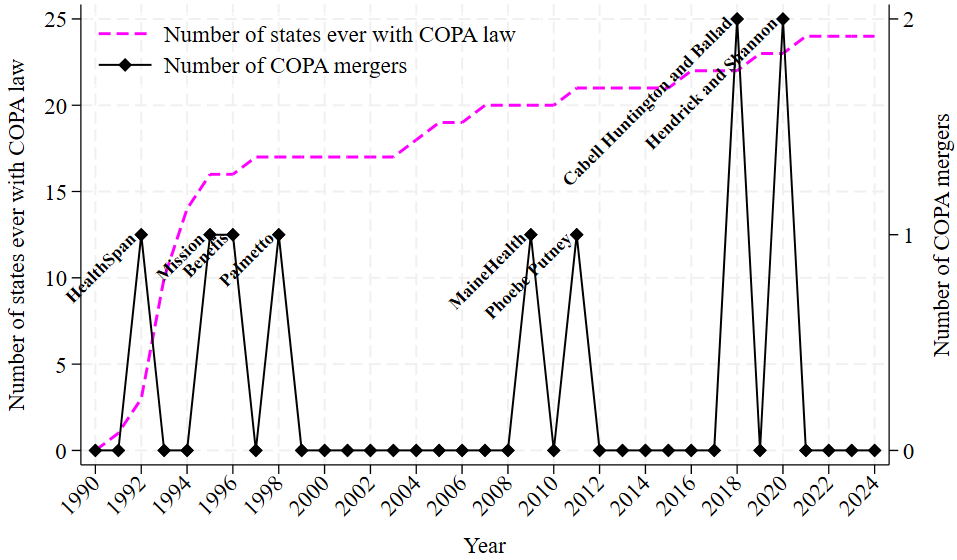
Figure 1: Map of COPA Law Adoption




*Notes:* Years superimposed on states correspond to the year in which the state's COPA law was passed, after which it was active until present. Ranges correspond to the years in which a state's COPA law was active, prior to its repeal.

There have been ten mergers consummated under COPAs, the first being the merger of the HealthOne and LifeSpan systems in Minnesota in 1992. An illustration of the timing of each COPA merger can be found in Figure 2. Hospital mergers pursued under COPA law are frequently justified by the merging parties on the grounds that one of the participants would be forced to close absent the merger, reducing healthcare access in what are often rural or underserved regions. For instance, in the 1995 merger of Mission Hospital and St. Joseph's Hospital in Asheville, North Carolina, the hospitals successfully lobbied for COPA protection by emphasizing St. Joseph's financial vulnerability and the risk of service loss (Hall, 2024). Similar justifications were presented in the Benefis Health merger in Montana and in the Cabell Huntington merger in West Virginia, where each hospital was the sole acute care provider in its county (FTC, 1996; West Virginia Health Care Authority, 2016). These narratives frame consolidation as necessary to maintain community health infrastructure.

Figure 2: Timing of COPA Mergers and Adoption




*Notes:* Solid black line shows the number of COPA mergers in each year. Labels next to diamond markers display the name of mergers occurring in that year. Dashed magenta line shows the number of states to have ever adopted a COPA law as of that year.

The process of evaluating whether to grant a merger federal antitrust immunity typically involves a combination of legislative review and administrative oversight by state health authorities or attorneys general. States assess whether the merger's purported benefits—such as cost savings, improved quality, and sustained access—outweigh the risks associated with reduced competition. This evaluation sometimes involves economic modeling; more often, it is grounded in public comments and consultations with stakeholders. When a COPA is approved, states typically impose a suite of regulatory conditions designed to constrain the newly created market power. While terms vary across COPA contracts, they all include restrictions on price growth or operating margins. Online Appendix A summarizes the terms of each COPA. There are no COPA contracts that impose any kind of structural remedies on the merging parties.

To enforce these conditions, states employ a range of monitoring tools, including mandatory reporting on financial performance, quality metrics, and access indicators, as well as periodic audits and third-party evaluations. Typically, oversight bodies are established within

existing state agencies, as in Texas, where the state health department supervises COPA hospitals (Texas Department of State Health Services, 2024). However, oversight effectiveness can be undermined by political turnover or fiscal constraints. Some COPAs include sunset clauses or periodic renegotiation, while others are structured with open-ended regulatory commitments. The repeal of COPAs raises concerns about regulatory capture. Garmon and Bhatt (2022) document significant increases in prices in the post-repeal period relative to the post-merger, pre-repeal period. Our focus in this paper will be on the efficacy of COPA contracts while they are enforced, prior to any revision or repeal.

# 3 Merger Effects Under Incomplete Contracts

In writing COPA contracts, local regulators face the difficult task of designing regulations for the various dimensions of firm behavior along which the merging hospitals may exercise their market power. Naturally, some dimensions are easier to quantify and monitor than others. For example, virtually all COPAs have some form of price regulation, but there is significant heterogeneity in the presence and scope of quality regulations, as health care quality is multidimensional and condition-dependent. These measurement issues make it likely that COPA contracts are incomplete and do not fully regulate all potentially anticompetitive behaviors. This incomplete regulation may not only fail to prevent market power exercise in unregulated areas, but may actually redirect anticompetitive behavior toward dimensions that are harder to monitor and control.

In what follows, we present a model of a hospital's joint decision over price and quality and contrast the forces at play under different regulatory regimes: market-determined prices and regulated prices. The exposition follows Gaynor (2007) and its application of the Dorfman and Steiner (1954) model of advertising in a monopolistic competition setting. Consider a hospital facing a downward-sloped residual demand curve $q(p, z)$ that is decreasing in price $p$ and increasing in a vertical measure of quality $z$. Assuming a linear cost function

$C(q(p,z),z) \equiv q(p,z)(c+dz)$ with $c$ and $d$ being positive parameters, the hospital chooses price and quality to maximize its profit function:

$$\pi(p,z) = q(p,z)(p-c-dz).$$

A merger affects the hospital's choices through its impact on both marginal cost and residual demand, the latter interpreted as (potentially unequal) declines in the absolute value of the price and quality elasticities of demand, $\varepsilon_p$ and $\varepsilon_z$.

**Market-determined prices:** In the absence of price regulation, hospitals choose price and quality jointly to maximize profits. Solving for the optimal price and quality as a function of demand elasticities and marginal cost terms yields:

$$p^* = c\left(\frac{\varepsilon_p}{\varepsilon_p - \varepsilon_z - 1}\right)$$
$$z^* = \frac{c}{d}\left(\frac{\varepsilon_z}{\varepsilon_p - \varepsilon_z - 1}\right)$$

Taking the total derivative with respect to the price and quality elasticity, it is easy to see that the optimal price and quality increase with a decrease in the price elasticity and decrease with a decrease in the quality elasticity (so long as $\varepsilon_p > 1$). Hence, the impact of consolidation on both price and quality is ambiguous and depends on the relative impacts on the quality and price elasticities of demand. In particular, if the ratio of the quality and price elasticities of demand falls by more than price increases, then quality will fall as a result of consolidation. We note that if large cost synergies are realized, they will put downward pressure on this correlation and result in lower prices and higher quality. This theoretical ambiguity is consistent with the mixed findings of empirical work on the price and quality effects of mergers (Ho and Hamilton, 2000; Mukamel, Zwanziger, and Bamezai, 2002; Volpp, Williams, Waldfogel, Silber, Schwartz, and Pauly, 2003; Gowrisankaran and Town, 2003;

Capps, 2005; Propper, Burgess, and Gossage, 2008).

**Regulated prices:** Under price regulation, hospitals maximize profits subject to a price constraint $p \leq \bar{p}$. Under a binding post-merger price constraint, the hospital's optimal quality choice is

$$z^* = \left(\frac{\varepsilon_z}{1+\varepsilon_z}\right)\left(\frac{\bar{p}-c}{d}\right).$$

This condition highlights that the impact of mergers on quality is more predictable when prices are regulated. Specifically, mergers that reduce sensitivity to quality put downward pressure on the hospital's quality choice. Under regulated prices, the usual countervailing force—where higher price margins due to lower price sensitivity put upward pressure on the hospitals' quality choice—is effectively neutralized. Additionally, under binding price constraints, price and quality effects should be positively correlated: lower price constraints imply lower returns to quality investment, and thus lower levels of quality in equilibrium. In the case where quality is measured by, e.g., mortality rates, we predict that lower price constraints imply higher mortality rates and a negative correlation between price and mortality effects.

# 4 Data

Building on the preceding discussion, our empirical analysis pursues two main objectives: First, we aim to compare the effects of hospital mergers subject to state-imposed behavioral remedies to similar mergers free from behavioral remedies. Second, we aim to assess the relationship between price and quality effects across the two regulatory regimes.

To evaluate the effect of hospital mergers on a range of price and non-price outcomes, we require information on hospital outcomes and characteristics over time. We make use of three main data sources: the American Hospital Association (AHA) annual survey, the Centers for Medicare and Medicaid Services (CMS) hospital cost reports, and the CMS

Hospital Compare data. For data on hospital ownership, we use the AHA annual survey data, which identifies the system to which each hospital belongs. We validate the merger activity implied by changes in the AHA system affiliation variable using hospital merger reports compiled by Irving Levin Associates. We collect publicly available data, including the FTC's list of merger filings and internet searches, to identify hospitals that participated in COPA mergers.

We use hospital cost reports from CMS from 1996-2022 to construct commercial inpatient prices as the net revenue for non-Medicare inpatient admissions divided by the number of non-Medicare admissions (Dafny, 2009; Dafny, Ho, and Lee, 2019). We follow Schmitt (2017) in calculating costs per adjusted discharge as the sum of total salaries and other costs across all cost centers divided by adjusted discharges (including estimated outpatient discharges based on charge shares). We construct hourly wages by aggregating workers into three categories—blue-collar, white-collar, and nursing and pharmacy—and calculating average hourly wages as total salaries divided by total work hours (Prager and Schmitt, 2021). Employment is measured as full-time equivalent employees per category.

Hospital characteristics include beds, outpatient charge fraction, Medicaid discharge share, and case-mix index, the latter of which comes from Medicare's annual impact files. Quality measures from CMS Hospital Compare data beginning in 2008 include Medicare mortality and readmission rates for heart failure, heart attack, and pneumonia, aggregated using admission-weighted averages.[3] Note that our inpatient price measures pertain to commercial patients, while our mortality and readmission measures pertain to Medicare patients. We assume that quality effects estimated for Medicare patients are informative about changes in quality to non-Medicare patients.[4]

We supplement our hospital-level data with county-level socio-economic measures. Data

[3]Online Appendix B describes how missing data for mortality and readmission rates impacts our estimation.

[4]Prior work provides evidence consistent with this assumption. For example, Barnett, Olenski, and Sacarny (2023) shows that a Medicare informational intervention changed provider behavior for both Medicare beneficiaries and patients covered by other insurers.

on population and the proportion of residents over 65 come from the U.S. Census Bureau. Measures of one-bedroom rent values come from the Department of Housing and Urban Development's Fair Market Rents data. We obtain per capita income from the Bureau of Economic Analysis' Personal Income data.

## 4.1 Sample Construction and Summary Statistics

For our main specification, we construct two treatment groups. The first treatment group consists of hospital mergers consummated under COPAs. We exclude the three COPA mergers before or in 1996, as they fall outside the range of years for which treatment effects can be estimated. This leaves us with seven COPA mergers.

The second treatment group comprises mergers that were consummated in the absence of any behavioral remedy, in states that have enacted COPA laws at some point in time, but either prior to the passage or after the repeal of the COPA law. We impose these restrictions because states that pass COPA laws may share common hospital market or regulatory characteristics that correlate with merger outcomes. We refer to this second treatment group as "non-COPA" mergers.

For all treated hospitals, we retain the six years preceding and four years following the year in which the merger was consummated. This restriction ensures that COPA regulations are active during the entirety of the post-merger period for hospitals in our COPA treatment sample. To avoid contamination between merger effects, we exclude from the non-COPA treatment group any mergers occurring within 10 years of a previous merger, since our estimation window requires six years of pre-treatment and four years of post-treatment data. There are a total of 44 non-COPA mergers that satisfy the criteria for inclusion in this second treatment group.

We follow Dafny et al. (2019) and construct a control group of hospitals that do not engage in merger activity and are part of a hospital system (i.e., are not independent) for the entirety of our sample period.

**Matching treatment groups:** COPA hospitals seek antitrust immunity from the state because they would presumably face federal antitrust action without it. This suggests that COPA mergers are likely to take place in more concentrated markets and involve greater increases in concentration than mergers in our comparison group. That there exist unchallenged mergers as anticompetitive as COPA ones follows from the results of Brot et al. (2024), which shows that that 20% of the over 1,000 mergers consummated between 2002 and 2020 could have been predicted to meaningfully lessen competition.

Measuring concentration requires us to define geographic markets. We follow Brot et al. (2024) and define markets based on distance, measuring a hospital's market share as its share of inpatient hospital beds among all hospitals within 25 miles of it. Using these definitions, we compare the change in and resultant HHI for COPA and non-COPA mergers in Table A1. As expected, COPA mergers result in an average post-merger HHI that is 1,239 points higher and an average change in HHI that is 341 points larger than non-COPA mergers.

We match COPA and non-COPA mergers using the Coarsened Exact Matching (CEM) algorithm. The algorithm computes weights that generate balance across the groups based on the change in HHI ($\Delta$HHI) and HHI level induced by the merger (Iacus et al., 2012). The CEM algorithm coarsens the values of HHI and $\Delta$HHI into strata. COPA (non-COPA) hospitals receive zero weight if there are no non-COPA (COPA) hospitals in their stratum. Weighting non-COPA mergers in this way mitigates the possibility that differences in treatment effects across the two groups are driven by differences in the extent to which they reduce competition. In balancing the tradeoff between minimizing differences across the two groups and reducing sample size, we opt to employ the maximum number of strata for which no COPA hospitals are dropped.

Matching on concentration measures addresses selection into behavioral remedies on concentration. Another form of selection that may bias our estimates is on cost synergies: e.g., hospitals may be approved to merge under COPA law by virtue of their large cost synergies. Rather than matching on additional variables which would further reduce sample size, we

demonstrate in Section 3 that selection on cost synergies of any kind is not consistent with negative price effects, negative quality effects, and a positive correlation between the two. More details on the matching process and measures described above can be found in Online Appendix B.

Table 1: Summary Statistics for Matched Treatment and Control Groups

| | Treatment group | | | |
|---|---|---|---|---|
| | COPA (1) | Non-COPA (2) | $p$-value (3) | Control group (4) |
| Number of hospitals | 39 | 166 | | 266 |
| *Outcome variables* | | | | |
| Inpatient price (thous. $) | 10.45 (7.04) | 8.96 (6.55) | 0.553 | 10.42 (11.41) |
| Cost per adjusted discharge (thous. $) | 9.88 (3.91) | 9.41 (5.72) | 0.754 | 11.58 (8.94) |
| 30-day mortality rate | 13.77 (2.03) | 13.03 (2.24) | 0.409 | 10.98 (4.05) |
| 30-day readmission rate | 20.10 (3.27) | 19.18 (2.83) | 0.422 | 18.00 (6.80) |
| Hourly wage | 22.74 (4.88) | 19.24 (7.51) | 0.050 | 23.06 (11.65) |
| Employment | 160.50 (201.07) | 141.05 (160.60) | 0.706 | 98.13 (155.45) |
| *Hospital characteristics* | | | | |
| % Medicaid | 10.26 (9.63) | 13.88 (11.01) | 0.306 | 16.39 (18.22) |
| % Outpatient charges | 60.29 (17.52) | 51.40 (20.38) | 0.075 | 34.08 (26.45) |
| Case-mix index | 1.49 (0.26) | 1.38 (0.29) | 0.036 | 1.29 (0.31) |
| Non-profit | 0.85 (0.37) | 0.71 (0.46) | 0.311 | 0.30 (0.46) |
| Number of beds | 132.92 (138.11) | 118.84 (114.33) | 0.599 | 105.34 (126.10) |
| *County characteristics* | | | | |
| Population (thous.) | 94.43 (86.24) | 201.27 (273.55) | 0.018 | 1138.42 (1823.04) |
| Share age 65 or older | 18.18 (3.98) | 14.79 (2.86) | 0.052 | 14.24 (4.09) |
| Per capita income (thous. $) | 36.63 (7.45) | 32.29 (9.32) | 0.060 | 40.99 (19.20) |
| One-bedroom rent | 561.19 (84.40) | 513.29 (140.01) | 0.173 | 676.05 (229.96) |
| Share rural | 0.10 (0.23) | 0.13 (0.21) | 0.589 | 0.05 (0.12) |
| *Other characteristics* | | | | |
| System size | 10.54 (7.09) | 5.56 (5.17) | 0.222 | 5.24 (4.47) |
| Pre-merger HHI | 3088.72 (1378.85) | 3184.62 (1417.56) | 0.788 | 2004.25 (1902.23) |
| Merger induced change in HHI | 410.29 (920.45) | 275.56 (706.64) | 0.679 | 0.00 (0.00) |

*Notes:* Values are for the year in which the hospital participated in its merger, or the first year in which it appears in the data. Cells show mean (standard deviation) for continuous measures and percentages for binary and categorical variables. HHI stands for Herfindahl–Hirschman index and is computed using bed shares for hospitals within 25 miles of one another. Non-COPA observations are weighted using weights from coarsened exact matching of non-COPA hospitals to COPA hospitals on HHI and $\Delta$HHI. $p$-values are for difference in means between COPA and non-COPA hospitals, computed with standard errors clustered at the system level.

Table 1 provides summary statistics for our analysis sample, broken into treatment and control groups. $p$-values test the null hypothesis that the COPA and non-COPA mergers have the same population means of the given variable. We see that the matching results in

comparable HHI levels and changes across the two treatment groups. COPA and non-COPA mergers are comparable with respect to rurality, per capita income, and rents, though COPA mergers tend to take place in counties with lower populations on average. Overall, these patterns suggest that the groups are largely balanced across many dimensions, minimizing potential confounding in our treatment effect comparisons.

# 5 Empirical Strategy

The exposition of our empirical strategy follows Borusyak et al. (2024). Let $D_{ht}$ be the binary treatment status of hospital $h$ in year $t$ relative to a specific merger event, and $Y_{ht}$ an outcome of interest. We represent our panel as a set of hospital-year observations $ht \in \Omega$. We define two sets of treated units, $\Omega_1^{COPA} = \{ht : D_{ht} = 1 \text{ and } h \in COPA\}$ and $\Omega_1^{nonCOPA} = \{ht : D_{ht} = 1 \text{ and } h \in nonCOPA\}$, which contain hospital-year units following COPA and non-COPA mergers, respectively. The complementary set of non-treated hospital-year pairs $\Omega_0 = \{ht : D_{ht} = 0\}$ contains both untreated units and treated hospitals in the years prior to merging. Let $Y_{ht}(0)$ be the stochastic potential outcome of hospital $h$ during year $t$ absent treatment.

For $\Omega_1 \equiv \Omega_1^{COPA} \cup \Omega_1^{nonCOPA}$, we define the causal effect of a merger on a single treated observation $ht \in \Omega_1$ as $\tau_{ht}^Y \equiv \mathbb{E}[Y_{ht} - Y_{ht}(0)]$. We are interested in estimands that are functions of treatment effects $\tau_{ht}^Y$ as well as of pre-specified weights $w_{ht}$ that are independent of realized outcomes. We make three assumptions in order to identify these estimands.

**Assumption 1** (Parallel Trends)**.** *For all* $ht \in \Omega$, $\mathbb{E}[Y_{ht}(0)] = \alpha_h + \beta_t + X_{ht}'\gamma$, *where* $X_{ht}$ *are time-varying covariates.*

Assumption 1 implies that hospitals' potential outcomes can be estimated using a two-way fixed-effects model, and that treated and untreated hospitals' potential outcomes trend in a parallel fashion. Selection into merging may result in violation of the parallel trends assumption, which we account for by controlling for time-varying area and hospital char-

acteristics. In Appendix 7, we provide illustrations of lead-lags estimates that allow for a visual inspection of the validity of the parallel trends assumption.

**Assumption 2** (No-anticipation effects)**.** $Y_{ht} = Y_{ht}(0)$ *for all* $ht \in \Omega_0$

This assumption implies that the outcomes observed for control units are informative about potential outcomes and that neither control nor pre-intervention observations are affected by the anticipation of merger events. Visual inspection of results from the lead-lag section can help verify the validity of this assumption.

**Assumption 3** (Model of causal effects)**.** *We allow for unrestricted individual effects for all treated units; i.e., for a vector of unknown parameters* $\theta$*,* $\tau^Y_{ht} = \theta_{ht}$ *for all* $ht \in \Omega_1$*.*

In principle, we could impose restrictions on the estimated effects by, for instance, allowing them to vary only across treatment groups, as in a standard "static" two-way fixed effects specification. While such restrictions may increase estimation efficiency, our setting necessitates a more agnostic approach that accounts for the idiosyncrasies of each merger event. Furthermore, estimating hospital-specific effects enables us to examine correlations between multiple outcomes, a key advantage of the Borusyak et al. (2024) methodology.

**Estimands:** Letting $\varepsilon_{ht} = Y_{ht} - \mathbb{E}[Y_{ht}]$ for all $ht \in \Omega$, our potential outcome model can be represented by the following equation:

$$Y_{ht} = \alpha_h + \beta_t + X'_{ht}\gamma + D_{ht}\tau^Y_{ht} + \varepsilon_{ht}. \tag{1}$$

The first of our two target estimands is the difference in average treatment effects (ATEs):

$$\begin{aligned} \tau^Y &= \sum_{ht \in \Omega_1^{COPA}} w_h^{COPA} \tau^Y_{ht} - \sum_{ht \in \Omega_1^{nonCOPA}} w_h^{nonCOPA} \tau^Y_{ht} \\ &= \tau^Y_{COPA} - \tau^Y_{nonCOPA}. \end{aligned} \tag{2}$$

Here, $w_h^{COPA}$ and $w_h^{nonCOPA}$ are the CEM weights for COPA and non-COPA hospitals, respectively.[5] $\tau^Y$ is our estimate for how behavioral remedies influence merger effects for outcome $Y$.[6]

Our second target estimand is the correlation between the treatment effects of two different outcomes. Specifically, we compute the correlation in treatment effects for outcomes $Y$ and $Y'$ for COPA hospitals as:

$$\rho_{COPA}^{Y,Y'} = \frac{\sum_{ht \in \Omega_1^{COPA}} (\tau_{ht}^Y - \bar{\tau}^Y)(\tau_{ht}^{Y'} - \bar{\tau}^{Y'})}{(T_1 \times N^{COPA} - 1) s_{COPA}^Y s_{COPA}^{Y'}}, \tag{3}$$

where $\bar{\tau}^Y$ and $s_{COPA}^Y$ are the mean and standard deviation of $\tau_{ht}^Y$ across $(h,t) \in \Omega_1^{COPA}$, respectively. This allows us to evaluate the theoretical prediction outlined in Section 3 that tighter price constraints are correlated with reductions in quality. An analogous version of Equation (3) is estimated for non-COPA hospitals.

**Estimator:** We compute the target estimands using an imputation-based approach. Specifically, we proceed in three steps. First, we estimate the parameters $(\alpha, \beta, \gamma)$ in Equation (1) by OLS, using only non-treated observations $ht \in \Omega_0$. For each treated observation $ht \in \Omega_1$, we then calculate the predicted counterfactual outcome $\hat{Y}_{ht}(0)$. We compute an observation-specific effect based on the difference between the observed and predicted outcome, $\hat{\tau}_{ht}^Y = Y_{ht} - \hat{Y}_{ht}(0)$. Given the vector of observation-specific effect estimates, we can compute target estimands according to Equation (2). Borusyak et al. (2024) show that under Assumptions 1-3, this procedure allows for an unbiased estimation of any linear combination of the treatment effects. To conduct inference, we follow the asymptotically conservative approach of Borusyak et al. (2024) to compute standard errors for ATEs.

The sample correlation of estimated treatment effects given by Equation (3) is poten-

[5] See Appendix B for more details on the construction of these weights.

[6] Equation (2) is operationally equivalent to a triple difference in estimating the differential treatment effect between two groups under the assumption that they would have followed parallel trends absent treatment, which we validate by evaluating whether leads for the difference in average treatment effects are statistically different from zero.

tially biased because it includes a component from the correlation of estimation errors. To correct this, we estimate treatment effects via a leave-one-out procedure where each hospital's treatment effects are computed using parameters estimated using all other hospitals. This approach eliminates the primary source of bias and yields a consistent estimator for the correlation of treatment effects as the number of hospitals grows large (see Online Appendix C for details).

# 6 Results

In Table 2, we present average treatment effect (ATE) estimates for six outcome variables: inpatient price, cost per adjusted discharge, 30-day mortality rates, 30-day readmission rates, hourly wages, and employment. ATEs for COPA hospitals are presented in column (1), for non-COPA hospitals in column (2), and the difference between COPA and non-COPA ATEs in column (3). Coefficients for log-transformed outcomes have been exponentiated and can be interpreted as percent changes.

We estimate that non-COPA mergers result in a 6% increase in price growth, which is in line with the 5.4% price effects from anticompetitive mergers from Brot et al. (2024). In contrast, we estimate that COPA mergers result in a 7.3% reduction in inpatient price growth on average relative to control hospitals. These estimates imply that behavioral remedies imposed as part of COPAs reduce inpatient price growth by 13.3 p.p. Effects on cost per adjusted discharge are negative but insignificant for both treatment groups.

We estimate statistically significant increases in 30-day mortality for both treatment groups on the order of 0.8 p.p. for COPA mergers and 0.5 p.p. for non-COPA mergers. The difference in treatment effects is not statistically significant for overall mortality rates, but is significant for heart attack and heart failure mortality, for which COPA mergers generate 0.9 p.p. and 0.7 p.p. greater increases in mortality than non-COPA ones, respectively.[7]

[7]This result is consistent with the findings of Kessler and McClellan (2000), who show that risk-adjusted one-year mortality for Medicare heart attack patients is significantly higher in more concentrated hospital

Table 2: Average Treatment Effects

| | COPA (1) | | Non-COPA (2) | | Difference in ATEs (1) minus (2) | |
|---|---|---|---|---|---|---|
| Inpatient price | $-0.073^{***}$ | (0.028) | $0.060^{**}$ | (0.028) | $-0.133^{***}$ | (0.041) |
| Cost per adjusted discharge | $-0.025$ | (0.022) | $-0.013$ | (0.014) | $-0.012$ | (0.026) |
| 30-day mortality rate | $0.008^{***}$ | (0.002) | $0.005^{**}$ | (0.002) | 0.003 | (0.002) |
| Heart attack | $0.006^{***}$ | (0.001) | $-0.002$ | (0.002) | $0.009^{***}$ | (0.002) |
| Heart failure | $0.010^{***}$ | (0.001) | $0.003^{**}$ | (0.001) | $0.007^{***}$ | (0.001) |
| Pneumonia | $0.006^{**}$ | (0.002) | $0.009^{***}$ | (0.002) | $-0.003$ | (0.003) |
| 30-day readmission rate | $0.010^{**}$ | (0.004) | $-0.005^{**}$ | (0.002) | $0.015^{***}$ | (0.004) |
| Heart attack | $-0.006$ | (0.003) | 0.002 | (0.003) | $-0.008^{**}$ | (0.004) |
| Heart failure | $0.020^{***}$ | (0.006) | 0.001 | (0.002) | $0.019^{***}$ | (0.006) |
| Pneumonia | $0.007^{*}$ | (0.004) | $-0.001$ | (0.002) | $0.008^{**}$ | (0.004) |
| Hourly wage | $0.081^{***}$ | (0.027) | 0.017 | (0.011) | $0.064^{**}$ | (0.027) |
| Blue collar | 0.034 | (0.019) | 0.016 | (0.009) | 0.018 | (0.020) |
| White collar | $0.103^{***}$ | (0.036) | 0.003 | (0.017) | $0.100^{**}$ | (0.039) |
| Nursing & pharmacy | $-0.023$ | (0.014) | $0.039^{**}$ | (0.019) | $-0.062^{***}$ | (0.023) |
| Employment | $-0.262^{***}$ | (0.058) | $-0.101^{***}$ | (0.028) | $-0.161^{**}$ | (0.063) |
| Blue collar | $-0.272^{***}$ | (0.056) | $-0.104^{***}$ | (0.033) | $-0.168^{***}$ | (0.063) |
| White collar | $-0.326^{***}$ | (0.069) | $-0.168^{***}$ | (0.036) | $-0.157^{**}$ | (0.076) |
| Nursing & pharmacy | 0.008 | (0.029) | $-0.072^{**}$ | (0.035) | $0.079^{**}$ | (0.035) |
| Number of observations | 3,267 | | | | | |

*Notes:* Table shows average treatment effects (ATEs) taken over the four years following a merger. Column (1) reports ATEs for hospitals undergoing COPA mergers. Column (2) reports ATEs for hospitals undergoing non-COPA mergers. Column (3) provides the difference between Columns (1) and (2). Averages are weighted using weights from coarsened exact matching of non-COPA hospitals to COPA hospitals on HHI and $\Delta$HHI. Standard errors clustered at the system level are provided in parentheses. $***p < 0.01, **p < 0.05, *p < 0.10$.

We likewise estimate that COPA mergers raise readmission rates by 1.5 p.p. more than non-COPA ones. These readmission effects stem primarily from a 2 p.p. increase in the likelihood of readmission for heart failure among hospitals undergoing COPA mergers.

We estimate positive and significant wage effects for both treatment groups on the order of 8.1% and 1.7% for COPA and non-COPA mergers, respectively. This differential is driven

markets. Similarly, Garmon and Bhatt (2022) document increases in heart-attack mortality following the Maine and Phoebe Putney COPA mergers when compared with a control group of non-merging hospitals.

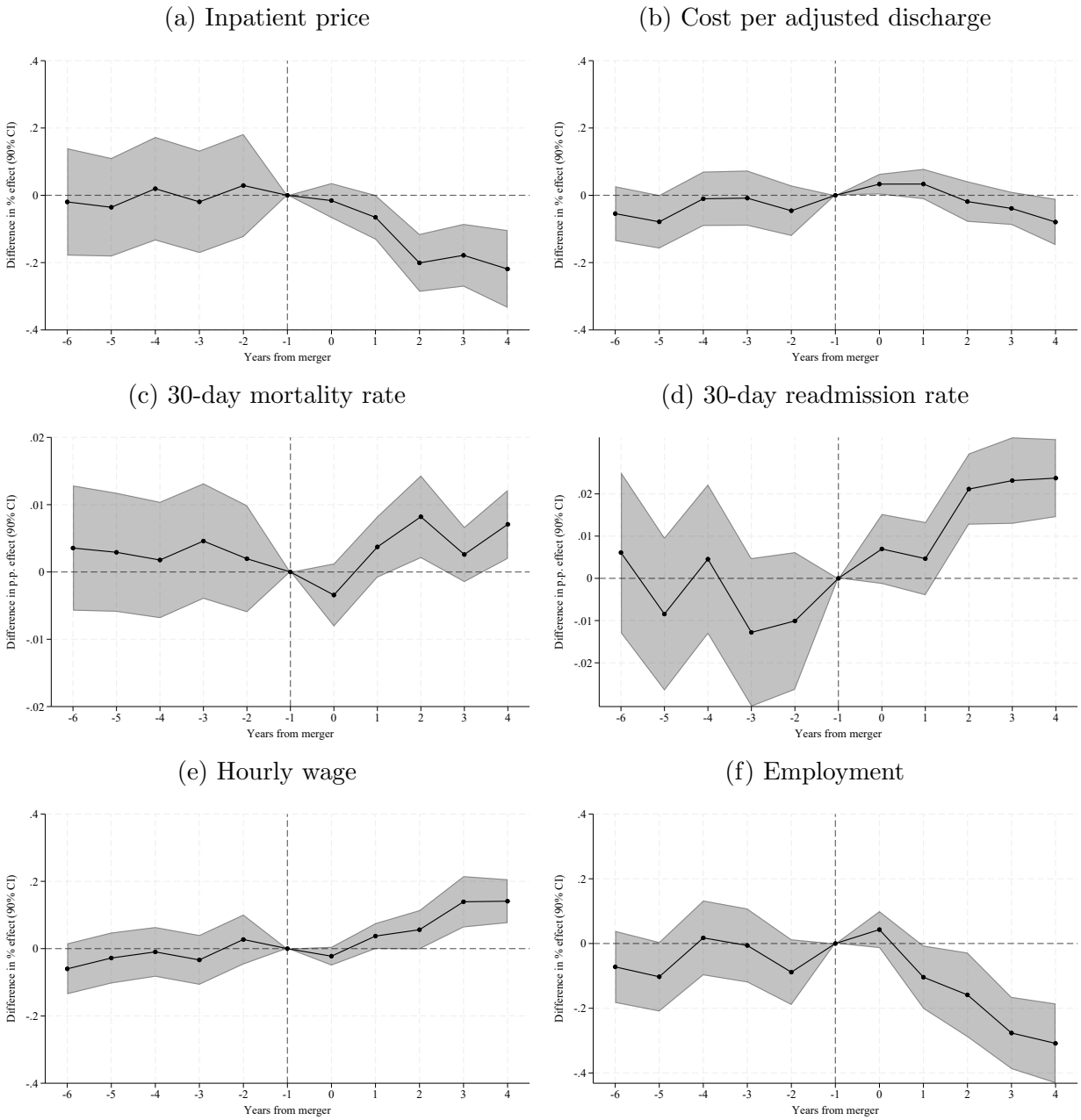
largely by a substantial 10.3% increase in the wages of white collar workers at COPA hospitals. Effects for employment suggest that the wage effects are most probably the result of post-merger changes in workforce composition: at both COPA and non-COPA hospitals, we estimate large negative employment effects for blue and white collar workers, though both of these effects are approximately 16 p.p. higher for COPA mergers relative to non-COPA ones. We estimate no effect of COPA mergers on nursing and pharmacy employment.

Leads and lags for the difference between COPA and non-COPA ATEs are presented in Figure 3 and for each treatment group's ATEs separately in Figure A1. Both sets of results support our assumptions of parallel trends and no anticipation.

**Results for hospitals experiencing relatively large increases in concentration:** To evaluate the extent to which the results we estimate in Table 2 are driven by market power, we estimate average treatment effects for hospitals in our sample that experienced relatively large increases in concentration. We subset the data while also maintaining balance between treatment groups by performing a ranked partition of the mutually exclusive strata developed as part of Coarsened Exact Matching into two sets, each containing approximately equal numbers of hospitals. Hospitals are considered to have experienced a relatively anticompetitive merger if they belong to the upper half of this ranked partition, which corresponds to strata with relatively high values of HHI and $\Delta$HHI.

Panel (a) of Table 3 summarizes the change in and resultant level of HHI for hospitals that underwent relatively anticompetitive mergers. The averages are substantially higher than those observed for the full sample but not statistically different from one another. We estimate that COPAs lower merger-induced price growth by 11.1 p.p. and result in increases in 30-day mortality and readmission rates that are 0.5 p.p. and 1.8 p.p. greater than those from non-COPA mergers, respectively. These results are consistent with the theory laid out in Section 3: if price constraints are more likely to be binding among hospitals experiencing relatively large declines in competition, then we should observe greater spillovers to the

Figure 3: Leads and Lags for Difference in Average Treatment Effects




*Notes:* Figures show differences in average treatment effects (ATEs) taken over hospitals within each of the six years preceding and four years following a hospital merger. Differences are between the ATEs for COPA and non-COPA mergers. Black dots provide the point estimates for the difference in ATEs. Gray regions correspond to 90% confidence intervals computed using standard errors clustered at the system level. Averages are weighted using weights from coarsened exact matching of non-COPA hospitals to COPA hospitals on HHI and $\Delta$HHI.

unregulated quality dimension among hospitals in this group.

Table 3: Average Treatment Effects for Hospitals Experiencing Relatively Large Increases in Concentration

| | COPA (1) | | Non-COPA (2) | | Difference (1) minus (2) | |
|---|---|---|---|---|---|---|
| *Panel (a): Mean (standard deviation) of change in and resultant HHI* | | | | | | |
| HHI | 3898.31 (988.24) | | 3883.15 (1107.77) | | -15.16 (296.77) | |
| ΔHHI | 569.13 (1100.95) | | 385.52 (824.84) | | -183.61 (239.37) | |
| *Panel (b): Coefficient (standard error) for average treatment effects* | | | | | | |
| Inpatient price | −0.041 | (0.039) | 0.071** | (0.028) | −0.111** | (0.049) |
| Cost per adjusted discharge | −0.016 | (0.020) | 0.002 | (0.016) | −0.018 | (0.025) |
| 30-day mortality rate | 0.010*** | (0.001) | 0.005*** | (0.002) | 0.005*** | (0.001) |
| 30-day readmission rate | 0.013*** | (0.004) | −0.005* | (0.003) | 0.018*** | (0.005) |
| Wages | 0.124*** | (0.024) | 0.020* | (0.011) | 0.104*** | (0.025) |
| Employment | −0.273*** | (0.041) | −0.134*** | (0.028) | −0.139*** | (0.047) |
| Number of hospitals | 25 | | 69 | | | |
| Number of observations | 2,449 | | | | | |

*Notes:* Panel (a) summarizes the HHI and merger-induced change in HHI for hospitals undergoing relatively large increase in concentration. Treated hospitals are divided into strata based on their values of HHI and ΔHHI as part of coarsened exact matching. We performed a ranked partition of these strata into two sets, each containing approximately equal numbers of hospitals. Hospitals are considered to have experienced a relatively anticompetitive merger if they belong to the upper half of this ranked partition. Panel (b) shows average treatment effects (ATEs) taken over the four years following a merger for hospitals undergoing relatively large declines in competition. Column (1) reports ATEs for hospitals undergoing COPA mergers. Column (2) reports ATEs for hospitals undergoing non-COPA mergers. Column (3) provides the difference between Columns (1) and (2). Averages are weighted using weights from coarsened exact matching of non-COPA hospitals to COPA hospitals on HHI and ΔHHI. Standard errors clustered at the system level are provided in parentheses. Bottom panel provides the unique number of COPA and non-COPA hospitals experiencing relatively large declines in concentration, respectively, as well as the number of hospital-year pairs used in estimation of ATEs. $***p < 0.01, **p < 0.05, *p < 0.10$.

**Robustness** We perform several robustness checks for our main results. Table A2 provides ATEs for inpatient price and mortality using 10- and 30-mile radius as alternative market definitions. Because the number of COPA mergers is small, in Table A3, we check that no single COPA merger drives our results by dropping all hospitals that participate in a given COPA merger, one merger at a time, and re-estimating our ATEs. In Table A4, we provide results generated by dropping each member of our set of time-varying controls. This last check serves as a test of the assumption of the Borusyak et al. (2024) estimator that controls are unaffected by treatment. Our results are robust across all of these checks. We provide further evidence of the exogeneity of our controls in Table A5, which presents ATEs using as

outcomes each of our hospital-level time-varying controls. We find no significant differences for any of these outcomes.

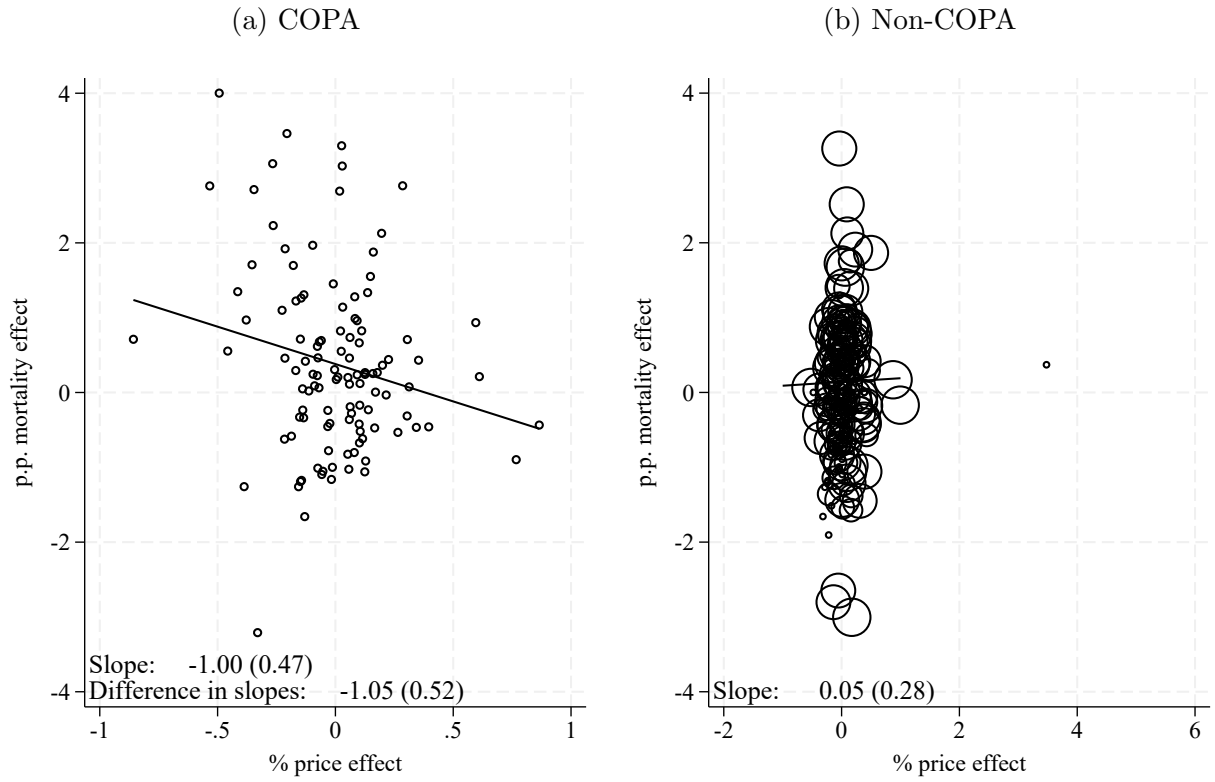
**Correlation between price and mortality effects:** The results above indicate that price and quality effects from COPAs are negative. In our final analysis, we verify that these results are reflective of a tradeoff over regulated and unregulated dimensions of hospital behavior by estimating the correlation between individual-level price and mortality effects separately for COPA and non-COPA mergers. Based on our theory from Section 3, our hypothesis is that this correlation will be negative for COPA mergers.

Results are illustrated in Figure 4. Panel (a) provides a scatterplot of individual (hospital-year) 30-day mortality effects against individual price effects. A line of best fit is drawn through these points, the slope of which is provided in the bottom left corner of the plot. We estimate a statistically significant slope between mortality and price effects for COPA mergers, implying that 1 p.p. lower merger-induced price growth is associated with a 1 p.p. increase in 30-day mortality.

Panel (b) presents the analogous results for non-COPA hospitals, where treatment effects are weighted by each hospital's CEM weight. We estimate no statistically significant association between price and mortality effects for non-COPA mergers, which is perhaps reflective of substantial variation across transactions in the magnitude of reductions in quality elasticities relative to price elasticities.

Figure A2 presents the correlations between price and readmission effects. The difference between correlations is negative, but is not statistically significant. Figure A3 shows that the negative correlation between price and mortality effects among COPA mergers is maintained when using a 30-mile radius in defining markets.

Figure 4: Correlations between Price and Mortality Effects




*Notes:* Panel (a) provides a scatter plot of bias-corrected individual (hospital-year) 30-day mortality effects against bias-corrected individual inpatient price effects across all hospitals undergoing COPA mergers. Panel (b) provides the analogous scatter plot for non-COPA mergers, where treatment effects are weighted and dot size is proportional to weights from Coarsened Exact Matching. Bias-corrected treatment effects are constructed by applying leave-one-out bias correction to individual treatment effect estimates (see Appendix C). The slope estimate in the bottom left corner of each panel is the coefficient from a regression of individual mortality effects on individual price effects and corresponds to the slope of the line of best fit. The difference between estimated slopes for Panels (a) and (b) is provided in the bottom right corner of Panel (b). Standard errors clustered at the merger level shown in parentheses.

**Consent decrees:** As discussed in Section 1, COPAs are one of two types of conditional agreements state regulators may enter into with merging hospitals, the other being consent decrees. Mergers under state-issued consent decrees are not as well documented and employ a wider variety of conditions on prices. Montague, Davison, Gudiksen, and King (2023) identify five state-issued consent decrees issued between 2012 and 2022. All five regulated post-merger prices, though only one does so through the imposition of post-merger price caps, with the remainder requiring "good faith negotiations" with payers. Online Appendix

D provides a summary of these consent decree mergers as well as ATEs from models estimated with the inclusion of a third treatment group of consent decree mergers. Results for COPA and non-COPA mergers are stable relative to our main results, and estimated effects for consent decree mergers are mixed and largely insignificant, reflecting the large degree of heterogeneity in this group relative to its sample size.

# 7 Discussion and Conclusion

This paper evaluates how behavioral remedies enforced by state regulators impact the effects of hospital mergers. We compare several outcomes for hospitals that merge under Certificates of Public Advantage, which regulate some but not all dimensions of merging hospitals' behavior, to unregulated mergers that generate similar changes in concentration.

We find evidence that COPAs constrain prices but do so at the expense of quality. For both COPA mergers and their unregulated counterparts, treatment effects for wages and employment are reflective of substantial organizational restructuring following consolidation: employment, particularly of blue- and white-collar workers, is reduced, which raises average hourly wages. These effects are more pronounced among COPA mergers.

Reductions in employment following mergers may reflect cost synergies but also disinvestments in quality. In particular, the substantial reductions in blue-collar employment—which includes housekeeping, dietary, and maintenance staff—combined with cuts to white-collar administrative support, may contribute to the observed increases in mortality rates by reducing the workforce responsible for hospital cleanliness, patient nutrition, and care coordination.

One piece of evidence that disinvestments in quality are part of the difference in labor market effects between the two treatment groups is that increases in mortality are larger among COPA mergers. In the case of heart attacks, which Gaynor (2007) refers to as the "canary in the mine shaft" for possible "general slack in the hospital," we estimate that

COPA mergers generate 0.9 percentage point greater increases in 30-day mortality than similar unregulated mergers. Another piece of evidence is that price effects are negatively correlated with mortality effects across COPA mergers, consistent with lower margins incentivizing lower post-merger quality. The same association is not observed for non-COPA mergers.

A natural question is how the benefits of constrained price growth compare to the lost value of life under COPAs. Using pre-merger levels of commercial spending and mortality rates at COPA hospitals along with the ATEs from Table 2, we conclude that COPAs generate average annual spending reductions of $551 million, but result in an average annual loss in value of statistical life of $575 million[8]. Together, these results indicate that COPAs are poor substitutes for federal antitrust scrutiny.

To provide an extensive evaluation of COPA cases, our analysis is limited in that it focuses on variables that are broadly available across the universe of hospitals, e.g., price information for only inpatients, and quality measures constructed from only Medicare patients' information. Although we do not see those limitations critically restricting the validity of our results, we do see room for future research that can access claim-level data for specific COPA cases to confirm the insights we provide here. Moreover, the impact of COPA mergers on extensive margin outcomes, such as hospital exit, is interesting but outside the scope of this paper.

Overall, the limitations of state-imposed behavioral constraints that we document here have implications for the application of state-action immunity under Parker (1943) that go beyond COPAs. The Supreme Court has established two conditions for states to confer antitrust immunity: the state must (i) clearly articulate a policy to displace competition

[8]Average change in hospital spending on commercial inpatients computed as $\frac{0.133}{T_0 \times N^{COPA}} \times \sum_{ht \in \Omega_0^{COPA}} Spend_{ht}$ where $Spend$ is inpatient commercial spending from the CMS hospital cost reports. Average change in value of statistical life is computed as $\sum_Y \left( \frac{\tau^Y_{COPA} VSL}{T_0 \times N^{COPA}} \times \sum_{ht \in \Omega_0^{COPA}} Denominator^Y_{ht} \right)$ where $Denominator^Y$ is the volume of patients admitted with condition $Y \in$ {heart attack, heart failure, pneumonia} from the Hospital Compare Data and $VSL$ is our estimate of value of statistical life for Medicare beneficiaries. We use a conservative estimate of the value of statistical life for 60-year olds from Aldy and Viscusi (2003) of $2.5 million.

and (ii) actively supervise the private conduct in question. The Court has set a high bar for what constitutes active supervision, emphasizing that states must "exercise ultimate control over the challenged anticompetitive conduct" and that "the mere presence of some state involvement or monitoring does not suffice" (California Retail Liquor Dealers Ass'n v. Midcal Aluminum, Inc., 1980). Our findings suggest that COPA contracts do not facilitate effective oversight of all relevant dimensions of hospital behavior, calling into question the legitimacy of state-action protections in this context.

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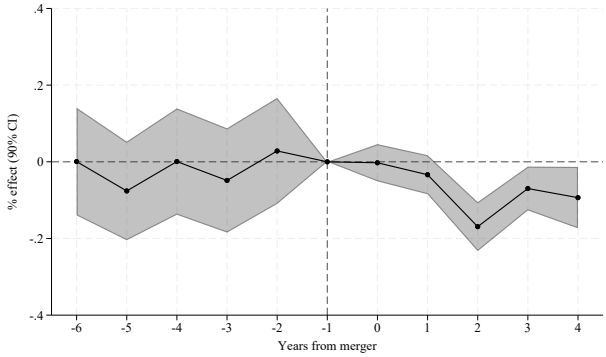
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# Appendix Tables and Figures

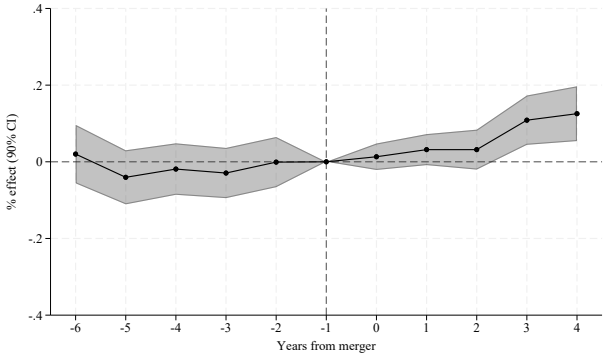
Figure A1: Leads and Lags for Average Treatment Effects

(a) Inpatient price

(a1) COPA

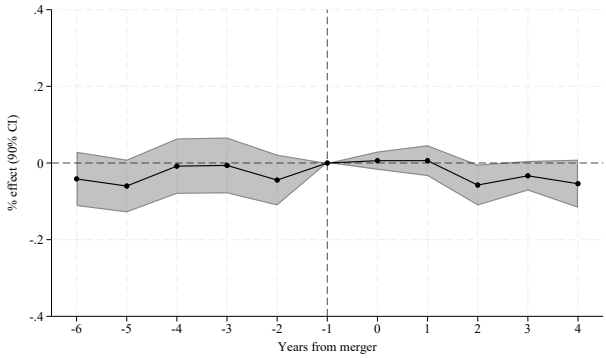



(a2) Non-COPA

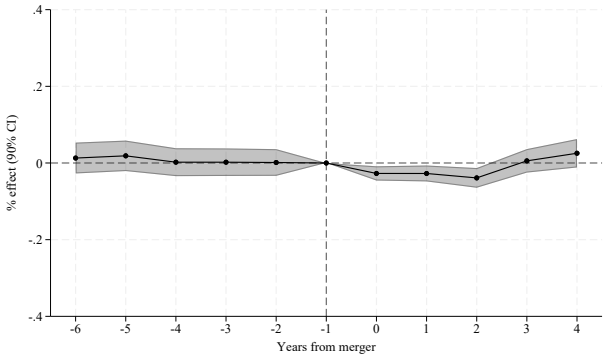



(b) Cost per adjusted discharge

(b1) COPA

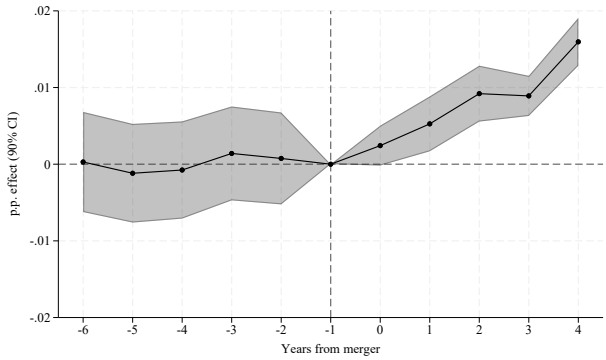



(b2) Non-COPA

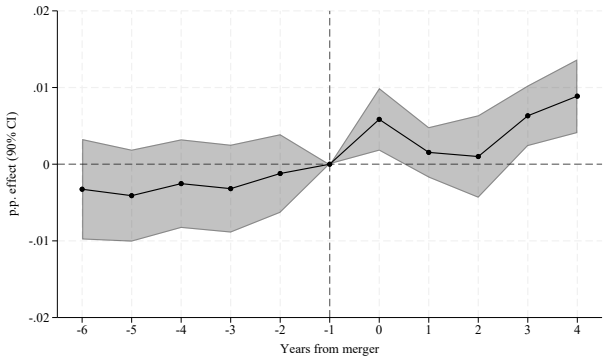



(c) 30-day mortality rate

(c1) COPA



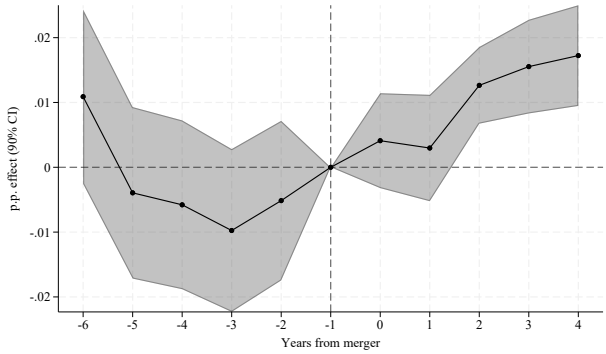

(c2) Non-COPA



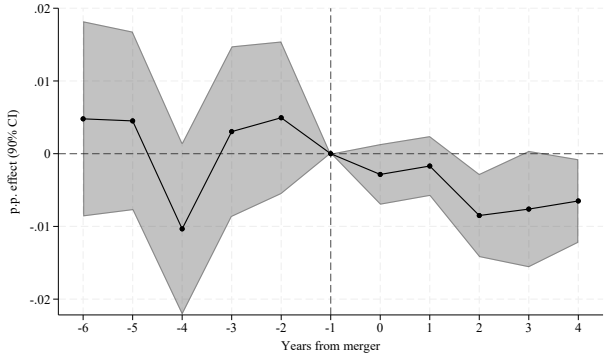
Leads and Lags for Average Treatment Effects, continued

(d) 30-day readmission rate

(d1) COPA

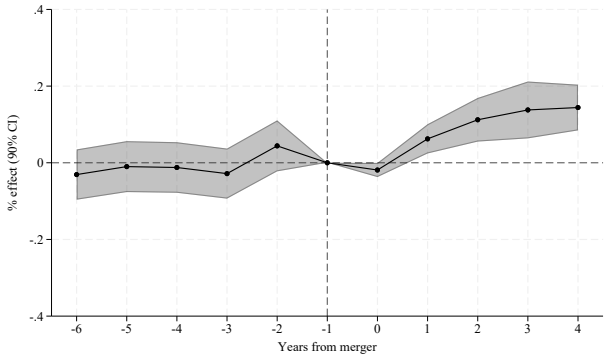



(d2) Non-COPA

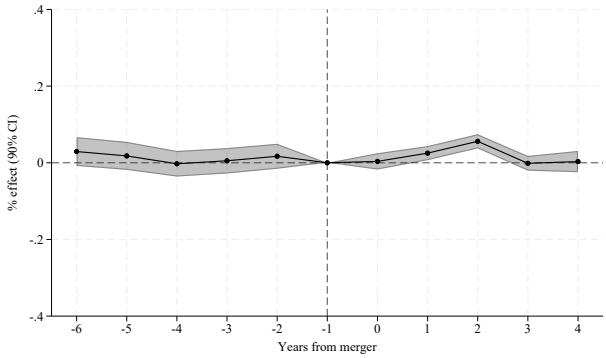



(e) Wages

(e1) COPA

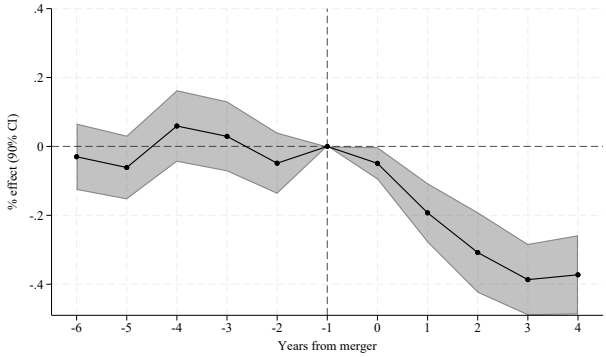



(e2) Non-COPA

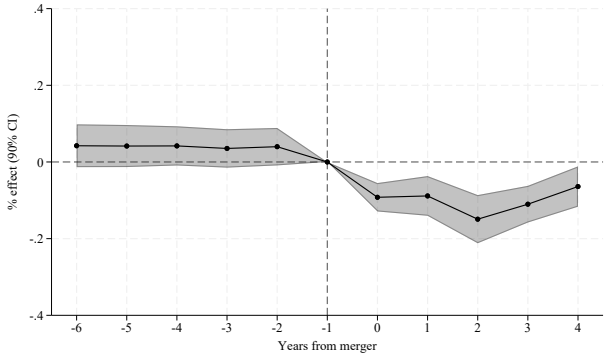



(f) Employment

(f1) COPA

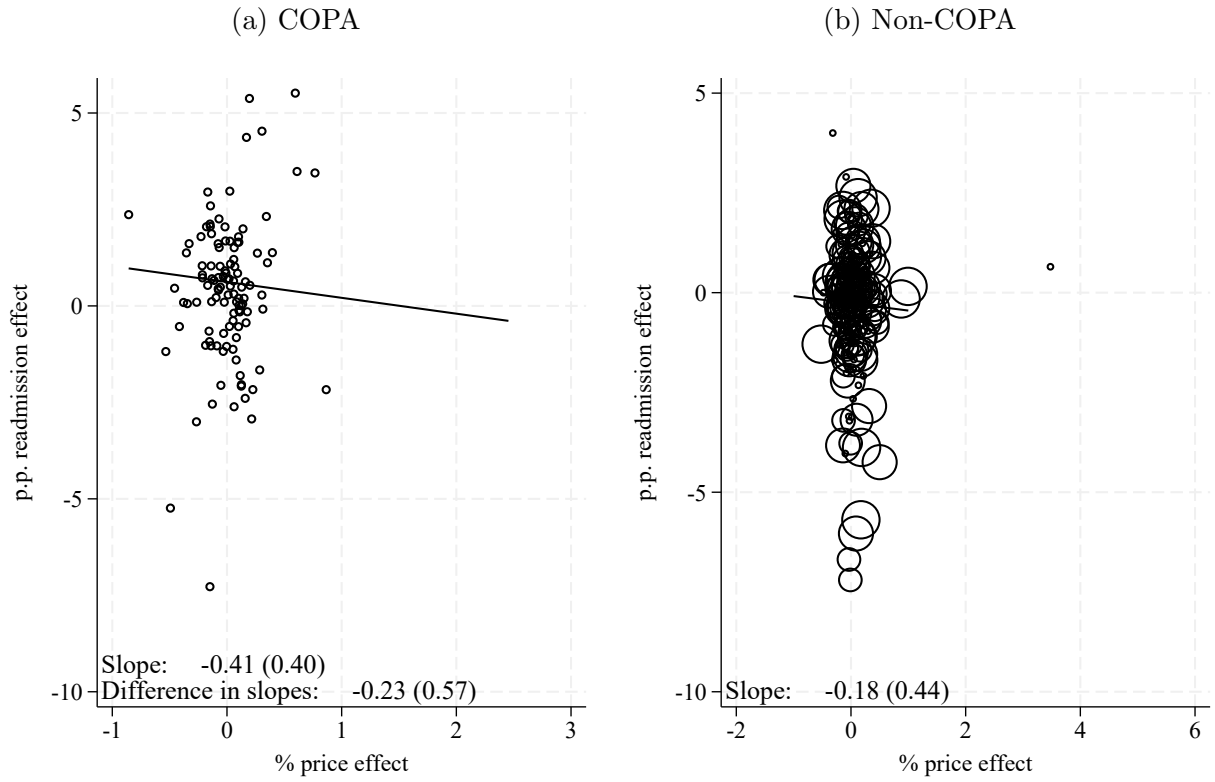



(f2) Non-COPA



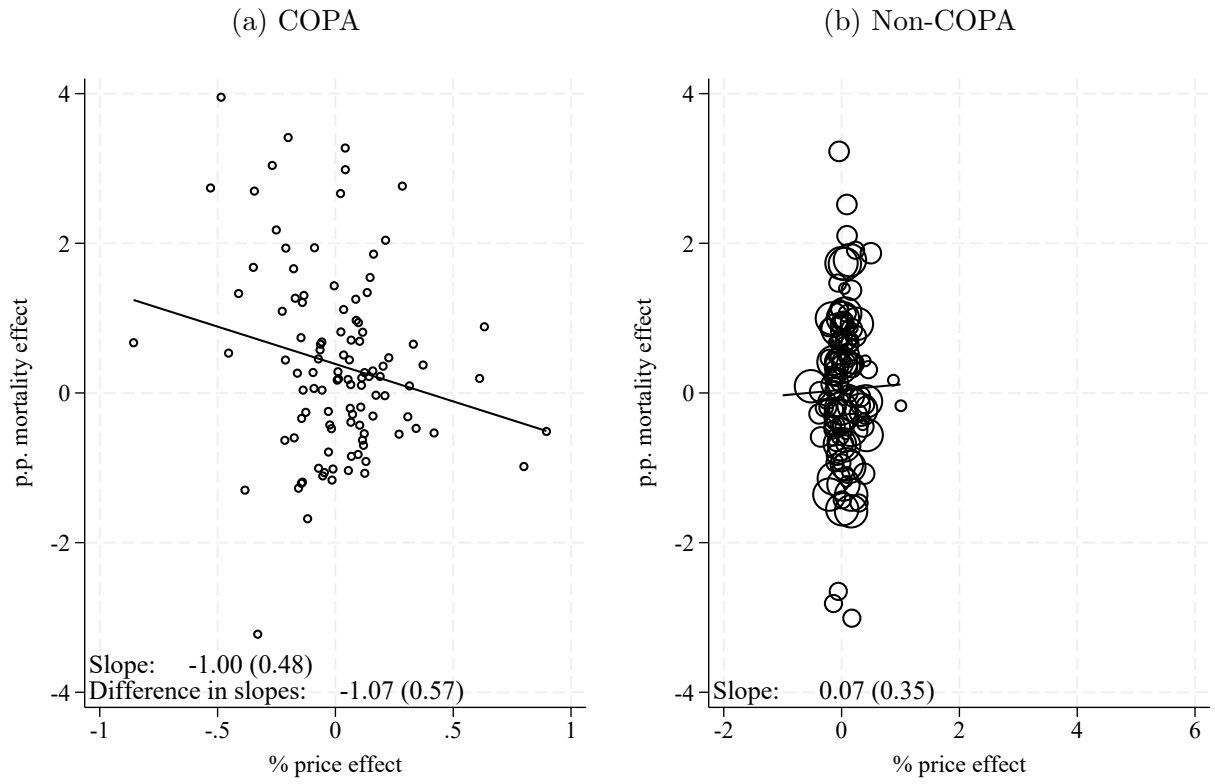

*Notes:* Figures show average treatment effects (ATEs) taken over hospitals within each of the six years preceding and four years following a hospital merger. Averages are plotted across time for COPA mergers in the left column and for non-COPA mergers in the right column. Black dots provide the point estimates for the difference in ATEs. Gray regions correspond to 90% confidence intervals computed using standard errors clustered at the system level. Averages are weighted using weights from coarsened exact matching of non-COPA hospitals to COPA hospitals on HHI and $\Delta$HHI.

Figure A2: Correlations between Price and Readmission Effects




*Notes:* Panel (a) provides a scatter plot of bias-corrected individual (hospital-year) 30-day readmission effects against bias-corrected individual inpatient price effects across all hospitals undergoing COPA mergers. Panel (b) provides the analogous scatter plot for non-COPA mergers, where treatment effects are weighted and dot size is proportional to weights from Coarsened Exact Matching. Bias-corrected treatment effects are constructed by applying leave-one-out bias correction to individual treatment effect estimates (see Appendix C). The slope estimate in the bottom left corner of each panel is the coefficient from a regression of individual readmission effects on individual price effects and corresponds to the slope of the line of best fit. The difference between estimated slopes for Panels (a) and (b) is provided in the bottom right corner of Panel (b). Standard errors clustered at the merger level in parentheses.

Figure A3: Correlations between Price and Mortality Effects Using 30-Mile Radius as Market Definition




*Notes:* Results are generated using 30-mile radius in defining markets. Panel (a) provides a scatter plot of bias-corrected individual (hospital-year) 30-day mortality effects against bias-corrected individual inpatient price effects across all hospitals undergoing COPA mergers. Panel (b) provides the analogous scatter plot for non-COPA mergers, where treatment effects are weighted and dot size is proportional to weights from Coarsened Exact Matching. Bias-corrected treatment effects are constructed by applying leave-one-out bias correction to individual treatment effect estimates (see Appendix C). The slope estimate in the bottom left corner of each panel is the coefficient from a regression of individual mortality effects on individual price effects and corresponds to the slope of the line of best fit. The difference between estimated slopes for Panels (a) and (b) is provided in the bottom right corner of Panel (b). Standard errors clustered at the merger level in parentheses.

Table A1: Comparison of Changes in and Resultant HHI for Hospitals that Undergo COPA and non-COPA Mergers

| | COPA (1) | Non-COPA (2) | Difference (1) minus (2) |
|---|---|---|---|
| HHI | 3088.71 (220.79) | 1849.40 (125.48) | 1239.31 (280.25) |
| ΔHHI | 410.29 (147.38) | 68.99 (22.59) | 341.30 (84.86) |

*Notes:* Values are for the year in which the hospital participated in its merger. Cells show mean (standard error). HHI stands for Herfindahl–Hirschman index and is computed using bed shares for hospitals within 25 miles of one another.

Table A2: Difference in Average Treatment Effects Using Alternative Market Definitions

| | All (1) | | Relatively large increase in concentration (2) | |
|---|---|---|---|---|
| 10-mile radius | | | | |
| Inpatient price | $-0.110^{***}$ | (0.036) | $-0.003$ | (0.053) |
| 30-day mortality rate | $0.007^{***}$ | (0.002) | $0.008^{***}$ | (0.003) |
| 30-mile radius | | | | |
| Inpatient price | $-0.103^{***}$ | (0.037) | $-0.073$ | (0.047) |
| 30-day mortality rate | $0.006^{***}$ | (0.002) | $0.006^{***}$ | (0.001) |

*Notes:* Table shows the difference between average treatment effects (ATEs) for COPA and non-COPA mergers taken over the four years following a merger defining a hospital's market using a 10-mile radius (Column 1) and 30-mile radius (Column 2) rather than a 25-mile radius as in the main specifications. Averages are weighted using weights from coarsened exact matching of comparison group hospitals to COPA hospitals on HHI and $\Delta$HHI. Standard errors clustered at the system level are shown in parentheses. $***p < 0.01, **p < 0.05, *p < 0.10$.

Table A3: Difference in Average Treatment Effects Leaving Out One COPA Merger at a Time

| | All (1) | | Relatively large increase in concentration (2) | |
|---|---|---|---|---|
| Palmetto | | | | |
| Inpatient price | $-0.102^{**}$ | (0.046) | $-0.111^{**}$ | (0.049) |
| 30-day mortality rate | 0.003 | (0.002) | $0.005^{***}$ | (0.001) |
| Maine Health | | | | |
| Inpatient price | $-0.158^{***}$ | (0.045) | $-0.146^{**}$ | (0.055) |
| 30-day mortality rate | 0.003 | (0.002) | $0.005^{**}$ | (0.002) |
| Phoebe Putney | | | | |
| Inpatient price | $-0.129^{**}$ | (0.044) | $-0.097^{*}$ | (0.057) |
| 30-day mortality rate | 0.002 | (0.002) | $0.004^{**}$ | (0.002) |
| Cabell Huntington | | | | |
| Inpatient price | $-0.136^{***}$ | (0.040) | $-0.082^{*}$ | (0.042) |
| 30-day mortality rate | 0.004 | (0.002) | $0.005^{***}$ | (0.001) |
| Ballad | | | | |
| Inpatient price | $-0.128^{***}$ | (0.032) | $-0.097^{**}$ | (0.039) |
| 30-day mortality rate | 0.000 | (0.002) | $0.006^{***}$ | (0.001) |
| Hendrick | | | | |
| Inpatient price | $-0.133^{*}$ | (0.041) | $-0.120^{**}$ | (0.050) |
| 30-day mortality rate | $0.004^{*}$ | (0.002) | $0.005^{***}$ | (0.002) |
| Shannon | | | | |
| Inpatient price | $-0.138^{***}$ | (0.041) | $-0.119^{**}$ | (0.050) |
| 30-day mortality rate | 0.003 | (0.002) | $0.005^{***}$ | (0.001) |

*Notes:* Table shows the difference between average treatment effects (ATEs) for COPA and non-COPA mergers taken over the four years following a merger defining a hospital's market using all treated hospitals (Column 1) those experiencing relatively large increases in concentration (Column 2) where the latter are defined as in Table 3. Each panel computes ATEs excluding hospitals that participate in the corresponding COPA merger. Averages are weighted using weights from coarsened exact matching of comparison group hospitals to COPA hospitals on HHI and $\Delta$HHI. Standard errors clustered at the system level are shown in parentheses. $***p < 0.01, **p < 0.05, *p < 0.10$.

Table A4: Difference in Average Treatment Effects Leaving Out One Control Variable at a Time

| | All (1) | | Relatively large increase in concentration (2) | |
|---|---|---|---|---|
| % Medicaid | | | | |
| Inpatient price | −0.139*** | (0.042) | −0.125** | (0.049) |
| 30-day mortality rate | 0.003 | (0.002) | 0.004*** | (0.001) |
| % Outpatient charges | | | | |
| Inpatient price | −0.154*** | (0.045) | −0.122** | (0.053) |
| 30-day mortality rate | 0.003 | (0.002) | 0.005*** | (0.001) |
| Case-mix index | | | | |
| Inpatient price | −0.101*** | (0.034) | −0.090* | (0.047) |
| 30-day mortality rate | 0.002 | (0.002) | 0.005*** | (0.002) |
| Non-profit | | | | |
| Inpatient price | −0.134*** | (0.041) | −0.115** | (0.049) |
| 30-day mortality rate | 0.003 | (0.002) | 0.005*** | (0.001) |
| ln(Beds) | | | | |
| Inpatient price | −0.131*** | (0.041) | −0.108** | (0.049) |
| 30-day mortality rate | 0.003 | (0.002) | 0.005*** | (0.001) |
| ln(Population) | | | | |
| Inpatient price | −0.135*** | (0.040) | −0.119** | (0.048) |
| 30-day mortality rate | 0.003 | (0.002) | 0.005*** | (0.001) |
| Share age 65 or older | | | | |
| Inpatient price | −0.133*** | (0.041) | −0.112** | (0.049) |
| 30-day mortality rate | 0.003 | (0.002) | 0.005*** | (0.001) |
| ln(Per capita income) | | | | |
| Inpatient price | −0.151*** | (0.043) | −0.146*** | (0.046) |
| 30-day mortality rate | 0.005 | (0.002) | 0.007*** | (0.002) |
| ln(One-bedroom rent) | | | | |
| Inpatient price | −0.133*** | (0.041) | −0.111** | (0.049) |
| 30-day mortality rate | 0.003 | (0.002) | 0.005*** | (0.002) |

*Notes:* Table shows the difference between average treatment effects (ATEs) for COPA and non-COPA mergers taken over the four years following a merger defining a hospital's market using all treated hospitals (Column 1) those experiencing relatively large increases in concentration (Column 2) where the latter are defined as in Table 3. Each panel computes ATEs excluding the corresponding variable from the set of time-varying controls. Averages are weighted using weights from coarsened exact matching of comparison group hospitals to COPA hospitals on HHI and ΔHHI. Standard errors clustered at the system level are shown in parentheses. $***p < 0.01, **p < 0.05, *p < 0.10$.

Table A5: Average Treatment Effects for Hospital-level Time-varying Controls

| | COPA (1) | | Non-COPA (2) | | Difference in ATEs (1) minus (2) | |
|---|---|---|---|---|---|---|
| % Medicaid | $0.014^{**}$ | (0.007) | $0.021^{**}$ | (0.010) | $-0.008$ | (0.010) |
| % Outpatient charges | $0.011^{*}$ | (0.006) | 0.002 | (0.006) | 0.009 | (0.009) |
| Case-mix index | $0.000^{**}$ | (0.000) | 0.000 | (0.000) | 0.000 | (0.000) |
| Non-profit | $0.001^{***}$ | (0.000) | $0.001^{***}$ | (0.000) | $-0.000$ | (0.000) |
| ln(Beds) | $-0.002^{***}$ | (0.000) | $-0.000^{*}$ | (0.000) | $-0.001^{***}$ | (0.000) |
| Number of observations | 3,267 | | | | | |

*Notes:* Table shows average treatment effects (ATEs) for hospital-level time-varying controls taken over the four years following a merger. Column (1) reports ATEs for hospitals undergoing COPA mergers. Column (2) reports ATEs for hospitals undergoing non-COPA mergers. Column (3) provides the difference between Columns (1) and (2). Averages are weighted using weights from coarsened exact matching of non-COPA hospitals to COPA hospitals on HHI and $\Delta$HHI. Standard errors clustered at the system level are shown in parentheses. $***p < 0.01, **p < 0.05, *p < 0.10$.

# Online Appendix A COPA Merger Details

In what follows, we present a short description of each COPA merger that has been consummated. The first consummated COPA merger was the merger of HealthSpan and Medina in Minnesota in 1994. This was a vertical merger of hospitals, health insurance, and physician groups. Since this paper is concerned with horizontal integration in hospital markets, we omit the HealthSpan merger from our discussion.

- **Mission, 1995:** In the early 1990s, Mission Hospital and St. Joseph's Hospital were Asheville's only two private acute care hospitals. They lobbied the North Carolina General Assembly to enact COPA legislation in 1993 to facilitate a partnership, claiming they would maintain separate identities. The hospitals lobbied for COPA amendments in 1995 to gain antitrust immunity. Mission bought St. Joseph's outright in 1998, with COPA amended again to facilitate the merger (United States District Court Western District of North Carolina, 2022). Mission was subject to a margin cap which stipulated that "the operating margin of MHS over any three-year period shall not exceed by more than one percent the mean of the median operating margin of comparable hospitals (provided that this cap will not fall below three percent)" (Vistnes, 2011; Bovbjerg and Berenson, 2015). By 1998, the merged entity controlled 91% of Madison County and 87% of Buncombe County admissions, growing to 93% in the Asheville Region by 2016. North Carolina repealed its COPA law in late 2015, effective in January 2018, ending price oversight while leaving Mission's monopoly intact (Khan et al., 2022).

- **Benefis, 1996:** In July 1996, Montana's Department of Justice approved a COPA allowing Columbus Hospital and Montana Deaconess Medical Center—Great Falls' only two acute care hospitals—to merge and form Benefis Health System. The COPA "include[d] the establishment of a 'patient revenue cap' to ensure that the consolidated hospitals do not generate revenues in excess of those sufficient to provide the profit margin approved by the Department [of Justice]." Regulators "expect[ed] the COPA's requirements to re-

sult in price reductions of approximately 18% to 23% during the first four years after the consolidation. The Department will conduct an annual audit to assure proper implementation of this rate regulation" (Leibenluft, 1996). In 2007, Benefis lobbied the Montana legislature to terminate the COPA agreement, despite objections from the Montana Attorney General. The legislation took effect in 2009, ending regulatory oversight (Khan et al., 2022).

- **Palmetto, 1998:** Palmetto Health was formed in 1998 through the merging of Baptist Healthcare and Richland Memorial Hospital. These hospitals comprised 66 percent of the local market for acute care services in Columbia, South Carolina at the time. The COPA required Palmetto to reduce gross charges to all payers, invest in outreach programs, and submit to annual audits until 2003 when many of the COPA terms were repealed in 2003 (Garmon and Bhatt, 2022; Bhatt, 2019).

- **MaineHealth, 2009:** The first COPA merger of the 21st century was the acquisition of Southern Maine Medical Center (SMMC) by Maine Health in 2009. The flagship general acute care hospital of Maine Health, Maine Medical Center, was located 20 miles northeast of SMMC. The COPA required that "During the term of the certificate, SMMC will set its annual consolidated budgeted operating margin at a level less than or equal to 3% of its total operating revenue" (Maine Department of Health and Human Services, 2009). The COPA did not impose any restrictions on other MaineHealth hospitals and was repealed in 2015 (Garmon and Bhatt, 2022).

- **Phoebe Putney, 2011**: Although it was not a COPA merger per se, we include the Phoebe Putney merger in our analysis sample as it received exemption from antitrust enforcement action despite the FTC's efforts to block the merger. Phoebe Putney had long been managed by the Albany-Dougherty County Authority when it acquired Palmyra Medical Center in Albany, Georgia in 2011. The FTC challenged the merger, but its challenge was struck down when the courts determined that the county Health Authority was

exempt from antitrust scrutiny under the state-action immunity. In 2015, the FTC and Phoebe Putney settled an agreement (1) "requiring Phoebe Putney and the Hospital Authority must notify the FTC in advance of acquiring any part of a hospital or a controlling interest in other healthcare providers in the Albany, Georgia area for the next 10 years," and (2) "prohibiting them from objecting to regulatory applications made by potential new hospital providers in the same area for up to five years" (FTC, 2014).

- **Cabell Huntington, 2018**: In 2018, Cabell Huntington Hospital acquired St. Mary's Medical Center under the West Virginia COPA law. Both target and acquirer were located in the same county. The West Virginia Health Care Authority required that "Neither CHH nor SMMC [seek] an increase in Hospital Rates beyond Benchmark Rates established by the West Virginia Health Care Authority...through a mathematical process that computes a charge increase based on a provider charge and cost comparison within a peer group" (West Virginia Health Care Authority, 2016; Office of the West Virginia Attorney General, 2015). This COPA is set to expire in 2025.

- **Ballad, 2018:** Ballad Health was established in 2018 following the merger of Mountain States Health Alliance and Wellmonth Health System. These systems spanned northeastern Tennessee and southwestern Virginia, and the merger was approved under the two states' COPA laws. The COPA conditions placed "limits upon, measurement, and reporting of price increases for specific services, including hospital inpatient and outpatient, non-hospital outpatient, physician and physician extender, charge-based and cost-based services" (Tennessee Department of Health, 2018; Virginia Department of Health, 2017).

- **Shannon and Hendrick, 2020:** In 2020, the Texas legislature provided two hospital merger antitrust exemptions under its COPA law. One was the merger of Shannon Medical Center and San Angelo Community Medical Center in Tom Green County; the other was the merger of Hendrick Medical Center and Abilene Regional Medical Center in Taylor County. The state authority approved these mergers' antitrust immunity applications

despite the FTC's public concern that the mergers would harm competition. The Texas COPAs "Prohibit an increase in rates for hospital services by a hospital operating under a certificate of public advantage...from taking effect without prior approval of HHSC....." (Texas Health and Human Services Commission, 2023; Texas Department of State Health Services, 2024).

# Online Appendix B Additional Information on Data and Matching

**Inpatient prices:** Using data from the CMS hospital cost reports and following the approach described in Dafny (2009) and used in numerous other papers (Garmon and Bhatt, 2022; Dafny et al., 2019; Garmon, 2017; Lewis and Pflum, 2017), we construct the inpatient price for hospital $h$ in year $t$ as

$$\text{price} = \frac{\text{inpatient revenue} \times \left(1 - \frac{\text{contractual discounts}}{\text{total charges}}\right) - \text{Medicare revenue}}{\text{total inpatient discharges} - \text{Medicare inpatient discharges}}$$

where

$$\begin{aligned}\text{inpatient revenue} =& \text{inpatient routine service charges} + \text{inpatient intensive care charges} + \\ & \text{inpatient ancillary service charges}\end{aligned}$$

and

$$\text{Medicare revenue} = \text{Medicare primary payer amounts} + \text{Medicare total amount payable.}$$

A strength of using this measure of prices is that it can be computed for all hospitals for the entirety of our sample period. However, this measure is computed as the net revenue for

non-Medicare inpatient admissions divided by the number of non-Medicare admissions and does not exclude revenues from non-Medicare, non-commercial sources, such as Medicaid. We address this issue by controlling for each hospital's share of Medicaid admissions throughout our analyses. We additionally follow Dafny et al. (2019) and opt to control for each hospital's case mix index, acquired from the Medicare Inpatient Prospective Payment System's annual impact files, rather than include it in the denominator of the inpatient price calculation.

**Cost per adjusted discharge:** Using data from the CMS hospital cost reports, we follow the approach described in Schmitt (2017) to construct cost per adjusted discharge. The cost reports provide each hospital's annual costs broken out by cost center. Cost centers include, but are not limited to, inpatient, ancillary, and outpatient services. We compute total costs as the sum of salaries and other costs across all cost centers. We then derive the total cost per adjusted discharge by dividing total cost by adjusted discharges, the latter of which includes both inpatient discharges and an estimate of outpatient discharges based on the share of hospital charges attributable to outpatient services:

$$\text{cost} = \frac{\text{total cost}}{\text{total inpatient discharges} \times \left(1 + \frac{\text{outpatient charges}}{\text{inpatient charges}}\right)}.$$

**Hourly wages and employment:** Using data from the CMS hospital cost reports, we construct wages and employment following the approach described in Prager and Schmitt (2021). For employees grouped into narrow labor categories, the cost reports include the total amount paid and the hours worked for each hospital. We aggregate these worker types into three broad categories based on education levels, specificity of skills, and similarity of hourly wages: blue-collar workers, white-collar workers, and nursing and pharmacy workers. We then calculate the average hourly wage of each category $c$ by dividing total salaries paid by the total work hours:

$$\text{wage}_c = \frac{\text{total salaries}_c}{\text{total hours}_c}.$$

We measure employment as each hospital's number of full-time equivalent employees per labor category, which is calculated by dividing total hours by the number of hours in a cost report year:

$$\text{fte}_c = \frac{365}{\text{cost report days}} \times \frac{\text{total hours}_c}{52 \times 40}.$$

The mapping of narrow employment categories to broad ones is as follows:

- Blue collar: Maintenance & Repairs, Operation of Plant, Laundry & Linen Service, Housekeeping, Dietary, Cafeteria, Central Services & Supply, and Medical Records & Medical Records Library line items
- White collar: Employee Benefits Department, Administrative & General, Maintenance of Personnel, and Social Service line items
- Nursing and pharmacy: Nursing Administration and Pharmacy line items

Appendix Table OA1 tabulates the COPA mergers that contribute to estimation of each of the leads and lags for average treatment effects for each of the above outcomes. All COPA mergers contribute to all estimates with the exception of four years after merger, to which the most recent mergers Hendrick and Shannon do not contribute, and Palmetto, which does not contribute to estimation of leads for three or more years preceding treatment, as it takes place two years after the start of our sample period.

Table OA1: Tabulation of COPA Mergers by the Leads and Lags for ATEs they Contribute to Estimation of for Outcomes Other than 30-Day Mortality and Readmission

| Years from merger | Merger | | | | | | |
|---|---|---|---|---|---|---|---|
| | Palmetto | Maine Health | Phoebe Putney | Cabell Huntington | Ballad | Hendrick | Shannon |
| -6 | | ✗ | ✗ | ✗ | ✗ | ✗ | ✗ |
| -5 | | ✗ | ✗ | ✗ | ✗ | ✗ | ✗ |
| -4 | | ✗ | ✗ | ✗ | ✗ | ✗ | ✗ |
| -3 | | ✗ | ✗ | ✗ | ✗ | ✗ | ✗ |
| -2 | ✗ | ✗ | ✗ | ✗ | ✗ | ✗ | ✗ |
| -1 | ✗ | ✗ | ✗ | ✗ | ✗ | ✗ | ✗ |
| 0 | ✗ | ✗ | ✗ | ✗ | ✗ | ✗ | ✗ |
| 1 | ✗ | ✗ | ✗ | ✗ | ✗ | ✗ | ✗ |
| 2 | ✗ | ✗ | ✗ | ✗ | ✗ | ✗ | ✗ |
| 3 | ✗ | ✗ | ✗ | ✗ | ✗ | ✗ | ✗ |
| 4 | ✗ | ✗ | ✗ | ✗ | ✗ | | |

**Mortality and readmissions:** We derive quality measures from the CMS Hospital Compare Database. These data are available for the years 2009 through 2023. Mortality and readmission rates are constructed from Medicare Fee-for-Service claims and are provided for three underlying conditions: heart failure, heart attack, and pneumonia. As described in Dunn, Knepper, and Dauda (2021), Hospital Compare measures are typically lagged one year and therefore correspond to the years 2008 through 2022 in our sample. They are also constructed over three-year rolling periods, and the effects of mergers on these measures may be delayed. We construct aggregate measures of mortality and readmission by taking an admission-weighted average over the rates for each of these three conditions.

Appendix Table OA2 provides an analogous tabulation for 30-day mortality and readmission as was provided for all other outcomes in Appendix Table OA1. Leads and lags corresponding to years prior to 2008 will not contribute to estimation for these outcomes.

Table OA2: Tabulation of COPA Mergers by the Leads and Lags for ATEs they Contribute to Estimation of for 30-Day Mortality and Readmission

| Years from merger | Merger | | | | | | |
|---|---|---|---|---|---|---|---|
| | Palmetto | Maine Health | Phoebe Putney | Cabell Huntington | Ballad | Hendrick | Shannon |
| -6 | | | | ✗ | ✗ | ✗ | ✗ |
| -5 | | | | ✗ | ✗ | ✗ | ✗ |
| -4 | | | | ✗ | ✗ | ✗ | ✗ |
| -3 | | | ✗ | ✗ | ✗ | ✗ | ✗ |
| -2 | | ✗ | ✗ | ✗ | ✗ | ✗ | ✗ |
| -1 | | ✗ | ✗ | ✗ | ✗ | ✗ | ✗ |
| 0 | | ✗ | ✗ | ✗ | ✗ | ✗ | ✗ |
| 1 | | ✗ | ✗ | ✗ | ✗ | ✗ | ✗ |
| 2 | | ✗ | ✗ | ✗ | ✗ | ✗ | ✗ |
| 3 | | ✗ | ✗ | ✗ | ✗ | ✗ | ✗ |
| 4 | | ✗ | ✗ | ✗ | ✗ | | |

**Matching:** We match COPA hospitals to non-COPA ones using the Coarsened Exact Matching (CEM) algorithm to compute weights that generate balance across the groups in the change in and resultant HHI induced by the merger Iacus et al. (2012). The CEM algorithm coarsens the values of ΔHHI and $HHI$ into strata $s$. COPA (non-COPA) hospitals receive zero weight if there are no non-COPA (COPA) hospitals in their stratum. Letting $s$ denote the stratum to which an observation belongs, and assuming $s$ is restricted to stratum for which at least one COPA and non-COPA hospital are observed, weights for all matched units are computed as

$$\omega_h^{COPA} = \frac{1}{T_1 \times N^{COPA}}$$
$$\omega_h^{nonCOPA} = \left(\frac{N^{nonCOPA}}{N^{COPA}}\right)\left(\frac{N_{s(h)}^{COPA}}{N_{s(h)}^{nonCOPA}}\right)\left(\frac{1}{T_1 \times N^{nonCOPA}}\right),$$

where each weight's denominator includes the number of post-merger periods and unique hospitals of the respective merger type.

Sturge's rule is used to construct strata for HHI (Sturges, 1926). For ΔHHI, we use the

largest number of equally-sized strata for which no COPA hospitals go unmatched in the main results. This implies three ΔHHI strata.

# Online Appendix C Bias Correction for Correlation between ATEs Estimand

Our effect estimate can be decomposed into the true effect plus an unobserved shock:

$$\begin{aligned}\hat{\tau}_{ht}^{Y} =& Y_{ht}(1) - \hat{\alpha}_{h}^{Y} - \hat{\beta}_{t}^{Y} - X_{ht}'\hat{\gamma}^{Y} \\ =& Y_{ht}(1) - Y_{ht}(0) + Y_{ht}(0) - \hat{\alpha}_{h}^{Y} - \hat{\beta}_{t}^{Y} - X_{ht}'\hat{\gamma}^{Y} \\ =& \tau_{ht}^{Y} + Y_{ht}(0) - \hat{\alpha}_{h}^{Y} - \hat{\beta}_{t}^{Y} - X_{ht}'\hat{\gamma}^{Y} \equiv \tau_{ht}^{Y} + \varepsilon_{ht}^{Y}\end{aligned} \tag{4}$$

Assuming $E[\varepsilon_{ht}^{Y} \mid X] = 0$, our estimator is unbiased for linear combinations of individual effects:

$$E\left[\sum_{h,t} w_{ht}\hat{\tau}_{ht}^{Y} \mid X\right] = \sum_{h,t} w_{ht} E[\tau_{ht}^{Y} \mid X]$$

However, for the correlation estimand, bias arises from the product of residuals:

$$\frac{1}{NT}\sum_{ht}\hat{\tau}_{ht}^{Y}\hat{\tau}_{ht}^{Y'} = \frac{1}{NT}\sum_{ht}\tau_{ht}^{Y}\tau_{ht}^{Y'} + \frac{1}{NT}\sum_{ht}\varepsilon_{ht}^{Y}\varepsilon_{ht}^{Y'} + \frac{1}{NT}\sum_{ht}\tau_{ht}^{Y}\varepsilon_{ht}^{Y'} + \frac{1}{NT}\sum_{ht}\varepsilon_{ht}^{Y}\tau_{ht}^{Y'} \tag{5}$$

**Addressing bias via leave-one-out estimation:** To eliminate the bias in our estimate of the correlation between treatment effects, we use a leave-one-out (jackknife) approach. For each hospital $h$, we estimate $\hat{\beta}_{t}^{-h}$ and $\hat{\gamma}^{-h}$ using all hospitals except $h$. We then estimate $\hat{\alpha}_{h}$ using only hospital $h$'s data with these parameters. Following the same steps as Equation (4), we see that the leave-one-out treatment effect estimate is:

$$\hat{\tau}_{ht}^{Y,-h} = Y_{ht} - \hat{\alpha}_{h}^{Y} - \hat{\beta}_{t}^{Y,-h} - X_{ht}'\hat{\gamma}^{Y,-h} = \tau_{ht}^{Y} + \varepsilon_{ht}^{Y,-h}$$

where $\varepsilon_{ht}^{Y,-h} = Y_{ht}(0) - \hat{\alpha}_{h}^{Y} - \hat{\beta}_{t}^{Y,-h} - X_{ht}'\hat{\gamma}^{Y,-h}$.

We can construct an analog of Equation (5) for $\hat{\tau}_{ht}^{Y,-h}$. The only parameter estimated using hospital $h$'s data is $\hat{\alpha}_h$, which is computed as the mean residual across hospital $h$'s $T$ observations. The estimation error $\hat{\alpha}_h - \alpha_h$ is $O_p(T^{-1/2})$. This introduces two sources of $O(T^{-1})$ bias in the correlation. First, since both $\tau_{ht}^Y$ and $\hat{\alpha}_h$ depend on hospital $h$'s outcomes, their covariance is proportional to the variance of $\hat{\alpha}_h$, yielding $E[\tau_{ht}^{Y'}\varepsilon_{ht}^{Y,-h}] = O(T^{-1})$. To see this, note that the residual can be decomposed as:

$$\begin{aligned}
\varepsilon_{ht}^{Y,-h} &= Y_{ht}(0) - \hat{\alpha}_h^Y - \hat{\beta}_t^{Y,-h} - X_{ht}'\hat{\gamma}^{Y,-h} \\
&= [Y_{ht}(0) - \alpha_h^Y - \beta_t^Y - X_{ht}'\gamma^Y] + [\alpha_h^Y - \hat{\alpha}_h^Y] + [\beta_t^Y - \beta_t^{Y,-h}] + [X_{ht}'\gamma^Y - X_{ht}'\hat{\gamma}^{Y,-h}] \\
&= [Y_{ht}(0) - \alpha_h^Y - \beta_t^Y - X_{ht}'\gamma^Y] + [\alpha_h^Y - \hat{\alpha}_h^Y] + o_p(1)
\end{aligned}$$

where the last line follows as $N \to \infty$.

The first bracketed component is the structural error with mean zero by assumption, so:

$$\begin{aligned}
E[\tau_{ht}^Y \varepsilon_{ht}^{Y',-h}] &= E[\tau_{ht}^Y(\alpha_h^{Y'} - \hat{\alpha}_h^{Y'})] + o(1) \\
&= \mathrm{Cov}(\tau_{ht}^Y, \hat{\alpha}_h^{Y'}) + E[\tau_{ht}^Y]E[\alpha_h^{Y'} - \hat{\alpha}_h^{Y'}] \\
&= \mathrm{Cov}(\tau_{ht}^Y, \hat{\alpha}_h^{Y'}) \\
&= \mathrm{Cov}\left(\tau_{ht}^Y, \frac{1}{T}\sum_{s=1}^{T} Y_{hs}'(0)\right) \\
&= \frac{1}{T}\sum_{s=1}^{T} \mathrm{Cov}\big(\tau_{ht}^Y, Y_{hs}'(0)\big) \\
&\approx \frac{1}{T}\mathrm{Cov}\big(\tau_{ht}^Y, Y_{ht}'(0)\big) = O(T^{-1})
\end{aligned}$$

where the last line follows from the assumption that $\mathrm{Cov}(\tau_{ht}^Y, Y_{hs}')$ are small for $s \neq t$; i.e. we assume weak temporal correlation within hospitals. Both $\tau_{ht}^Y$ and $\hat{\alpha}_h^{Y'}$ are functions of hospital $h$'s potential outcomes. Similar logic and the same assumption of limited temporal correlation within hospitals yields $E[\varepsilon_{ht}^{Y,-h}\varepsilon_{ht}^{Y',-h}] = O(T^{-1})$.

With the leave-one-out estimator and $N \to \infty$ and $T$ fixed, the leave-one-out correlation

estimator satisfies:

$$E\left[\frac{1}{NT}\sum_{h,t}\hat{\tau}_{ht}^{Y,-h}\hat{\tau}_{ht}^{Y',-h}\right] = E\left[\frac{1}{NT}\sum_{h,t}\tau_{ht}^{Y}\tau_{ht}^{Y'}\right] + O(T^{-1})$$

If both $N, T \to \infty$, the estimator is consistent:

$$\frac{1}{NT}\sum_{h,t}\hat{\tau}_{ht}^{Y,-h}\hat{\tau}_{ht}^{Y',-h} \xrightarrow{p} \frac{1}{NT}\sum_{h,t}\tau_{ht}^{Y}\tau_{ht}^{Y'}$$

# Online Appendix D State-Issued Consent Decree Details and Results

Like its federal counterpart, a state-level consent decree is a negotiated settlement between the state attorney general and the merging parties following a challenge to the merger based on antitrust law. Unlike federal agreements, state-issued consent decrees are much more likely to impose behavioral rather than structural remedies.

Using Montague et al. (2023)'s survey on state-regulated healthcare provider transactions, we identify five hospital mergers consummated under state-issued consent decree that include some kind of remedy for increased product market power. We note that the set of mergers subject to state-imposed consent decrees collected by Montague et al. (2023) is limited to those issued between 2012 and 2022 for which publicly available information on the website of some healthcare market oversight entity exists. We are aware of no other catalog of mergers subject to state-imposed consent decrees, nor any other way of cataloging such data. We briefly summarize these mergers below, noting that only one regulates prices via caps on growth, all others requiring only "good faith negotiations" between the merging parties and payers.

- **Geisinger/Bloomsburg, 2011:** In 2012, the Commonwealth of Pennsylvania challenged Geisinger Health System Foundation's acquisition of Bloomsburg Hospital, alleging it

would substantially lessen competition in Columbia County. Under the consent decree, Geisinger agreed to operate Bloomsburg Hospital as an acute care hospital for eight years and maintain separate corporate status. The agreement required G-BH to enter into good faith negotiations with health plans for new contracts and established a third-party review process to validate negotiation methodologies and ensure fair pricing (U.S. District Court for the Middle District of Pennsylvania, 2012).

- **Faxton/SEMC, 2013:** In 2013, the New York AG approved a hospital merger between Faxton-St. Luke's Healthcare and St. Elizabeth Medical Center (one secular and one Catholic, respectively), arguing that the union was necessary for the hospitals' survival (Urban Institute, 2019; New York Attorney General Eric Schneiderman, 2013). The agreement requires the hospitals to negotiate in good faith with payers and grants payers the right to continue their currently-existing contracts with the hospital for 5 years at current prices if they believe the hospitals are negotiating unfairly (Montague et al., 2023).

- **Geisinger/Lewistown, 2014:** In 2013, the Commonwealth of Pennsylvania challenged Geisinger Health System Foundation's acquisition of Lewistown Health Care Foundation, alleging it would substantially lessen competition in Mifflin and Juniata Counties. Under the consent decree, Geisinger agreed to maintain Lewistown Hospital as a 123-bed acute care hospital for eight years. The agreement required G-LH and GC to enter into good faith negotiations with health plans for new contracts and established a third-party review process with independent consultant validation of negotiation methodologies (U.S. District Court for the Middle District of Pennsylvania, 2013).

- **Massachusetts healthcare providers, 2018:** In 2018, the Massachusetts Attorney General approved the merger of Lahey Health System and CareGroup (which included Beth Israel Deaconess Medical Center) to form Beth Israel Lahey Health. The merger brought together multiple hospitals and health systems under the Beth Israel Lahey Health umbrella, including Beth Israel Deaconess Medical Center, Lahey Hospital & Medical Cen-

ter, Mount Auburn Hospital, New England Baptist Hospital, and several other facilities. The agreement capped commercial unit price increases at the Health Care Cost Growth Benchmark minus 0.1% and froze managed Medicare rates (Massachusetts Office of the Attorney General, 2018).

- **Concord/LRG, 2021:** In 2021, the New Hampshire Attorney General's Consumer Protection and Antitrust Bureau filed a final judgment regarding Concord Hospital's acquisition of bankrupt LRGHealthcare, which operated Lakes Region General Hospital in Laconia and Franklin Regional Hospital. The agreement created a binding arbitration process that is available to payers if they believe the hospitals to be negotiating unfairly (Montague et al., 2023; New Hampshire Consumer Protection and Antitrust Bureau, 2021).

We incorporate consent decree mergers into our analysis by adding them to the treatment group that in our main analysis contains only COPA mergers. We then match mergers free from any state-imposed behavioral remedies (what we term "non-COPA" in the main text) to this treatment group of COPA and consent decree mergers using the same conditional exact matching algorithm as in the main results. Appendix Table OA3 provides summary statistics for consent decree hospitals and the weighted non-conditional agreement comparison group; summary statistics for the COPA treatment group and control group are unchanged relative to those presented in Table 1. Appendix Table OA4 provides average treatment effects and differences in ATEs for a specification of our event study model with three treatment groups.

Table OA3: Summary Statistics for Consent Decree Hospitals and Matched Non-Conditional Agreement Hospitals

| | Consent decree | Non-conditional agreement |
|---|---|---|
| Number of hospitals | 19 | 166 |
| *Outcome variables* | | |
| Inpatient price (thous. $) | 8.06 (6.57) | 9.12 (7.14) |
| Cost per adjusted discharge (thous. $) | 10.96 (5.89) | 9.67 (5.80) |
| 30-day mortality rate | 12.29 (1.90) | 12.82 (2.12) |
| 30-day readmission rate | 19.40 (1.93) | 19.26 (3.30) |
| Hourly wage | 31.16 (6.36) | 19.96 (7.88) |
| Employment | 346.75 (208.97) | 154.84 (172.44) |
| *Hospital characteristics* | | |
| % Medicaid | 7.69 (5.10) | 13.40 (10.90) |
| % Outpatient charges | 63.29 (14.54) | 49.15 (20.57) |
| Case-mix index | 1.61 (0.30) | 1.40 (0.29) |
| Non-profit | 0.95 (0.23) | 0.72 (0.45) |
| Number of beds | 175.47 (125.92) | 135.39 (131.55) |
| *County characteristics* | | |
| Population (thous.) | 599.19 (529.30) | 301.11 (425.70) |
| Share age 65 or older | 16.88 (2.40) | 14.49 (2.93) |
| Per capita income (thous. $) | 62.64 (19.29) | 34.20 (12.54) |
| One-bedroom rent | 1025.66 (370.75) | 561.28 (190.08) |
| Share rural | 0.05 (0.14) | 0.11 (0.20) |
| *Other characteristics* | | |
| System size | 3.89 (2.85) | 3.37 (3.72) |
| HHI | 1334.75 (1084.82) | 2647.32 (1582.12) |
| $\Delta$HHI | 274.21 (471.16) | 205.84 (606.17) |

*Notes:* Values are for the year in which the hospital participated in its merger, or the first year in which it appears in the data. Cells show mean (standard deviation) for continuous measures and percentages for binary and categorical variables. HHI stands for Herfindahl–Hirschman index and is computed using bed shares for hospitals within 25 miles of one another. Non-conditional agreement observations are weighted using weights from coarsened exact matching of non-conditional agreement hospitals to COPA and consent decree hospitals on HHI and $\Delta$HHI.

Table OA4: Average Treatment Effects Including Consent Decree Treatment Group

| | COPA | | Consent decree | | Non-conditional agreement | | Difference in ATEs | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| | | | | | | | (1) minus (3) | | (2) minus (3) | |
| | (1) | | (2) | | (3) | | (4) | | (5) | |
| Inpatient price | $-0.073^{***}$ | (0.023) | $-0.025$ | (0.031) | $0.043^{***}$ | (0.016) | $-0.116^{***}$ | (0.027) | $-0.068^{**}$ | (0.031) |
| Cost per adjusted discharge | $-0.025$ | (0.017) | $-0.113^{***}$ | (0.021) | $-0.011$ | (0.009) | $-0.014$ | (0.018) | $-0.102^{***}$ | (0.018) |
| 30-day mortality rate | $0.007^{***}$ | (0.001) | $-0.004^{**}$ | (0.002) | $0.002^{*}$ | (0.001) | $0.005^{***}$ | (0.002) | $-0.006^{***}$ | (0.002) |
| 30-day readmission rate | $0.009^{***}$ | (0.003) | 0.001 | (0.002) | $-0.006^{**}$ | (0.002) | $0.015^{***}$ | (0.003) | $0.006^{**}$ | (0.003) |
| Wages | $0.080^{***}$ | (0.029) | $-0.033^{**}$ | (0.015) | 0.009 | (0.008) | $0.071^{**}$ | (0.030) | $-0.042^{***}$ | (0.013) |
| Employment | $-0.263^{***}$ | (0.035) | $-0.019$ | (0.027) | $-0.084^{***}$ | (0.019) | $-0.179^{***}$ | (0.039) | $0.065^{**}$ | (0.025) |

*Notes:* Table shows average treatment effects (ATEs) taken over the four years following a merger. Column (1) reports ATEs for hospitals undergoing COPA mergers. Column (2) reports ATEs for hospitals undergoing consent decree mergers. Column (3) reports ATEs for non-conditional agreement mergers. Column (4) provides the difference between Columns (1) and (3). Column (5) provides the difference between Columns (2) and (3). Averages are weighted using weights from coarsened exact matching of non-conditional agreement hospitals to COPA and consent decree hospitals on HHI and $\Delta$HHI. Standard errors clustered at the system level are shown in parentheses. $***p < 0.01, **p < 0.05, *p < 0.10$.