THE IMPACT OF PRIVATIZATION: EVIDENCE FROM THE HOSPITAL SECTOR*

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Abstract

Privatization has been shown to increase growth and profitability of public firms. However, effects on consumers are understudied. We study potential trade-offs in the US hospital sector where public control declined by 42% over 1983–2019. Private operators may improve hospitals' financial performance, but a focus on profitability may adversely affect access to care for certain patients. Using national data across all hospitals and patients, we study 258 hospital privatizations over the 2000–2018 period. Private operators improve profitability so that hospitals generate a modest surplus, primarily by increasing mean revenue per patient. However, this is partly achieved by differentially reducing the intake of low-income Medicaid patients, who are typically less profitable than other groups due to lower reimbursement rates. While other patients appear to be absorbed by neighboring hospitals, Medicaid patients experience an aggregate decline in utilization at the market-level, which we interpret as a decline in access to care. Hospital privatization therefore partially offsets the benefits of providing publicly funded health insurance through Medicaid, and our estimates imply it is quantitatively important. The aggregate decline in Medicaid volume is detected only in more concentrated hospital markets, suggesting market power is a key driver.

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1 Introduction

When should governments rather than private firms provide goods and services? Economists have long been interested in this question without reaching a consensus (Shleifer, 1998). Meanwhile, privatization is an important global phenomenon, with nearly a trillion dollars raised over 2013–16 through the sale of government assets (Megginson, 2017). The balance of empirical evidence suggests that privatization improves the efficiency and growth of government-owned firms (Ehrlich et al., 1994; World Bank, 1995). However, effects on consumers are understudied (Megginson and Netter, 2001; Galiani, Gertler and Schargrodsky, 2005). This is a key limitation since the privatization debate now centers around the delivery of social services, which has traditionally been the domain of governments (Stiglitz, 2005). Several countries, including Germany and Sweden, have privatized or are contemplating the privatization of healthcare providers; however, it remains contentious (Bergman et al., 2016; Knutsson and Tyrefors, 2022).[1] This paper begins to fill this gap by studying the costs and benefits of privatizing hospitals in the US.

Economic theory has long recognized the potential benefits of privatization. The performance of public firms may suffer due to misaligned incentives of employees and managers, soft budget constraints, and political interference (Shleifer and Vishny, 1994; Sheshinski and López-Calva, 2003). Private management can resolve these agency problems and improve profitability and growth, also potentially benefiting consumers. This works particularly well in industries with sufficient competition and without other market failures (Vickers and Yarrow, 1991). However, in the presence of market imperfections, public firms may increase consumer welfare by setting prices or quantities that account for social marginal benefits (La Porta and López-de Silanes, 1999). These concerns are exacerbated in the case of hospitals where a combination of profit maximization and clinical discretion may lead to a reduction in access for unprofitable or less profitable patients. Hart, Shleifer and Vishny (1997) hypothesize these responses by private contractors as a form of shirking on non-contractible quality. Indeed, government hospitals have been historically justified as "safety-net" facilities for such patients. Further, local hospital markets are highly concentrated, implying that private operators will not feel pressured to improve quality of care and may leverage market power to

[1] Several developed countries have privatized or attempted to privatize healthcare delivery. Germany has privatized more than a hundred hospitals since the 1990s (Heimeshoff, Schreyögg and Tiemann, 2014). England introduced legislation in 2012 to encourage private delivery of healthcare (Goodair and Reeves, 2022). Sweden implemented a multi-phase, multi-sector privatization over 1990–2013 with mixed effects (Dahlgren, 2014). Madrid, Spain attempted to privatize hospitals, but the proposal had to be scrapped due to protests by unions. The provincial governments of Quebec and Ontario in Canada are reportedly considering privatizing hospitals. The city of Detroit privatized its public health department in 2012 but reversed course following the Covid-19 pandemic.

generate greater profits (Bloom et al., 2015; Fulton, 2017).[2]

The hospital sector is an important and suitable empirical setting to test these theoretical arguments. It involves more than a trillion dollars in annual spending and employs more than 6.5 million people in the US, comparable in size to the entire Construction sector.[3] Because of the safety net considerations discussed above, the hospital sector employs more government employees in the US than any other sector except education.[4] Government hospitals contributed over 20% of bed capacity and employment in 2019 even though their presence has steadily diminished over the last few decades (see Figure 1a). The number of public hospitals has declined through two broad channels: conversion to private managerial control and closure.[5] Privatization is the dominant mechanism: there are more than 6 privatizations for every public hospital closure during our sample period.

This paper examines the effects of nearly 260 privatizations of non-federal public hospitals that occurred between 2000 and 2018. Our main data source is annual hospital surveys administered by the American Hospital Association (AHA), which allows us to observe hospital attributes and three key outcomes of interest: finances, patient volume across payers, and employment. For one key insurer, Medicare fee-for-service, we also use confidential, patient-level claims data.[6] A key virtue of our data is its comprehensive national coverage, allowing us to assess effects for the entire hospital sector, the key objective in this paper. We complement these sources with publicly available files from the Medicare cost reports and the US Census.

We employ a staggered difference-in-difference research design to estimate the effects of privatizations on the treated hospital, as well as spillovers on the market where the hospital is located. This follows the approach used by recent studies examining privatization (Galiani, Gertler and Schargrodsky, 2005; Arnold, forthcoming) as well as the organization of healthcare markets (Eliason et al. 2020, Craig, Grennan and Swanson 2021). We compare outcomes at the treated hospitals (markets) following the privatization with other public hospitals (markets) that did not experience a change in ownership during our sample period.

While our research design is standard in this literature, we recognize that privatizations are not exoge-

[2]In our sample, the mean Herfindahl-Hirschmann Index (HHI) across hospital markets was about 5,600 in 1999, well over the US Department of Justice threshold of 2,500 to determine a highly concentrated market (DOJ, 2010).

[3]Source: 2019 Quarterly Census of Employment and Wages – The Bureau of Labor Services (BLS). The Construction sector, NAICS code 23, employed about 7 million individuals in 2019.

[4]Source: Current Employment Statistics – The Bureau of Labor Services (BLS).

[5]When a hospital stops providing inpatient care, we consider that to be a closure even if it continues to provide outpatient care such as maintaining an Emergency Department or physician clinics. These latter cases are sometimes called partial closures.

[6]This is also sometimes known as Traditional Medicare, and is the largest payer of hospital services in the US. We are not able to repeat this exercise for Medicaid patients as there exists no equivalent national claims data since Medicaid is administered at the state level unlike federally administered Medicare.

nously assigned. We therefore take a number of precautions to probe the validity of our estimation strategy. We examine dynamic effects around the year of the privatization and find that the privatized hospitals do not differ from the comparison group prior to the change, but experience an immediate and persistent shift following the transition. We also subject the estimates to a number of robustness checks, including controlling for indicators of economic activity, using a matching design, and correcting for potential bias due to the staggered design. The estimates are qualitatively similar in all cases.

A key purpose to privatize hospitals is to improve their profitability so they become financially sustainable without government subsidy. The treated hospitals were unprofitable prior to privatization, with an average operating margin of -4% (of revenue). Their profitability suffered in comparison to private hospitals primarily due to lower mean revenue per patient even though they had lower personnel and total operating costs per patient before the transition. We find that private owners improve performance exactly on this dimension – mean revenue per patient increases by about 6%, alone sufficient to make a modest surplus. We also detect a substantial reduction in personnel spending, driven entirely by a reduction in employees per patient, suggesting greater operating efficiency. However, total operating cost per patient reduces by a statistically insignificant 3%. Overall, privatization improves profitability substantially.

Private control does not increase patient volume along with profitability. We find that patient volume decreases by 8.4% and it does not recover to pre-privatization levels even after five years.[7] Medicaid patients, who tend to be less profitable than other payers on average and reportedly pay less than the average cost of care, experience a disproportionate decline (Schulman and Milstein, 2019).[8] They experience nearly a 15% reduction in volume, implying that they account for 30% of the decline in patients even though they only form 20% of the patient base. Hence, private firms improve profitability partly by reducing the share of Medicaid patients at the hospital.

The strategy of changing patient mix at the privatized hospital could have larger implications for beneficiaries of Medicaid, the means-tested health insurance program that now covers one in four Americans and is the single largest insurer by number of enrollees.[9] To investigate the impact on access, we quantify the effect on aggregate market-level Medicaid hospital volume, assuming that a market is treated when a public

[7]We do not find a simultaneous increase in outpatient or Emergency Department (ED) volume to suggest a change in treatment style. The coefficient on outpatient volume is noisy, but has a negative sign, also implying a decline.

[8]National data on mean reimbursement rates by insurer does not exist, but some states report these values. Using data reported by California over 2008–10, we find that Medicaid mean inpatient reimbursement rates were 84% and 60% as large as mean Medicare and private insurance rates, respectively.

[9]The federal government's monthly enrollment reports indicate that Medicaid and Medicare had 84 million and 59 million enrollees, respectively, as of August 2022. US population was estimated at 332 million in 2022 by the Census Bureau.

hospital is first privatized.[10] We estimate a 4% decline in Medicaid utilization at the market-level, about the magnitude we would predict due to the decline at the privatized hospital without any compensating behavior by other facilities in the market.[11] Medicaid is the only payer that experiences an aggregate decline in volume. We interpret this decline as a reduction in access to hospital care for Medicaid patients, consistent with the claim that government hospitals are an essential component of the safety net.

The decline in Medicaid utilization could be socially valuable if the avoided stays were wasteful. We cannot rule out this possibility, but the evidence suggests this is not the main driver. We find substantial heterogeneity in the decline in Medicaid volume across different market types. Specifically, we detect a 12% decline in aggregate Medicaid volume in markets with above-median concentration and null effects in the less concentrated markets. The association with higher concentration level suggests that market power is a key mechanism. We also detect a large aggregate decline in Medicaid in markets with higher poverty rates. Hospitals in these markets have lower profit margins on average, and this result is consistent with the possibility that hospitals under greater financial stress implement deeper Medicaid cuts. In both these market types, the magnitude of the aggregate Medicaid decline is only partially explained by the direct effect on the privatized hospital, implying that competing hospitals also respond by cutting Medicaid when they are under financial stress or can exercise market power.

This paper makes three contributions. To our knowledge, we are the first to obtain nationally representative estimates of the causal effects of hospital privatization in the US, adding to the privatization literature in economics.[12] Chan, Card and Taylor (2022) study differences in quality of care between publicly owned Veterans Affairs (VA) hospitals and non-federal community hospitals using a quasi-experimental design. However, it is unclear how informative these results are about the privatization of the average non-federal government-owned hospital. Even outside of economics, we are aware of only a few relevant studies. Ramamonjiarivelo et al. (2016, 2020) study privatizations over an earlier period and document improved hospital profitability. However, they do not consider the market-level effects on access to care for Medicaid patients, nor the effects on employment. Our finding of improved profitability supports arguments favoring privati-

[10] We define hospital markets using Health Service Areas (HSAs). These are collections of contiguous counties that were delineated by the US Census to define self-contained hospital markets. There are approximately 900 HSAs in the US.

[11] Privatized hospitals account for about 24% of total market capacity on average. Hence, a 15% decline in Medicaid volume at the privatized hospital alone would predict a 3.6% decline in market-level volume if there was no counteracting response by other hospitals.

[12] Although there is extensive work on the deregulation in sectors such as airlines, telecommunications, and electricity, the evidence on privatization across sectors in the US is thin (Morrison and Winston, 2010; Davis and Wolfram, 2012; Levin and Tadelis, 2010; Borenstein and Bushnell, 2015; Howell et al., 2022).

zation, similar to much of the prior literature in this space, but the reduction in aggregate Medicaid volume validates the social argument against privatization (Boycko, Shleifer and Vishny, 1997; López-de Silanes, Shleifer and Vishny, 1997; La Porta and López-de Silanes, 1999; Megginson and Netter, 2001; Savas, 2005). We highlight this tradeoff in access versus profitability using data on all non-federal hospitals and across all patients. We document heterogeneity in the effects of privatizations across markets and payers which can guide further investigations. Future studies can build on this by focusing on a specific state or group of patients with more granular data to examine, for example, quality of care.

Second, we shed light on a novel link between provider operations and public health insurance. Prior studies in this literature have typically examined the effects of expanding or improving public insurance programs on downstream providers. For example, Garthwaite, Gross and Notowidigdo (2018) quantify the cost of uninsurance for hospitals in the form of uncompensated care and show that it is borne disproportionately by non-profit hospitals. Dunn et al. (2021) show that administrative frictions in Medicaid have substantial costs for physicians. Our results indicate the presence of a reverse causal link: reducing government control in the downstream hospital market limits the effectiveness of Medicaid. The effects are particularly pernicious in more concentrated markets.

Finally, our paper highlights the interplay between two uncoordinated policy choices: hospital privatization and Medicaid expansion. Federal and state governments have substantially expanded Medicaid eligibility over our sample period, which has been well studied (Cutler and Gruber 1996; Currie and Gruber 1996a; Garthwaite, Gross and Notowidigdo 2014; Miller, Johnson and Wherry 2019; among many others). However, the shrinking government provision of hospital care has received little attention. National statistics on enrollment and spending on Medicare and Medicaid show much lower growth in spending for the latter, implying some factors differentially suppress spending by Medicaid beneficiaries. Similarly, the observed Medicaid hospital volume growth (18% over 2000-19) is substantially lower than the 38% we would predict based on quasi-experimental estimates of Medicaid coverage expansion from prior studies, a 20 percentage point gap.[13] Our results imply that hospital privatization is one of the factors suppressing Medicaid volume, and can explain about 20% (4%/20%) of this gap in the markets that experienced privatizations.

The paper proceeds as follows. Section 2 provides the necessary background about public hospital ownership and privatization. We follow with a description of the data in Section 3, and our empirical

[13]Medicaid enrollment grew 38% between 2000-2019, from 15.7% to 21.7%, per data from the Kaiser Family Foundation. Applying the elasticity estimate of 100% reported by Duggan, Gupta and Jackson (2022) which reflects general equilibrium effects, we predict an increase of 38% in hospital stays as well.

strategy in Section 4. We present the estimated effects on the privatized hospitals in Section 5. We similarly examine effects on the affected markets in Section 6. Finally, Section 7 concludes.

2 Background

2.1 Hospital ownership

There is substantial heterogeneity in the ownership mix of hospitals across different geographies.[14] This is true not only of the share of publicly owned hospitals in a market, but also the type of privately owned hospital (non-profit or for-profit). Table 1 highlights this variation and presents the shares of bed capacity of four different owner types (public non-federal, public federal, private non-profit, and private for-profit) for a selected set of six large states with at least 100 hospitals in 2019 (AL, CA, TX, GA, IL, and PA). We also present the corresponding national means and standard deviations in column 7. The columns are ordered in descending order of non-federal public share of hospitals. For completeness, Appendix Table A.1 presents the corresponding values of non-federal share of bed capacity for all states. In these tables and throughout the paper we choose to focus on non-federal public hospitals since these usually serve the local community and are more comparable to private hospitals than federal hospitals, which mostly cater to military veterans or other designated populations (eg., Native Americans).

We note two interesting patterns in hospital ownership. First, states vary tremendously in their reliance on public hospitals. Pennsylvania has only 4% of beds at such hospitals, while 44% of Alabama hospital beds are at state or local government hospitals. This variation is even greater if we consider small states (Wyoming and Vermont have 71% and 2%, respectively). Second, the share of public hospitals does not necessarily track states' preferences over the size of government or the rural-urban split. For example, Texas and Alabama have higher public hospital shares than do Illinois and Pennsylvania. Of the 10 states with the greatest public shares of hospital capacity, only one (Washington) was more liberal than the average in 2018, and three were among the most conservative (Alabama, Mississippi, and Louisiana). The shares of public provision of hospital care do not track the state's rural share of population either. New Hampshire, Maine, South Dakota, North Dakota, and Vermont are all among the most rural states in the US and have

[14] The AHA survey reports hospital *control*, which could be recorded as one of non-profit, for-profit, or government. Control and ownership are identical except in the small number of cases where the owner outsources managerial control or leases the property to a contractor who happens to have a different organization structure. Such contractors are invariably private firms; hence there are some cases, as we shall discuss below, where the government *owns* the hospital, but it is *controlled* by a private firm. Unless specified otherwise, our focus is on which entity has managerial control.

the lowest shares of public hospital capacity.[15] We interpret this heterogeneity as a sign that there is little consensus across states on the appropriate scope for public delivery of hospital services.

2.2 Government and hospital care

Figure 1 presents national trends related to government involvement in hospital care over 1983–2019, all sourced from the American Hospital Association annual survey data. Panel (a) shows that the share of hospital beds at all non-federal government owned hospitals declined from 27% in 1983 to 17% in 2019, a drop of nearly 40%. If we include in this calculation ownership by the federal government, the share declined from 36% to 21%, more than a 40% decrease. There is a parallel, though slightly smaller, decline in the share of hospital employees working at public hospitals. Overall, public hospitals have consistently declined in importance over this period, though the decline was steeper in the 1980s and 1990s.

This pattern of declining public provision of hospital care is in stark contrast to the expansion of public insurance coverage for hospital care over the same period. Figure 1 panel (b) plots the trend in the share of patients covered by the two major public insurance programs at non-federal hospitals. Medicaid, the means-tested public insurance program *doubled* its share of hospital patients from 10% in 1983 to 22% in 2019. This is not surprising since Medicaid coverage has been expanded through several federal and state policy initiatives over this period, extending eligibility to an ever increasing share of the population. The share of Medicare – the public insurance program for the elderly – also increased from 32% to 45%.[16] Unlike Medicaid, eligibility for this program has been relatively stable and a large part of the increase is likely due to the aging of the population. By the end of the sample period, the two public insurance plans collectively sponsored care for about two-thirds of all hospital patients, an increase of nearly 60% relative to 1983.

The dramatic decline in government provision of hospital care is not part of a wider trend of a diminishing government provision of other services considered important for social welfare. For example, government provision of education services – a sector with similar market failures and policy considerations as healthcare – have remained remarkably stable over this period. For example, the share of students at public high schools has actually increased slightly, while the share of students at public degree granting institutions has decreased by 4 percentage points (pp).[17]

[15] Sources: Gallup poll on political preferences, 2018. See `https://news.gallup.com/poll/247016/conservatives-greatly-outnumber-liberals-states.aspx`. State rural share of population: `https://www.icip.iastate.edu/tables/population/urban-pct-states.`

[16] For all our analyses, Medicare includes patients on Traditional Medicare (TM) and Medicare Advantage (MA).

[17] Source: National Center for Education Statistics data on high school and higher education student enrollment.

Historically, a key justification for public ownership of hospitals has been to ensure access to hospital care for certain vulnerable patient groups that may be shunned by private care providers because they lack insurance coverage or are otherwise unprofitable. Perhaps state and local governments viewed the expansion of Medicaid coverage as an alternate means to ensure access to care, making it easier to justify the privatization or divestiture of public hospitals. While the national time series on public hospital control and public insurance coverage of hospital patients are negatively correlated, we formally tested whether this negative correlation also exists at the state-level.

We estimate the association between state-level changes in Medicaid's share of non-federal hospital patients (ΔM_{st}) and the corresponding changes in the public, non-federal share of hospital bed capacity (ΔP_{st}) over four periods – 1983–91, 1992–2000, 2001–09, and 2010–18 – using the following model, stacking all four periods together:

$$\Delta P_{st} = \alpha_t + \gamma \Delta M_{st} + \xi_{st}. \tag{1}$$

γ is the coefficient of interest in this model and captures the within-state correlation between changes in Medicaid coverage and public hospital capacity. We weight each cell by the respective state population to account for the heterogeneity in size across states. We obtain a statistically significant estimate of -0.41 (0.11) for γ, implying that an increase in Medicaid share of 10 pp in a state is associated with a decline in the government's share of bed capacity in that state of about 4 pp. Recall that the national share of non-federal public hospitals dropped by about 10 pp over this period; hence this effect size is economically meaningful. We emphasize that this estimate represents a *correlation* and may not be causal. However, it is consistent with the hypothesis that local and state government officials may view the expanded eligibility for Medicaid as an acceptable substitute for public hospitals to ensure access to care.

2.3 Privatization benefits and costs

There is an extensive theoretical and empirical literature on the costs and benefits of government ownership of firms, as well as on the effects of privatizing government-owned enterprises (Vickers and Yarrow, 1991; Megginson and Netter, 2001). Shleifer (1998) summarizes the theoretical arguments that have been made for and against privatization. The main argument in favor is to alleviate agency problems with government employees and managers. Agency problems can arise through several channels, including political interference, soft budget constraints, and poor performance management practices. Another argument is that private firms can ease capital and credit constraints, thereby enabling faster growth (Ehrlich et al., 1994). By resolving these frictions, private owners can improve the growth and profitability of formerly government-owned firms. This rewards not just the firm's managers and new shareholders, but potentially consumers as well. On the other hand, government ownership may also improve consumer welfare since public firms are supposed to maximize social welfare and not just profits. Therefore they may choose inputs and set prices and quantities to maximize social objectives. Alternatively, we may worry that private firms may inappropriately cut back on dimensions that cannot be contractually enforced, in order to maximize profits (Hart, Shleifer and Vishny, 1997).

Greater efficiency due to privatization is certainly plausible in the case of hospitals. Local politicians frequently have direct oversight and control over public hospitals, introducing the possibility of political interference in day-to-day operations. Managers at public hospitals may have limited flexibility to change operations due to greater unionization. Appendix Figure A.1 presents the share of unionized employees at public and private hospitals over 1995–2019 using data from the Current Population Survey (CPS). Unionization is consistently about twice as likely in public hospitals. This suggests greater employee protections at public hospitals and possibly lower efficiency. Private owners may improve operating efficiency by reducing the number of hospital employees per patient served, a strategy often deployed following privatization (Arnold, forthcoming). This has the potential to save substantial costs since personnel spend accounts for more than half the total operating costs of hospitals, on average.

At the same time, there is the potential for harm to some consumers due to privatization. Public hospitals consistently have a much higher share of low-income Medicaid and uninsured patients than their private counterparts, including private non-profit hospitals (Horwitz, 2005; Horwitz and Nichols, 2022). That private hospitals will be reluctant to admit patients covered by these payers is intuitive since hospitals are expected (and legally obligated) to provide the same quality of service to all patients although these payers reimburse at lower rates on average than private insurers and Medicare do.[18] Hence, a key concern is whether private management will adopt a strategy to reduce the intake of less lucrative patients, whether due to their insurer type or the nature of care required (eg., psychiatric patients).

[18]There is no national dataset that reports mean reimbursements by payer, but California does publicly report net reimbursements by payer. In California, average inpatient reimbursement rates for Medicaid and uninsured/self-pay patients were about 85% and 44%, respectively, of the mean rate for Medicare patients over 2008–10. Their rates were even lower relative to privately insured patients.

2.4 Hospital privatization in the US

The number of hospitals in the US under government control has declined through two channels. First, local governments relinquished operational control of the hospital to a private firm, which we call privatization. This accounts for 85% of the decline in the number of government hospitals. Second, public hospitals either closed operations entirely or converted to solely providing outpatient care. We identified 258 and 41 cases of privatizations and closures over 2000–18, respectively. Of the 1,060 public hospitals present in our sample in 1999, nearly a quarter were privatized during this period.

We classified privatization deals based on their key features. We did not have access to the contracts between governments and private firms and relied on press releases and independent reporting from the period around the transition. Appendix Table A.2 presents the distribution of the different types of deals represented in our sample, and whether the new operator is organized as a for-profit or non-profit. As the table shows, privatization can manifest in several different forms and one could argue that every case has some unique features. We find hospitals were brought under both non-profit and for-profit control, with the latter accounting for 28% of deals.

The private firm's operational control over the hospital after the transition varies in a continuum across different types of deal structures, ranging from limited control (short-term concessions) to complete control (ownership of all hospital assets). Appendix Section B.1 provides details on the different ways in which governments transfer hospital control. To simplify exposition, we group deals into two categories representing less and more private control. Following Hart, Shleifer and Vishny (1997), we do not rely on ownership of the assets alone to define control, rather we focus on operational control.

The first group accounts for nearly 60% of all deals and represents less control for the private operator. The government retains ownership of all assets, but outsources operational and managerial control to a private contractor. This structure was preferred to outright sales in some states (eg., Florida) because certain hospital sales required legislative approval, a lengthy and uncertain process (Needleman et al., 1997). The most common deal structure in this group was for the government to find a hospital management firm that would operate the hospital in return for a fixed monthly fee. We refer to this as "contract management." In another common approach, the government transfers operational control to a private firm specially incorporated to run the hospital. The government agency continues to oversee the new entity. It is unclear how much incentive the private operators have to improve the hospital's profitability under these arrangements and whether they have the autonomy to focus on more profitable patients and services or to cut staff. In general, we did not find language in the press reports suggesting the operators had constraints on their ability to make such changes.

Private operators enjoy substantially more operational control over the hospital in the second group of deals. This group contains three types of deal structures. The first is an outright sale of all hospital assets to the contractor. We assume the new owners operate the hospital to maximize their own objectives, as they would any of their existing hospitals. The second approach is for the government to award a long-term lease (usually more than 15 years), giving the contractor more autonomy to make changes to the buildings and other assets as well as day-to-day operational control. A third, related approach that also involves a long-term transfer of control along with autonomy over the assets is for the contractor to enter into a joint venture with the government.

Interestingly, for-profit operators are significantly more likely to be involved in the second group of deals than in the first. They are involved in more than 40% of the deals bestowing more control but in less than 20% of the deals bestowing less control. Specifically, ownership deals involve a non-profit and for-profit buyer in nearly equal proportions. Overall, we anticipate greater impacts of privatization in the second group of deals since the private firm likely has greater incentive and ability to make substantial operational changes and increase surplus. We test this hypothesis in Section 5.

3 Data and descriptive evidence

3.1 Data sources and sample construction

Our main data source comprises annual surveys of hospitals from the American Hospital Association (AHA). Our primary analysis relies on AHA files for the years 1995–2019. We use the AHA files to source key outcomes variables, information on hospital attributes such as ownership (public or private), size, and location. We exclude two types of hospitals from our main analysis sample. First, as discussed previously, we exclude federal hospitals since they typically cater to a distinct set of patients (such as veterans or Native Americans) rather than the local community at large. Our sample includes hospitals owned by a state, county, city, or by a hospital district.[19] Second, we exclude specialized hospitals such as psychiatric and

[19]Hospital districts are funded by taxpayers to own and operate public hospitals. These are mostly found in rural markets. They are typically comparable to a county in terms of size.

rehabilitation facilities. In addition to being highly specialized, these hospitals are often reimbursed in a distinct way from community hospitals. Therefore, our final sample contains non-federal, general acute care hospitals.[20]

Since the treatment of interest is the privatization of publicly owned hospitals, we took several steps to minimize measurement error in identifying hospital ownership transitions. We inferred changes in owner type if the value reported on the AHA survey changed from one year to the next. This naive approach yielded a total of 354 privatizations of public hospitals over 2000–18. We manually validated these implied changes in ownership by examining the annual summary of change files from the AHA, news articles, press releases, and hospital websites; and confirming the changes against proprietary databases such as the American Hospital Directory (AHD), which tracks hospital ownership over time. If we were not able to confirm a privatization, we dropped the relevant hospital from the sample for our baseline model. In several cases, the external data also helped us correct the year of privatization. Using this approach, we validated 258 privatizations.

The final sample contains public hospitals that were either treated (privatized) or that did not experience a change in ownership.[21] Our final analysis sample is an unbalanced panel at the hospital-year level. Figure A.2 presents a frequency distribution of the number of years we observe hospitals in the sample. About 90% of the hospitals are observed for the maximum possible 25 years. The patterns are nearly identical for the privatized and comparison hospitals.

A key limitation of the AHA data on hospital utilization is the inability to observe patient-level changes. We overcome this limitation in the case of Medicare patients, where we have access to confidential, administrative claims data on fee-for-service or Traditional Medicare (TM) patients, obtained from the Centers for Medicaid and Medicare Services (CMS). Although the popularity of managed care has grown in recent years, TM patients account for more than 70% of all Medicare enrollment during our sample period. This sample covers the period 2000–2017, largely overlapping with the main AHA sample. We apply the same sample construction rules to this data and examine the effects on total inpatient utilization among TM

[20] We identify general acute care hospitals using AHA's primary service code of 10, which are "general medical and surgical" hospitals. We include all hospitals whose most common service code is general medical and surgical. The predominant service code among excluded public hospitals is psychiatric.

[21] We cannot rule out the possibility of false negatives – public hospitals that were privatized but this transition was not reported to the AHA. We believe this is very unlikely since it would mean the change is not reported over multiple years, not just once. We conducted random checks and did not find any. This measurement error, if it exists, will tend to bias effects toward zero. We also validated 46 transitions of privately owned hospitals to public ownership during this period. We excluded these hospitals from the sample.

patients.

The key outcome variables are measures of hospital finances (revenue and operating expenses), patient volume, and employment. With the exception of patient revenue, which is sourced from the Healthcare Cost Reporting Information System (HCRIS) or Medicare cost reports, all the outcome variables are sourced from the AHA surveys. We inflation adjust the revenue and expenses and express them in 2019 dollars. We study patient volume by payer and in aggregate. Specifically, we observe volume for three payers: Medicare, Medicaid, and a residual group ('Others') that is largely composed of privately insured and uninsured patients. A limitation of the AHA data is that we cannot separately observe volume for uninsured and privately insured patients.[22] We examine total full-time equivalent (FTE) employed staff and the effects on different components (physicians, nurses, and others).

To circumvent potential bias due to the substantial skew in outcomes across hospitals of different sizes, we transform the variables in the regression analyses. In the case of revenue, expenses, and patient volume, we use logs rather than the level. In the case of labor inputs, we normalize by adjusted admissions, which include outpatient visits.[23] We also test sensitivity to normalizing by the contemporaneous number of hospital beds instead of by adjusted admissions. The results are qualitatively similar.

We supplement the main data sources with information on market-level attributes, such as county-level population, poverty, unemployment, and uninsurance rates from publicly available data sources like the US Census and the Bureau of Labor Services (BLS).

3.2 Descriptive evidence

Table 2 describes the hospital-level analysis sample. Across all columns, we present values from 1999, a year prior to the first privatization in our sample. Column 1 presents values for the 258 hospitals that would be privatized (treated) during the sample period. Column 2 describes the 802 remaining public hospitals that did not experience a change in ownership during this period and are located at least 15 miles away from any privatized hospital. This group comprises our primary comparison group. We imposed this distance requirement to mitigate the potential for spillover contamination.[24] Comparing values in these two columns

[22]To our knowledge, there is no national data that can do better. It is possible to use uncompensated care costs reported in the Medicare cost reports to *impute* the share of uninsured patients, but this approach is feasible only after 2010.

[23]Adjusted admissions are preferred over using inpatient volume alone since they also account for outpatient care which has rapidly grown over time. The AHA reports adjusted admissions and we use them directly. These are typically computed by scaling outpatient volume by the ratio of outpatient charges to inpatient charges (Schmitt, 2017).

[24]This restriction drops only 32 potential control hospitals. The choice of 15 miles is somewhat arbitrary and trades off the need to isolate comparison hospitals from treated facilities against the desire to retain a larger share of potential comparison hospitals in

reveals that privatized hospitals had about 22% fewer beds than the comparison hospitals, but were otherwise very similar: both types admitted about 35 patients per bed per year and about 65% of their patients were covered by public payers. The privatized hospitals had about 15% lower labor intensity (FTE staff per 100 adjusted admissions) and 17% lower operating expenses per adjusted admission at baseline, implying they were leaner than the comparison group prior to the change in control.

Column 3 presents the corresponding statistics on the 3,925 privately owned hospitals in the data. On almost all measures, private hospitals were noticeably different than their public counterparts. For example, they operated at much greater scale with twice the number of beds as the treated hospitals and discharged more patients per bed (40 versus 35). Public payers accounted for a lower share of their patients (58%). They had similar labor intensity but higher operating costs per admission than the privatized hospitals, suggesting a different cost structure. Hence, private hospitals differ substantially from public hospitals on important operational dimensions and are unlikely to offer a suitable counterfactual to the privatized hospitals. Column 4 presents the corresponding statistics for all 4,985 hospitals in the sample. Since about 80% of the hospitals are privately owned and they serve more patients, the aggregate statistics lean towards those for private hospitals.

Figure 2 describes the phenomenon of hospital privatization in the US over 2000–18. Panel (a) presents a heat map of the US based on the number of privatizations in the state. Privatization was widespread across the country with more than 40 states having at least one. States in the South and Midwest experienced the most number of privatization events during this period. Texas, Georgia, Louisiana, Indiana, and Minnesota are the five states with the most privatizations. Relative to the extant number of public hospitals, Louisiana and Indiana privatized a much greater share of their public hospitals than any other state. However, no state experienced more than 30 privatizations. Therefore, empirical strategies that focus on a few selected states may suffer from low statistical power, highlighting the value of examining this phenomenon using national data. Panel (b) presents the number of privatizations in each year. There were at least 10 privatizations in each year from 2002 through 2017, suggesting that the estimated treatment effects will not be dominated by a specific sub-period. Similarly, no single year accounts for more than 8% of the total number of privatizations. The trend of privatization accelerated following the Great Recession – there were about 16 conversions per year in 2009–18 versus 12 per year over 2000–09.

the sample. We found that about 75% of Medicare patients over 2000–16 were treated at a hospital located within 15 miles of their home zip code, suggesting this is an appropriate threshold. Appendix C.1 provides more details on sample selection.

4 Empirical Strategy

Our goal is to quantify the causal effects of privatization on the affected public hospitals and on the markets they serve. Our baseline models implement a staggered difference-in-differences (D-D) research design, following the recent literature on privatization and ownership in healthcare (Arnold, forthcoming; Eliason et al., 2020). We study privatizations executed over 2000–2018, so we observe each treated hospital for five years before and at least one year after the privatization.[25] Public hospitals that did not experience a change in ownership constitute the comparison group, and they offer an intuitive counterfactual for privatized hospitals.[26]

Equation 2 below presents our baseline model. Y_{ht} denotes the outcome of interest for hospital h in year t. We model the outcome as a function of hospital and year fixed effects, α_h and α_t, respectively. Recent studies of hospital closures have noted that markets experiencing closures had weak economic trends prior to the closures (Alexander and Richards, 2021; Chatterjee et al., 2022). Hence, we test sensitivity to including covariates X_{mt}, a vector of time-varying market attributes including population level, unemployment, poverty, and uninsurance rates for the county in which the treated hospital is located. We do not include time-varying hospital-level covariates (eg. bed capacity, services offered) in the models since most such attributes would plausibly be affected by the privatization. The key regressor of interest, D_{ht}, is a time-varying indicator variable that is equal to one starting in the year the hospital is privatized and zero otherwise. Finally, ϵ_{ht} denotes unobserved time varying factors. We cluster standard errors by hospital to account for the potential correlation of outcomes over time at the same hospital, which is the unit of treatment.

$$Y_{ht} = \alpha_h + \alpha_t + \beta\, D_{ht}\, [+X'_{mt}\,\delta] + \epsilon_{ht}. \tag{2}$$

While our approach is standard in this literature, we note that the privatizations were not randomly assigned, nor are we aware of a credible quasi-experimental instrument for these changes in control. Hence, one should interpret the coefficient of interest, β, with caution. However, our specifications control for the most important potential confounders. For example, hospital fixed effects eliminate persistent unobserved

[25] We require a longer period prior to the privatization in order to test for pre-trends. In robustness checks, we show the results are not sensitive to imposing the restriction that hospitals also be followed for five years after privatization.

[26] Hospitals that exit the sample are retained in the comparison group since this is a valid counterfactual to privatization.

differences between hospitals (and the markets they belong to), an important source of selection. Under the assumption that the privatized and comparison hospitals would have evolved on parallel trends in the absence of the transaction, β recovers the average treatment effect on the treated hospitals. We assess dynamic effects on treated hospital outcomes around the year of the privatization by estimating the event study model in Equation 3 for each outcome.

$$Y_{ht} = \alpha_h + \alpha_t + \sum_{s \neq -1} \beta_s \, D_{h,t+s} + \epsilon_{ht}. \tag{3}$$

A lack of differential trends in the years prior to the privatization is consistent with the identifying assumption. Reassuringly, the evidence suggests relatively large changes in trends following privatization that cannot be explained by pre-trends. We truncate the sample to five years before and after the year of privatization to focus on immediate changes in trajectory following the change in ownership. We also exclude the year of privatization (year zero) since it represents partial treatment and will add to measurement error. In our primary specifications, we estimate unweighted models, thus giving equal importance to all hospitals. Section 5.5 presents results from multiple checks where we assess robustness to using alternate modeling assumptions (including weighting), sample construction rules, and comparison groups.

5 Effects on the privatized hospital

5.1 Hospital finances

A key goal of privatization is to make public hospitals financially sustainable without the need for ongoing government subsidies. Hence, we begin our analysis by examining the effects on hospital finances. Table 3 presents the D-D coefficients obtained by estimating Equation 2 without including the covariate vector X_{mt} in Panel A, while Panel B presents the corresponding results obtained by including controls at the county-year level for population, percent in poverty, percent unemployed, and percent uninsured. For brevity, we examine four outcomes. Column 1 presents the effects on total revenue from patient care (inpatient and outpatient) after all discounts and adjustments. Columns 2, 3, and 4 present the effects on total operating expenses, personnel spending (including benefits), and all non-personnel expenses, respectively. We normalize revenue and expenses by total contemporaneous adjusted admissions in order to account for potential changes in patient volume following privatization. The outcomes should accordingly be interpreted as the mean value per patient. Finally, we use the log of the outcomes rather than the levels to mitigate the influence of outlier values. Accordingly, we interpret the coefficients as approximately estimating the percent change in mean revenue or costs. Figure 3 presents the corresponding event study plots with the dynamic effects on each outcome around the transition.

As the table shows, the estimates are very similar whether we include market-level covariates or not. This is reassuring since it mitigates the concern of model mis-specification and omitted variables like differences in the prevailing economic environment. We prefer to focus on the estimates obtained without including additional covariates as our primary results, hence throughout the text we will primarily discuss these estimates, unless they meaningfully diverge between the two panels. We detect a nearly 6% increase in mean revenue per patient, which is statistically significant at the 5% level. Since the mean revenue per patient is about \$8,100, this implies about a \$485 increase in mean reimbursement. Figure 3 panel (a) shows an increase in mean revenue in the year following privatization and the increase remains consistent in the 5-10% level over the 5 years we track following the transition. Reassuringly, there is no evidence of a differential pre-trend at the privatized hospitals prior to the intervention.

Table 3 Panel A column 2 presents the effect on total operating expense per patient and indicates a modest 3% decline that is statistically insignificant. Figure 3 panel (b) presents the corresponding event study, which confirms no trends before or after privatization. Columns 3 and 4 unpack this result by presenting the effects on personnel and non-personnel costs, respectively. There is a large decline in average personnel cost per patient (col. 3). This measure includes spending on salaries and benefits, normalized by the total adjusted admissions. The coefficient implies a nearly 9% decline in personnel cost. This appears to be a moderate decrease but is quite large when juxtaposed against the fact that privatized hospitals had lower personnel spend per patient at baseline than both private facilities and units in the comparison group (see Table 2). However, the decline in personnel spend is partially offset by small increases in costs elsewhere (col. 4). The coefficient is positive and statistically insignificant. Figure 3 panels (c) and (d) present the corresponding event study plots which are consistent with the average effects implied by the D-D coefficients.

Overall, it appears that privatization meaningfully improves hospital profitability. In the year before privatization, the treated hospitals had an operating margin of about -\$335 per patient, or 4% of mean revenue. Therefore, the 6% increase in revenue alone is sufficient to enable these hospitals to generate a modest surplus. If we include the 3% cost reduction (approximately \$255) in this calculation as well, ignoring the statistical insignificance for a moment, we estimate an increase in operating margin of about \$740 per patient, or 9% of mean revenue. Given the relatively healthy levels of operating cost per patient at baseline at the treated hospitals, it is intuitive that the new private management focuses on increasing mean revenue per patient to improve profitability.

Next, we examine the effects on payer mix and employment in order to discern the mechanisms behind the changes in revenue and personnel spend reported above.

5.2 Patient volume and payer mix

Table 4 presents the corresponding D-D estimates of the effect of privatization on patient volume and payer mix at the privatized hospital. The table follows the same format as Table 3 discussed previously. We present the effects on total patient admissions as well as on the component admissions by payer, to highlight potential heterogeneity in effects for patients accessing care through different payers. Columns 2–4 present results for patients covered by Medicaid, Medicare, and Other payers, which include private and uninsured patients.[27] Figure 4 presents the corresponding event study plots obtained by estimating Equation 3. Total patient admissions at the privatized hospital decline by 8.4% following privatization. This estimate is statistically significant at the 1% level and suggests a substantial contraction of the hospital's patient care services. Figure 4 Panel (a) presents the corresponding event study plot indicating a sharp and persistent decline in volume following the transition.

The decline in patient volume could be due to lower bed occupancy or a broader decline in operational capacity. The former implies a reduction in operating efficiency – a surprising outcome of privatization – while the latter could reflect a strategic decision by the new management to improve finances. To answer this question, we consider the effect on total volume per bed, where beds are updated contemporaneously to account for changes in capacity. Appendix Figure A.3 presents the dynamic effects on total patient volume per bed, obtained by estimating Equation 3. The figure shows a flat trend in total volume per bed following privatization, supporting the latter explanation. The corresponding D-D estimate is economically small and statistically insignificant: -0.3 patients per bed against a mean of 30.8, with a standard error of 0.7. Hence, private firms appear to downsize hospital capacity after they assume control.

Alternatively, the decline in inpatient volume could be partially explained by a change in treatment style

[27] Medicaid and Medicare include those on managed care plans, e.g., Medicare Advantage. The "Other" group is mostly composed of privately insured and self-pay patients. It also includes patients covered by small payers like government employee plans and workers' compensation. Unfortunately, the AHA survey does not provide a breakdown of "Others."

if the hospital treats more cases as outpatients following privatization. We test this conjecture and fail to detect an accompanying increase in outpatient care at the privatized hospitals. Appendix Table A.3 columns 1 and 2 present the corresponding effects on total Emergency Department (ED) and non-ED outpatient logged volumes, respectively. In both cases we find statistically insignificant and negative coefficients, implying, if anything, a decline in outpatient treatment. The coefficients are noisily estimated so we cannot rule out modest increases in outpatient volume. Appendix Figure A.4 panels (a) and (b) present the corresponding event study plots which support the interpretation of no changes in outpatient volume.

We then consider changes in payer mix. Although it appears volume declines across all payers, the decline is not evenly felt by all patient groups. While Medicaid admissions decline by nearly 15%, Medicare admissions only decline by about 5%, and the coefficient is statistically insignificant. Finally, we find a nearly 14% decrease in Other admissions, which includes uninsured admissions. Taken together, we infer that hospital privatization primarily affects non-Medicare patients. Although Medicaid accounted for only 20% of patients at the average treated hospital at baseline, it accounts for 30% of the decline in volume.[28] The event study plots in Figure 4 show that, relative to the non-treated public hospitals, privatized hospitals were not trending differentially on these outcomes prior to the year of transition. This is reassuring and supports the parallel trends identifying assumption. Further, the patterns are consistent with the coefficient magnitudes. For example, there is a noticeable, discrete drop in Medicaid and Other volume in the year after the transition (Panels b and d). As indicated by the dynamic coefficients, the magnitude of the drop in Medicaid admissions persists for at least the five years we follow. This pattern suggests the decline is not a transient phenomenon due to a one-off disruption in management. In contrast, there is little change in Medicare volume at privatized hospitals following the change (Panel c).

The analysis using the AHA data suggests that Medicare patients are not significantly affected by privatization. We further test this interpretation using claims data for Traditional Medicare (TM) patients, as discussed in Section 3. We implement our DD research design and present the results in Appendix Table A.4. We find a decline in TM admissions (Panel A row 1) that is not robust to using a matched sample (Panel B), nor to correcting for the staggered nature of the D-D, thus bolstering the conclusion that Medicare patient utilization is likely unaffected by privatization.[29]

[28]If we apply the estimated percent declines for each payer to the corresponding mean volume at baseline, we predict declines of 93, 68, and 146 patients, respectively, for Medicaid, Medicare, and Others. Hence, Medicaid accounts for 68/307 or 30% of the estimated decline in volume.

[29]We also examine the effects separately for those dually eligible for Medicaid and non-dual eligible TM patients and find a slightly larger decline among dual eligible patients in the baseline specification. However, the effects for both groups tend to be

We hypothesize that the differential decline in Medicaid patients is due to their lower reimbursement rates (Schulman and Milstein, 2019). Payer-specific reimbursement rates cannot be directly observed in the AHA, nor, to our knowledge, in any other national data source. We make some progress in characterizing the relative attractiveness of payers using detailed payer-specific patient volume and reimbursement data for the years 2008–10 reported by California's Office of Statewide Health Planning & Development (OSHPD). We confirm that Medicaid is less lucrative on average than Medicare and privately insured patients, but pays more than the average uninsured patient. Mean Medicaid rates are about 15% lower than mean Medicare rates for inpatient admissions and about 40% lower than the average privately insured patient. Uninsured or self-insured patients pay the lowest rates by far. In the AHA data, the "Other" group contains both privately insured and uninsured patients and its attractiveness therefore depends on the relative magnitude of private insurers. In the California data, privately insured patients contributed about 40% of this group at government hospitals on average. Using this benchmark, mean reimbursement rates for the Other group would be lower than both Medicaid and Medicare.[30] These patterns are consistent with private owners wanting to reduce the share of Medicaid and Other patients at their hospital in order to increase mean revenue per patient.

5.3 Employment

Table 5 presents the estimated effects on employment, suitably normalized by the hospital's adjusted admissions. Column 1 presents the effect on the number of full-time equivalent staff employed by the hospital, which includes both full-time and part-time employees. These are expressed per 100 adjusted admissions for ease of exposition. We find an economically meaningful reduction in total employment of 0.57 FTE per 100 admissions. Compared to the pre-privatization mean, this implies a decrease of 8% in labor intensity, implying that reducing employee count is the primary channel to decrease personnel spending rather than reducing employee compensation.[31] The lack of an effect on compensation differs from the general pattern of changes following privatization (Arnold, forthcoming). However, in this setting it is intuitive since public hospital employees received comparable mean wages at baseline to their counterparts at private hospitals.

non-robust and are statistically indistinguishable.

[30] Mean inpatient reimbursement rates for Medicare, Medicaid, Private insurance and uninsured are $16,360, $13,900, $23,500, and $7,160, respectively (values expressed in 2019 dollars). Assuming privately insured patients are 40% of the Other group, mean reimbursement for the composite group is about $13,700. Source: OHSPD hospital finances reports 2008–10.

[31] In results not summarized here, we estimated a statistically insignificant 3% decline in compensation per employee at the privatized hospital.

Although nurses account for 26% of total staff, we do not detect any reduction in nurse intensity. In contrast, physicians make a tiny share of employed staff, but decline by 30% relative to their strength.[32] The reduction in employment is driven mainly by the residual group, referred here as "Others." This group is disproportionately affected since it accounts for 70% of total FTE but contributes over 90% of the total reduction in labor intensity. This is a diverse group and includes patient care (eg., technicians), back office or overhead (eg., accounting), and managerial functions (eg., administrators).

We also test the possibility whether the decline in employed staff is partially offset by an increase in the use of contract labor following privatization. This is crucial since it affects how we interpret the decline in employment discussed above. If the decline in employment is partly or fully offset by an increase in contract staff, it implies that patient care is likely not affected, and the new management is just changing how it contracts with workers. However, the result in column 5 is near zero and statistically insignificant. We can rule out an increase in contract staff of more than 0.01 FTE per 100 cases (-0.01 + 2 x 0.01), which would offset less than 2% of the estimated decline in employment. We therefore interpret the results as implying a real reduction in labor intensity at the hospital.

Figure 5 presents the event study plots corresponding to each of these outcomes, except contract staff, where we find no effect and therefore exclude for brevity. The dynamic coefficients are consistent with the D-D estimates presented in Table 5. There is a noticeable decline in total physician and other FTEs per 100 adjusted admissions in the year following privatization, and it persists over the next five years. Appendix Table A.5 presents the corresponding estimates obtained using FTE per 100 beds instead. We update the beds contemporaneously to account for potential changes in bed capacity. The results are qualitatively similar, largely driven by the Others group, and suggest a slightly smaller decline in labor intensity than the estimates discussed above (7% vs. 8%). Appendix Figure A.5 presents the corresponding event study plots, which are consistent with the point estimates.

We use our estimated effects to help put in perspective the change in labor inputs following privatization. The average treated hospital had about 7,025 adjusted admissions per year at baseline, which declined by 6% (results not reported), implying a reduction of about 420 cases per year. If labor intensity were held constant at 7.4 FTE per 100 admissions (per Table 5), this alone would merit a reduction of about 31 FTE (7.4 x 4.2 = 31.1) for the average privatized hospital. But the private operators were able to reduce labor

[32] The figures here only account for employed physicians, such as hospitalists. However, for much of the sample period, hospitals typically did not employ physicians directly and this explains the low number of employed physicians.

intensity as well. For the average privatized hospital this implies a reduction of about 38 FTE (0.57 x (7025-420)/100 = 37.6). Hence, the average privatized hospital shed about 70 FTE (13.5% of baseline) in the five years following the change in control, of which 55% was due to more efficient use of labor.

5.4 Heterogeneity in treatment effects

Having discussed the average effects of privatization, we now briefly examine heterogeneity in treatment effects across transitions which plausibly confer different incentives or ability to maximize profits from the hospital. In the interest of brevity, we examine heterogeneity on two dimensions only – the extent of control over the hospital's operations and for-profit status of the new private owners. Greater residual control over the hospital enables greater ability to implement strategies to increase profits, such as cost-cutting. Similarly, private owners organized as for-profit will have a greater incentive to increase profits and growth than those organized as non-profits. We estimate triple difference models to recover the differential effect on the privatized hospital when the new owner has more control or is a for-profit. In both cases we anticipate greater reductions in operating costs as well as in Medicaid volume. However, the effect on total volume could go in either direction depending on how the new owners prioritize growth.

Appendix Table A.6 presents the corresponding main DD and triple difference coefficients. Columns 1–3, 4–7, and 8–10 present the effects on hospital finances, patient volume, and employment, respectively. We exclude outcomes on components (eg., non-personnel spend) where we did not detect changes in the baseline models for brevity. Panel A presents the baseline coefficients for ease of comparison. Panel B presents the triple difference results by residual control, which we classify as described in Section 2. The triple difference coefficients are noisily estimated, but are typically positive and do not fit a pattern of greater reduction in patient volume in deals where the private firm has greater control. In contrast, the effects on employment do seem to follow a consistent qualitative pattern. When the private operators have more control, they appear to eliminate about twice the number of employees. The effects on personnel spending and total expenses are also consistent with this pattern, although the triple difference coefficients are imprecisely estimated.

Table A.6, Panel C presents the corresponding results regarding for-profit status. We find for-profit buyers increase total volume by 6.6% (21.2-14.6 = 6.6), contrasted against non-profit buyers who experience a 14.6% decline in total volume. The difference is statistically significant. This suggests for-profit buyers are more focused on growth, while attempting to move to a more favorable payer mix. For-profit buyers increase the volume of patients covered by all payers *except* Medicaid, which declines by about 6%. The difference between the effect on Medicaid and total volume is far greater in the case of for-profit buyers (-12.4%) than in the case of non-profits (-4.0%). Perhaps this more favorable change in payer mix reflects in the greater increase in revenue at for-profit owned hospitals (Col. 1). In the case of employment, however, for-profit firms do not retrench more than non-profit firms do. In fact, the coefficients suggest that they make smaller magnitude cuts. Ultimately, due to the imprecision of the triple difference estimates, we cannot reject the null hypothesis that for-profits and non-profits behave differently, as with much of the prior literature (Duggan, 2000). Hence, we do not emphasize these results.

5.5 Robustness

We test the robustness of the main results presented above to different modeling assumptions and important validity concerns. Table 6 presents the corresponding results on finances, patient volume, and employment in columns 1–3, 4–7, and 8–10, respectively. Panel A repeats the baseline estimates, without including time-varying covariates, for ease of comparison. Across all checks, the models do not include market-level covariates. The results are very reassuring since the coefficients remain within two standard errors of the baseline estimates across all checks.

Panel B presents the coefficients obtained from regressions incorporating hospital bed capacity as weights.[33] This approach gives more weight to the changes at larger privatized hospitals. Interestingly, the effects on employment, personnel expenses and total expenses increase in magnitude, implying that larger hospitals experience greater cuts after privatization.

Panel C tests whether the estimates are robust to allowing the privatized hospitals to progress on a different linear trend. We estimate models including a linear trend interacted with an indicator for the treated units. The estimates largely remain qualitatively similar. The estimated decline in volume for Other patients is not robust to using this model.

The recent econometric literature on differences-in-differences has shown that estimates obtained from staggered treatment designs may suffer from biases due to the use of treated groups as controls for future treated units. To assess the importance of this potential threat, we report coefficients from the estimator proposed by Callaway and Sant'Anna (2020), which corrects for staggered designs and computes the weighted average treatment on the treated. Panel D presents the corresponding coefficients which are remarkably

[33]For this exercise, we hold bed capacity fixed. For treated hospitals we use the mean of pre-period beds, i.e., the mean of beds in the five years prior to privatization. For control hospitals we use the number of beds in 1999 or if the hospital was not in the sample in 1999 (rare), the first year we observe that hospital.

similar to the baseline estimates across the board.

Panels E and F report results from the baseline specification applied to different samples. Panel E assesses the importance of reducing imbalance in the panel which we allow in our baseline model for the privatized hospitals. That is, while we are able to follow some hospitals for five years following privatization, we can follow others for as little as one year. We assess the importance of this imbalance by limiting the sample to the privatized units that we can follow for at least five years. The results for Medicaid volume are nearly unchanged, while they move toward zero for the other payers, again underscoring the relative robustness of the effect on Medicaid.

Panel F reports results obtained using a matched sample. We implement propensity score matching to identify a subset of the comparison hospital group that resembles the privatized hospitals on key attributes like bed capacity and patient volume in the years just prior to the transition. We use matching to identify a single comparison hospital for each treated hospital without replacement and assign the same year of treatment to the control unit. We limit the data to years -5 through +5 around the year of privatization for *both* treated and control units. Appendix C.2 describes the matching exercise in more detail. The coefficients suggest slightly smaller effects on revenue and patient volume, but larger effects on employment and costs.

The next two panels help assess the importance of some of the sample construction rules we imposed in our baseline approach. We relax these restrictions one at a time and estimate the baseline specification so we can isolate their effect on the estimates. In Panel G, we retain all observations for the treated units, instead of censoring them at +/- 5 years around the privatization. We also retain data from the year of the privatization (year zero). The effect on revenue diminishes and on costs increase in magnitude. In Panel H, we retain hospitals in the comparison group that were recorded in the AHA data as switching between public and private control (and potentially back to public) but these transitions could not be manually validated. The estimates remain nearly unchanged.

6 Effects on the market

6.1 Patient volume

We find that public hospitals persistently admit fewer patients following privatization, and the decline is uneven across patients covered by different payers. From a policymaker's perspective, the result assumes more significance if privatization also causes an aggregate decline in utilization at the market-level, otherwise this could be interpreted as business stealing or patient reallocation. Being reallocated to a different hospital could potentially be harmful if the new hospital is further away or of worse quality than the privatized hospital, but at the same time it may also be an improvement if the public hospital was of lower quality. However, a reduction in access to care implies Medicaid (and perhaps other) patients are unambiguously worse off following a privatization. If Medicaid patients are perceived as unprofitable or undesirable, then alternate hospitals may be reluctant to step in and offset the decline at the privatized hospital.

To shed light on this concern, we adapt our research design and implement it at the market-level, which we define using Health Service Areas (HSAs). These were originally delineated by the US Census in a similar fashion and for the same purpose as the more commonly used Hospital Referral Regions (HRRs), developed by the Dartmouth Atlas group. We prefer to use HSAs for two reasons. First, they are smaller in size – there are 910 HSAs against 306 HRRs. The average HSA has about five hospitals (including both public and private owned), while the average HRR contains about 18. Hence, we will have greater statistical power to detect the market-level effects of a single hospital's privatization when we use a more granular market definition. At the same time, HSAs adequately capture patient hospital choice decisions.[34] Second, their borders follow county boundaries, while those of HRRs do not. This allows us to directly map the time-varying, county-level characteristics to HSAs.

To implement our analysis at the market-level, we tag the 204 markets in which privatized hospitals are located as treated, while the 706 remaining markets form the comparison group.[35] We then estimate an unweighted market-year level model equivalent to that presented in Equation 2. A market is considered treated when it first experiences a privatization during our sample period, and is then considered treated through the end of the sample (42 of the 204 markets experienced more than one privatization event). Table 7 describes the market-level analysis sample. Columns 1 and 2 are equivalent to the corresponding columns in Table 2. We also present some market-level economic characteristics, such as poverty and unemployment. The average treated market has 5.4 hospitals, out of which 1.3 or 24% are treated during the period. Market-level bed counts, payer mix, and economic indicators are as one would expect based on the hospital-level averages. Comparison markets are slightly smaller in size and have slightly better economic indicators on average (eg., lower poverty and unemployment).

[34]Using Medicare claims data, we confirm that more than 70% of TM patients choose a hospital located in the same HSA as their residence zipcode. The corresponding number for HRRs is about 80%.

[35]We considered imposing a non-neighbor rule for comparison markets to mitigate the potential for spillovers. But such a rule would nearly eliminate all potential untreated markets in the same states as the treated markets. It was unappealing to have the comparison group be restricted to an almost disjoint set of states.

Table 8 presents the estimated effects on market-level patient volume, with log of patient volume as the outcome. The columns present effects on total volume and by payer. Panels A and B present the average effects from specifications without and with the time-varying controls, respectively. Including market controls tends to magnify the point estimates but leads to similar interpretations, hence we continue to focus on the estimates without controls. Column 1 presents estimates on total volume and reports a 0.9 percentage point (pp) decline in admissions across all hospitals in the market. The direct effect at the treated hospitals was an 8.4% decline in admissions. Since treated hospitals are about 24% of total market capacity on average, we would expect a 2.0 pp decrease in total admissions at the market level based on the direct effect alone (24% of 8.4 pp). Hence, the point estimate we obtain suggests the presence of some offsetting responses by other hospitals in the same market. However, we are under-powered to statistically detect an effect of this magnitude at conventional levels of significance.

The key finding is that Medicaid patients experience a meaningful aggregate decline in volume at the market-level following a privatization. While the point estimates for Medicare and Others are close to zero, the effect on Medicaid is -4.2 pp – approximately what we would predict based on the privatized hospital's decline alone (24% of -14.9, or -3.6 pp). Hence, the point estimate suggests no offsetting responses by other local hospitals for Medicaid patients. The coefficient is noisily estimated so we cannot reject the null hypothesis of no decline in Medicaid volume, although it is larger and significant at the 5% level when we include controls. Figure 6 presents the corresponding event study plots for these outcomes. The estimated dynamic effects are consistent with the coefficients discussed above. Medicaid is the only payer for which the trend appears to be consistently negative following privatization.

Heterogeneity

We examine heterogeneity in the effects on aggregate patient volume across markets. This helps uncover whether the decline in Medicaid may be larger in certain types of markets, and is therefore of salience to policymakers considering privatization requests. Secondly, it also helps test whether the decline in Medicaid volume could be due to factors orthogonal to market structure. For example, one argument in favor of restricting Medicaid inpatient utilization is that some of these are potentially wasteful stays and the patients can be adequately treated in outpatient settings instead. However, if privatization simply allows hospitals to re-calibrate their admission protocols and filter out potentially wasteful stays, we would not expect much heterogeneity across markets based on baseline attributes of the hospitals, like their level of market power.

The first dimension of interest is the level of concentration in the local hospital market. Shleifer (1998) hypothesizes that privatization will have less beneficial effects in more concentrated markets since consumers have fewer outside options and therefore market forces cannot discipline the new managers at these firms. Vickers and Yarrow (1991) also note in their review of the evidence that privatization does not appear to have any benefits on productivity and growth when markets are not competitive. This is a highly pertinent issue in the case of hospitals since local hospital markets are concentrated on average – the mean Herfindahl Hirschman Index (HHI) in 2000 was nearly 3,000, well over the federal government's threshold for being "highly concentrated" (DOJ, 2010). The mean HHI further increased to about 4,000 by 2020.[36] We test this hypothesis in our setting by examining if the negative effect on Medicaid and total utilization is greater in more concentrated markets. We estimate triple difference models where we include an interaction term between treatment and a concentration level greater than the median in 1999. Note that concentrated markets have half the number of hospitals as the average treated market. Hence, the privatized hospital has more influence than usual. This fact alone suggests the effects will be larger in concentrated markets. Further, the remaining hospitals are more exposed to privatization than in the average market, amplifying the possibility of spillover responses.

Table 8, Panel C presents the estimated coefficients of interest from models without including market-level covariates. The results imply that the effects of privatization are diametrically opposed in markets with low versus high concentration. Utilization does not decline in competitive markets, and even increases a bit, though even in this case we do not detect a statistically significant increase for Medicaid patients. There is a sharp decline in utilization in concentrated markets, with a 6.2 pp decline in aggregate volume (4.3-10.5 = -6.2). While volume declines in more concentrated markets across all payers, the decline is most pronounced for Medicaid patients at -12.0% versus -6.0% for the next most affected payer. In results not reported here, we investigated the determinants of the larger decline in Medicaid volume in concentrated markets, relative to the average effect. We find the decline in volume at the privatized hospital is 10% larger in concentrated markets (16.5% vs. 15% overall). The privatized hospital also contributes a greater share of the market (44% vs. 24%) in these markets and together these predict a decline of about 7.3% (44% x 16.5%) without any response by the remaining hospitals. These results imply a substantial fraction of the aggregate decline cannot be explained by the actions of the privatized hospital alone (4.7/12 or 40%). Hence, the remaining

[36]We computed these HHI values using hospital bed shares recorded in the AHA and hospital referral regions to define hospital markets. Since HSAs are smaller and have fewer hospitals, the mean HHI would be greater if we used HSAs to define hospital markets.

hospitals in these markets also cut their Medicaid admissions when exposed to a privatization.

Table 8, Panel D presents the corresponding triple difference coefficients testing for differential effects in markets with greater poverty levels. The results clarify that privatizations barely register in markets with below-median poverty rates. All D-D coefficients, which estimate the effects for low poverty markets, are positive, small, and statistically insignificant. In contrast, markets with greater poverty rates experience an aggregate decline in patient volume of 3.4 pp (1.6 - 5.0 = -3.4), which is marginally significant. This is driven mostly by a large and statistically significant decline in Medicaid volume of 11.6 pp (3.2 - 14.8 = -11.6). In results not presented here we confirmed that privatized hospitals downsize similarly in both types of markets, hence this result does not reflect greater direct effects in markets with greater poverty levels. The contrasting effects on Medicaid volume in these two groups of markets suggests that neighboring hospitals in more affluent markets offset the decline of hospital operations following privatization. However, not only does this offsetting mechanism not operate in lower income markets, but also the neighboring hospitals appear to reduce their own intake of Medicaid patients since the impact on Medicaid is too large to be explained by the direct effect alone. Note that highly concentrated markets partially overlap with high poverty markets (56 out of 204 treated markets are above-median on both poverty and concentration levels), but overall, the two groups appear quite different. Concentrated markets are not as economically disadvantaged, having similar poverty, Medicaid, and uninsurance levels as the average treated market.

6.2 Employment

Previous studies on privatization have found spillover effects of privatization on market-level wages. Arnold (forthcoming) studies privatization in Brazil and finds substantial spillover effects on mean wages at exposed firms in the market. The aggregate decline in wages is nearly three times what would be predicted based on the effect on the privatized firm alone. In our setting, however, we did not find a direct effect on wages at the privatized hospital itself. We did find an effect on employment and therefore focus our attention on the aggregate effect on employment in the local hospital industry.

Table A.7 presents the corresponding effects on market-level hospital labor intensity obtained by applying our research design and in the usual format. Panels A and B present the average effects across all markets, while Panel C presents the results from a heterogeneity test described below. The columns present effects on total FTE (col. 1) and on the same components studied in Table 5. The average effects are estimated to be small and statistically insignificant for all outcomes. We can reject a change of more than

0.24 FTE per patient at the market-level. This range includes the effect we would expect based on the direct effect alone (24% of 0.57 = 0.14 FTE).

As with the market-level effects on utilization, the average effect on labor intensity may mask heterogeneity across different types of markets. Previous studies have noted the influential role of unions in preventing employment losses (Atanassov and Kim, 2009). Some have noted the importance of powerful unions in perhaps preventing privatization in the first place (Shleifer, 1998). We test the hypothesis that the private operator is able to implement greater staff cuts in markets where unions are less influential. We measure union prevalence using state-year level data on the share of unionized employees in the private services sector.[37] We assign markets with below-median values of this measure (about 4%) in 1999 to weak union presence. We then implement a triple difference specification testing for differential effects in markets with weaker unions. Table A.7, Panel C presents the corresponding results. Consistent with the hypothesis that unions prevent large employment cuts, we find a greater reduction in market-level labor intensity in markets with weaker unions. The difference between markets with low versus high union presence tends to be statistically significant across outcomes. Taking the case of total FTE, the aggregate effect in low-union presence markets is -0.2 FTE per 100 adjusted admissions (0.13 - 0.33 = -0.2). The direct effect on employment at the privatized hospital can account for about half of this aggregate effect.[38] This suggests that competing hospitals in these markets also reduce employment in response and these spillovers contribute the remaining half. It is unclear why the competing hospitals did not find these efficiency gains in employment earlier, but the result is consistent with the claim that unions provide some protection against job losses.

Appendix Table A.8 presents results from robustness checks on the baseline estimated effects on market-level patient volume and hospital employment. We implement the same checks as we did for the corresponding effects on the privatized hospital (Table 6), and hence the table follows the same format. Since the baseline coefficients on aggregate patient volume and employment were imprecisely estimated, we focus on the robustness of the decline in Medicaid volume and whether it remains greater than the overall decline in patient volume. Reassuringly, we find this same pattern repeats across the tests. Weighting markets by beds so that larger markets have more influence on the estimate pushes the decline in Medicaid toward zero, further corroborating the result in Section 6.1 that the decline in Medicaid volume is disproportionately felt

[37]The measure of unionization by state was computed using the data maintained by Hirsch and Macpherson (2022) at `www.unionstats.com`. They combine multiple data sources, but mainly rely on the CPS.

[38]In unreported results, we find an effect of -0.48 FTE at the privatized hospital in these markets, and they account for about 22% of the market. Hence, the direct effect alone predicts a reduction of 0.22 x 0.48 = 0.11 FTE at the market-level.

in markets with fewer hospitals.

6.3 Discussion

We find that privatizing public hospitals leads to a decline in the number of patients served by them and also to an aggregate decline in the number of hospital stays in the local market served by the privatized hospital. These effects are heterogeneous across payer types, with Medicaid patients being the most affected at the market-level, and across market types, with concentrated markets experiencing the greatest aggregate decline in volume.

One of the key channels through which Medicaid has been shown to improve health outcomes for beneficiaries is by improving access to and utilization of medical care (Currie and Gruber, 1996a,b; Miller et al., 2019). Hence, the aggregate decline in Medicaid volume potentially hurts its effectiveness as a social insurance program that ensures access to medical care for vulnerable low-income beneficiaries. Privatization therefore emerges as a channel that may curb utilization of care by Medicaid beneficiaries, joining other factors such as low reimbursement rates and various forms of utilization restrictions for those who receive their Medicaid coverage through managed care (Aizer, Currie and Moretti, 2007). National statistics on hospital utilization and spending on Medicaid are consistent with the notion that Medicaid beneficiaries face additional barriers to accessing care. For example, although Medicaid enrollment has grown faster than that of Medicare over 2000–19 (61% versus 55%), aggregate spending on Medicaid has grown at a much lower pace than for Medicare (106% versus 140%). Adjusting for the differential growth in enrollment enhances the difference – spending per enrollee has grown nearly twice as fast for Medicare than for Medicaid on average during this period (2.9% per year vs. 1.5%).[39]

Along similar lines, the growth in hospital utilization by Medicaid patients (about 18% over 2000–19) appears low compared to predictions using estimates, taken from the literature, of hospital utilization elasticity with respect to Medicaid coverage. The elasticities span a wide range, starting as low as 30% reported by the Oregon health insurance experiment, up to 280%, reported by a quasi-experimental study of the Medicaid expansion mandated by the Affordable Care Act (Finkelstein et al., 2012; Dunn et al., 2021). The estimate from the Oregon study represents a partial equilibrium elasticity and may understate the general equilibrium effect on utilization. We prefer the estimate of 100% reported by Duggan, Gupta

[39]Medicaid enrollment information sourced from Kaiser Family Foundation which provides granular data by state. Medicare enrollment and spending information on both programs sourced from the National Health Expenditure Tables, 2020. We deflate spending to 2019 dollars using the CPIU reported by the BLS.

and Jackson (2022) using data from California, which reflects general equilibrium effects and falls near the middle of this range. Applying this elasticity to the observed growth in Medicaid enrollment (38%, reported by the Kaiser Family Foundation), we expect a 38% increase in Medicaid hospital utilization. Hence, the observed growth in Medicaid hospital utilization is about 20 percentage points (pp) lower, or nearly half, than what estimates from the literature predict. Our results imply that privatization can explain about 4 pp, or about 20% of this gap in the markets exposed to privatization, assuming that the affected markets are representative of the average market.

7 Conclusion

Privatization can improve profitability and growth of public firms, but may hurt some stakeholders. This trade-off assumes greater significance in the case of hospital care, which has unique challenges and has experienced substantial privatization. Yet this phenomenon has largely been ignored in the economics literature. We provide novel evidence from the privatizations of public hospitals in the US over 2000–2018. We confirm that privatization improves hospital profitability sufficiently so that the hospitals transition from being loss-making to generating a modest surplus. The primary mechanism is increasing the average revenue per patient. The effects on operating costs are small and statistically insignificant, even though hospital employment declines following privatization.

Private owners increase mean revenue per patient by changing their patient mix toward more lucrative payers. As a result, the intake of low-income Medicaid patients disproportionately declines. This strategy has larger implications beyond the financial health of the privatized hospital since we detect a decline in aggregate Medicaid volume at the market-level, raising the concern of a reduction in access to care for these vulnerable patients. The evidence is consistent with exercise of market power since the reduction in Medicaid volume is magnified in concentrated markets. Our results imply privatization is quantitatively important in explaining why Medicaid utilization and spending has lagged enrollment.

There are several avenues for future research. Since we wanted to characterize the effects of privatizations nationally, we were limited to using hospital-level data. The lack of granular data imposes several limitations. We cannot describe whether hospital utilization declined to a greater extent for certain services based on profitability (eg., unprofitable services like psychiatric care) or their social value (eg., necessary versus avoidable stays). Most importantly, we could not shed light on the effects for the uninsured, another important vulnerable group of patients. Researchers with access to patient-level data, perhaps focused on narrower geographies or from specific payers, can make progress on these questions as well as quantifying the effects on quality of care. Along similar lines, future studies can use more granular employment data to examine labor and wage dynamics for workers in the local hospital market. Understanding these aspects will be key to comprehensively quantify the welfare effects of privatization and inform policy interventions.

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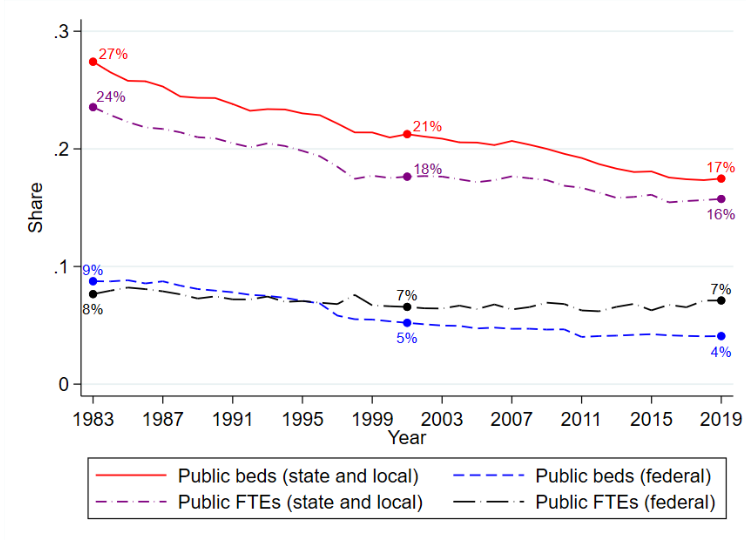
(a) As provider



(b) As payer

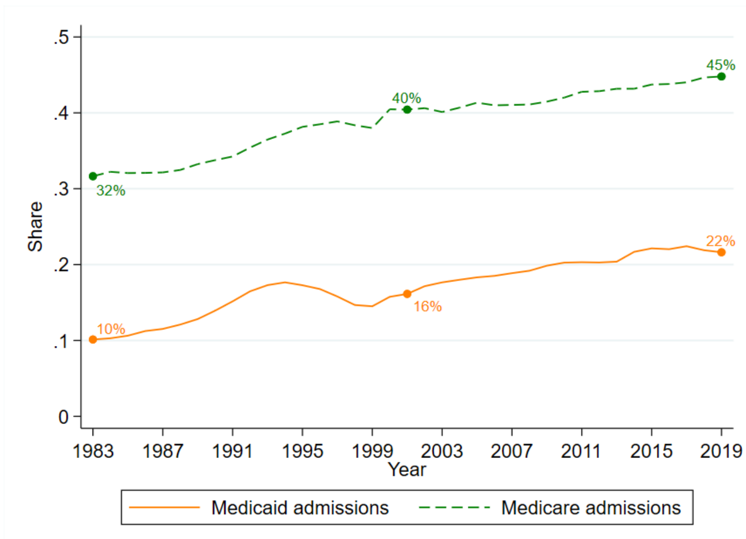
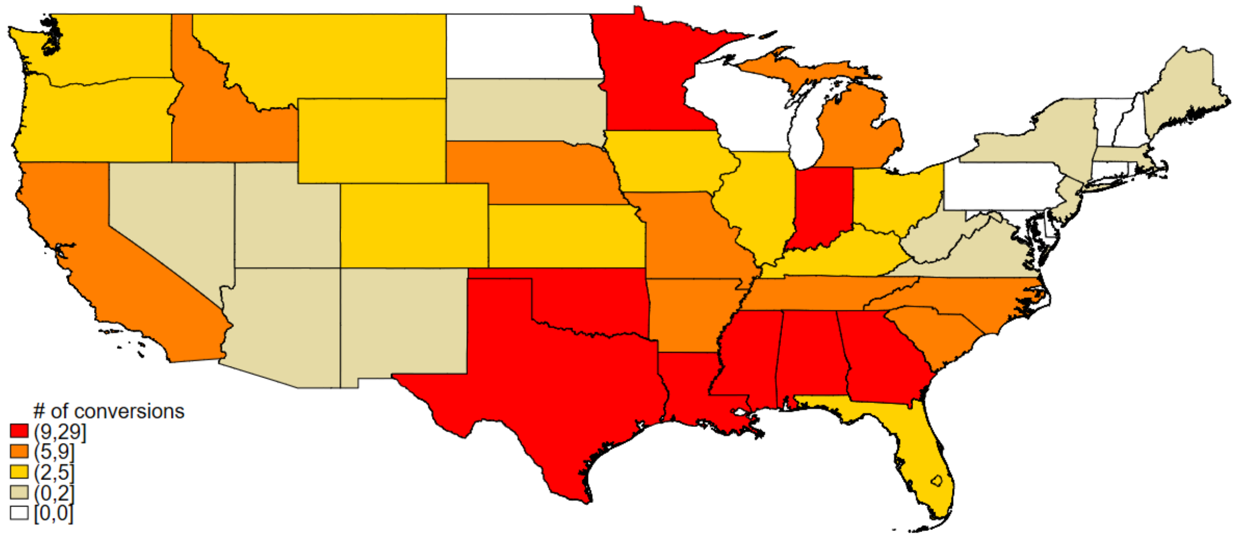


Figure 1: Government role in hospital care

Note: The figure presents overall shares in the US from 1983 through 2019 using American Hospital Association (AHA) survey data. Non-general-acute-care hospitals were included in the sample for share calculations. In Panel A we plot the share of total beds and full-time equivalent employees (FTEs) contributed by public, non-federal hospitals (red and purple lines, respectively) and by public, federal hospitals (blue and black lines, respectively). In Panel B, the share of Medicaid admissions is given by the orange, solid line; the share of Medicare admissions is given by the green, dashed line. For Panel B, the denominator comprises all non-federal hospitals present in the survey in each year.

(a) By state



(b) By year

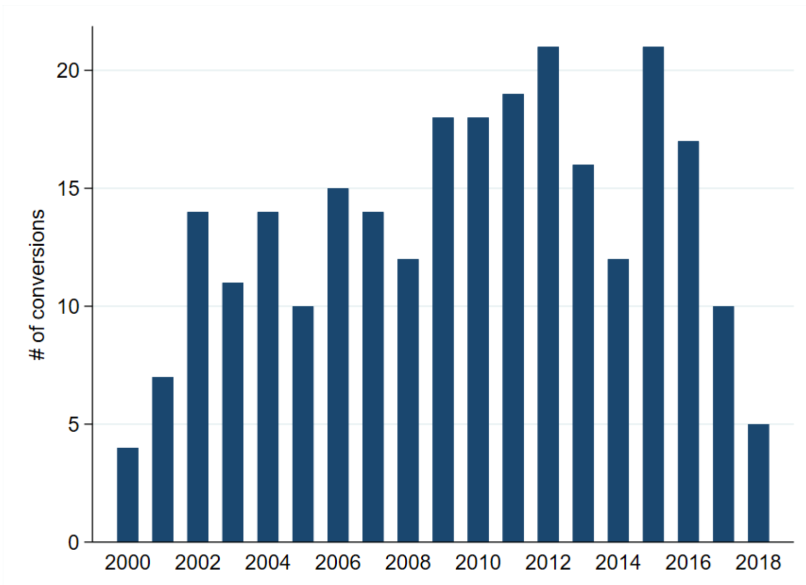
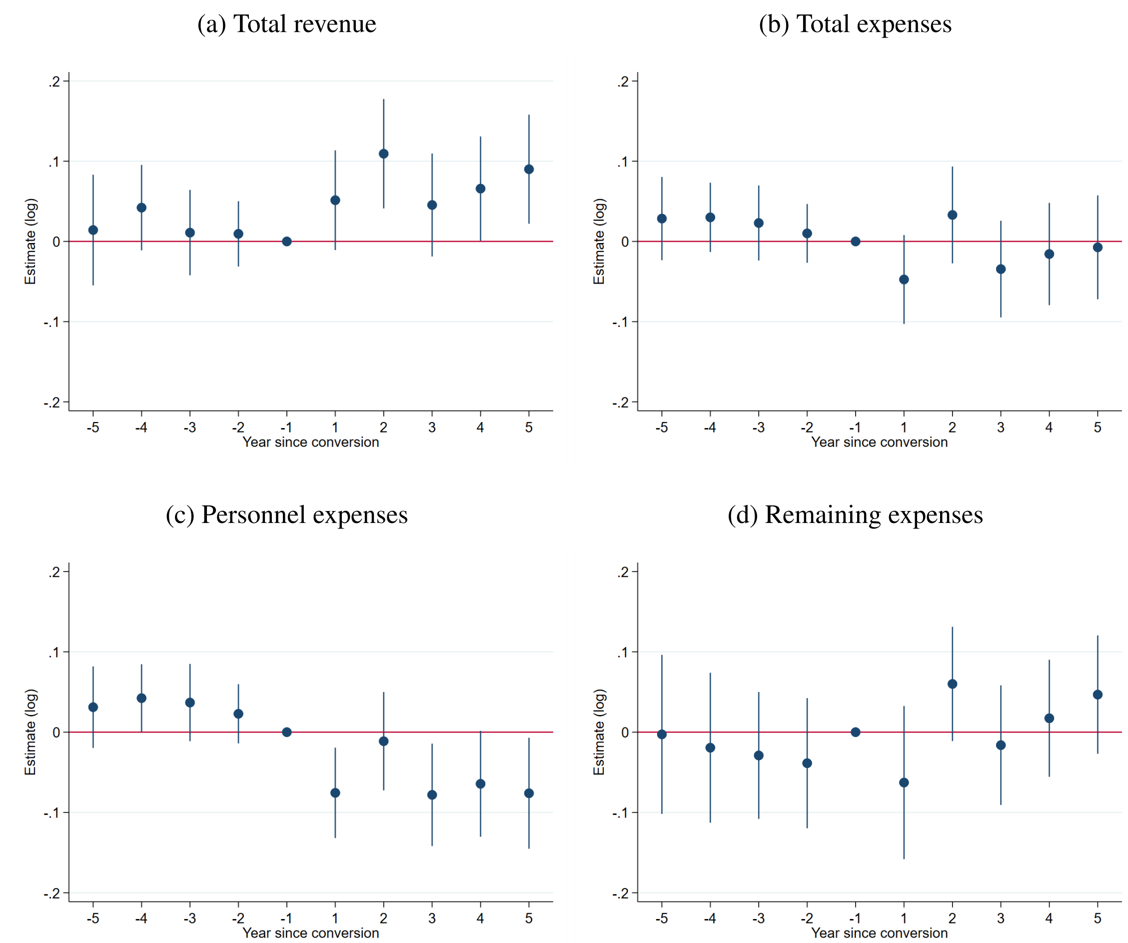


Figure 2: Privatizations

Note: The figure presents the distribution of non-federal, public-hospital privatizations during our sample period (2000–18). We restricted our sample to general-acute-care hospitals. Panels (a) and (b) present the distribution by state and by year, respectively. Hawaii and Alaska are not pictured and include 4 and 1 conversions, respectively. We manually validated each conversion.



(a) Total revenue

(b) Total expenses

(c) Personnel expenses

(d) Remaining expenses

Figure 3: Effects on (log) finances

Note: The figure presents event study plots obtained by estimating Equation 3 on hospital-year level data. The control group is comprised of hospitals that remain public throughout our sample period. Outcomes in panels (a) and (b) are total revenue (from Medicare cost reports) and total expenses (from AHA), respectively. Total expenses comprises personnel expenses and remaining expenses, shown in panels (c) and (d), respectively. All outcomes are normalized by contemporaneous, adjusted admissions and presented in logs. Adjusted admissions are admissions scaled by the ratio of outpatient to inpatient revenue. Year zero is the year of privatization and is excluded for treated hospitals since it represents partial treatment. The error bars present 95% confidence intervals. Standard errors are clustered by hospital.

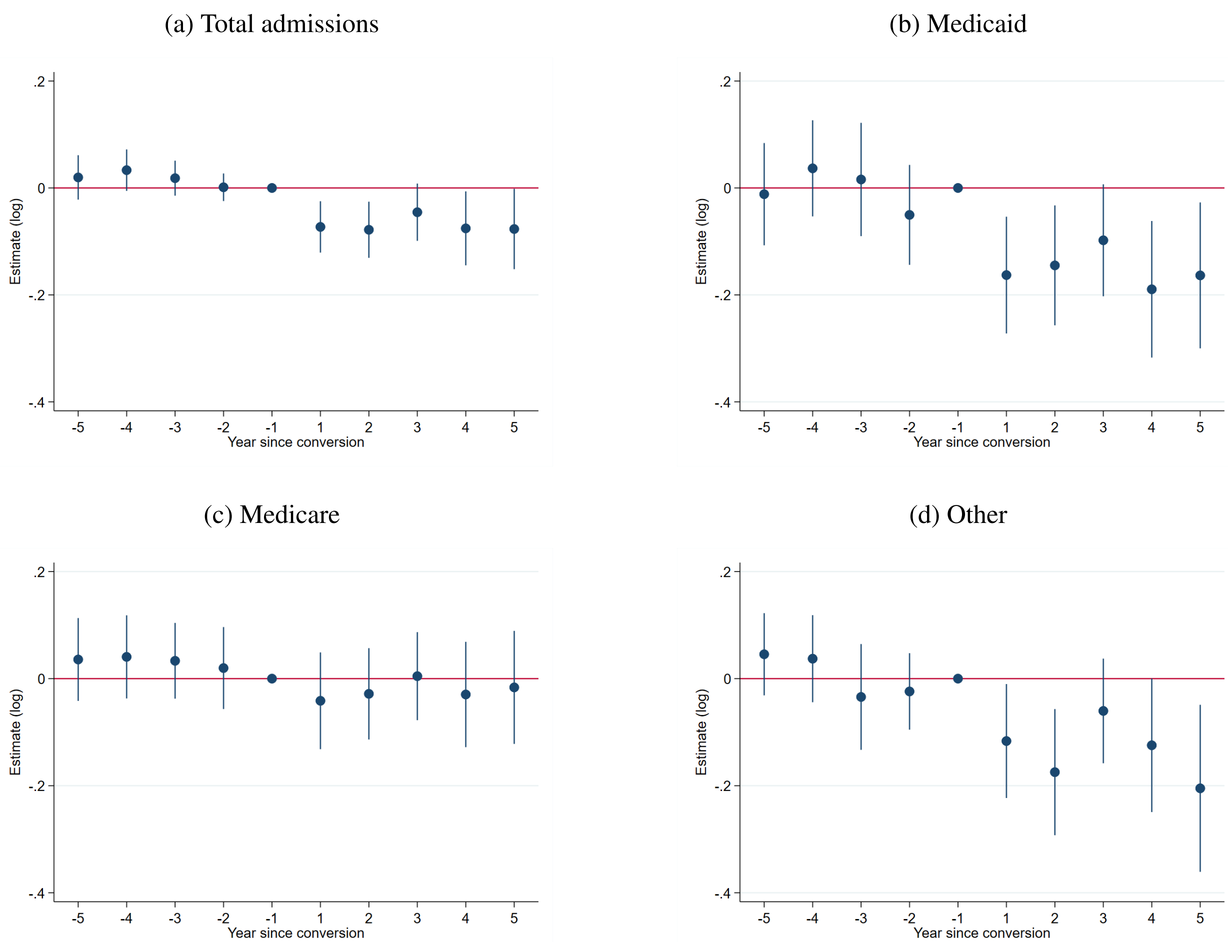


Figure 4: Effects on patient (log) volume

Note: The figure presents event study plots obtained by estimating Equation 3 on hospital-year level data. The control group is comprised of hospitals that remain public throughout our sample period. The outcomes are log total, Medicaid, Medicare, and other admissions in panels (a), (b), (c), and (d), respectively. ‘Other’ admissions refers to hospital admissions not covered by Medicaid or Medicare and mostly comprises privately insured and uninsured patients. Year zero is the year of privatization and is excluded for treated hospitals since it represents partial treatment. The error bars present 95% confidence intervals. Standard errors are clustered by hospital.

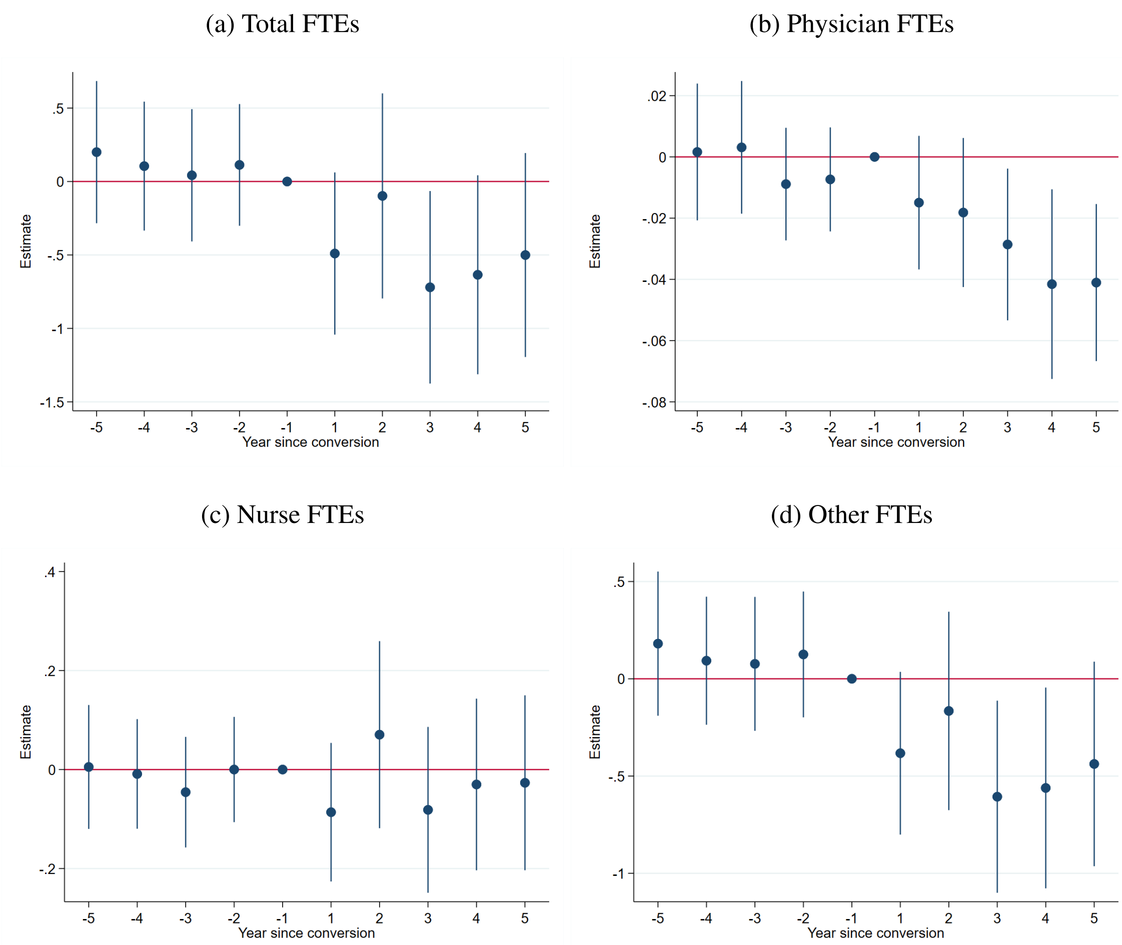


Figure 5: Effects on staff (per 100 adjusted admissions)

Note: The figure presents event study plots obtained by estimating Equation 3 on hospital-year level data. The control group is comprised of hospitals that remain public throughout our sample period. Outcomes are total full-time equivalent employees (FTEs), physician FTEs, nurse FTEs, and other FTEs in panels (a), (b), (c), and (d), respectively. We normalize staff levels in each column by contemporaneous, adjusted admissions, which scales admissions by the ratio of outpatient to inpatient revenue. We then multiply outcomes by 100 for ease of exposition. Year zero is the year of privatization and is excluded for treated hospitals since it represents partial treatment. The error bars present 95% confidence intervals. Standard errors are clustered by hospital.

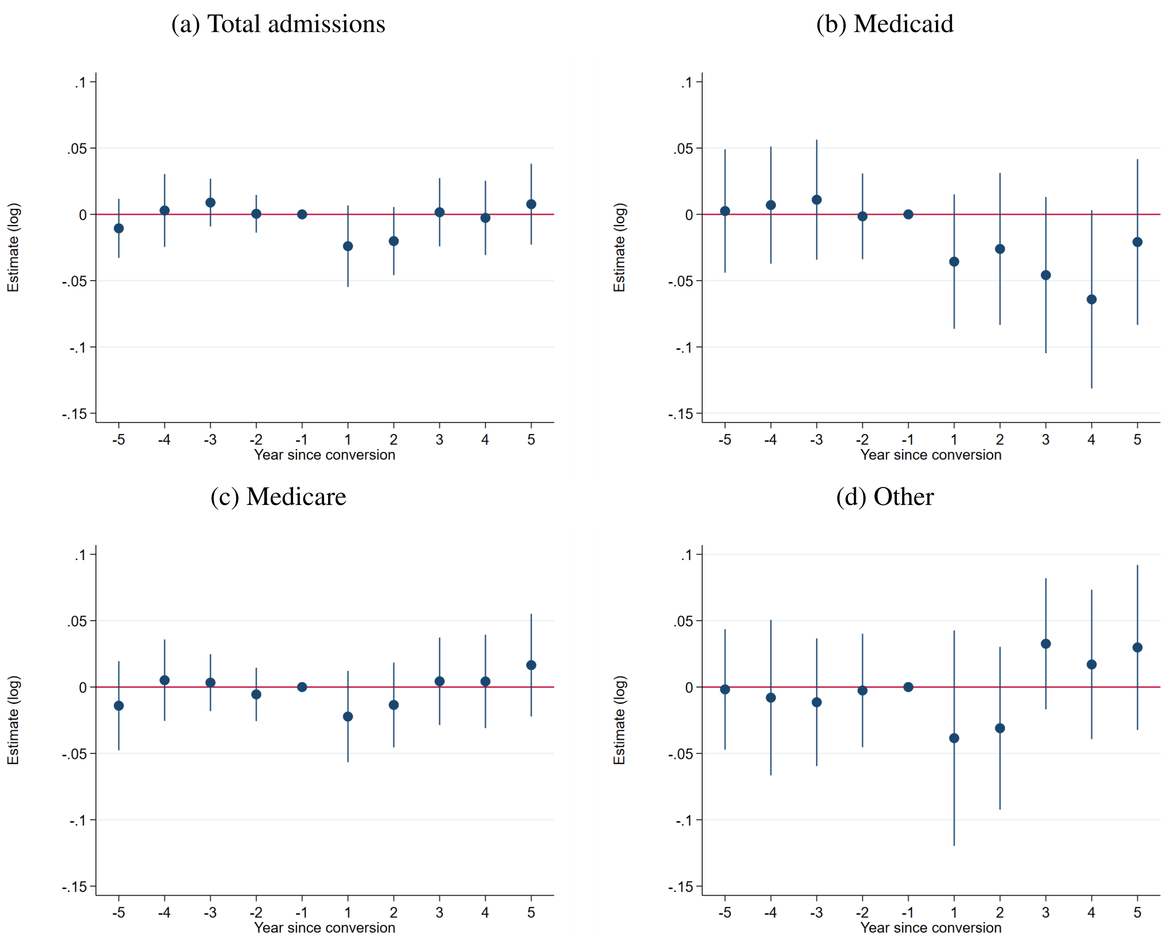


Figure 6: Effects on market (log) volume

Note: The figure presents event study plots obtained by estimating the market-level equivalent of Equation 3 on market-year level data. We define hospital markets using Health Service Areas (HSAs), as described in Section 6. The outcomes are log total, Medicaid, Medicare, and other admissions in panels (a), (b), (c), and (d), respectively. ‘Other’ admissions refers to hospital admissions not covered by Medicaid or Medicare and mostly comprises privately insured and uninsured patients. Year zero is the first year a privatization occurs in a given market and is excluded for treated markets since it represents partial treatment. The error bars present 95% confidence intervals. Standard errors are clustered by HSA.

Table 1: Shares of hospital beds by ownership type for select states in 2019

	(1) AL	(2) CA	(3) TX	(4) GA	(5) IL	(6) PA	(7) US Overall
Public (non-federal)	44.4	22.9	15.8	11.7	8.0	3.8	17.3
							(12.5)
Public (federal)	4.4	3.6	5.8	3.4	3.7	3.6	4.2
							(2.1)
Non-profit	23.4	56.8	37.1	71.5	80.8	79.3	62.9
							(19.2)
For-profit	27.8	16.8	41.3	13.4	7.5	13.3	15.6
							(12.4)
# hospitals	116	419	588	172	208	235	6,090

Notes: The table presents shares of hospital beds by ownership type for select states using American Hospital Association survey data from 2019. Appendix A.1 lists public (non-federal) hospital bed shares for all states. Non-general-acute-care hospitals were included in the sample for share calculations. Column 7 shows mean shares for the overall US; standard deviations are shown in parentheses.

Table 2: Descriptive statistics

	(1) Privatized	(2) Remaining Public	(3) Private	(4) All
% public	100.0	100.0	0.0	21.3
% for-profit	0.0	0.0	21.1	16.6
% non-profit	0.0	0.0	78.9	62.1
Admissions	3,095	4,154	7,461	6,703
	(4,404)	(6,917)	(7,702)	(7,587)
Beds	93	120	186	171
	(105)	(162)	(179)	(176)
% Medicaid adm	15.4	16.6	13.0	13.7
	(8.6)	(12.4)	(8.8)	(9.5)
% Medicare adm	49.2	47.4	44.5	45.2
	(15.6)	(16.8)	(13.1)	(13.9)
% other adm	35.4	36.1	42.4	41.0
	(14.4)	(14.0)	(13.9)	(14.2)
Total FTEs/100 adj adm	7.7	9.1	7.5	7.8
	(5.2)	(5.8)	(4.4)	(4.7)
Total expenses/adj adm	7,320	8,810	9,098	8,960
	(4,003)	(5,302)	(4,647)	(4,744)
Personnel expenses/adj adm	3,909	4,765	4,646	4,627
	(2,082)	(2,827)	(2,287)	(2,379)
# hospitals	258	802	3,925	4,985

Notes: The table presents descriptive statistics on a cross-section of hospitals in the analysis sample. We use values from 1999 for most hospitals. In rare instances in which we do not observe a hospital in 1999, we use values from that hospital's first year in the data. Appendix C.1 describes the sample construction restrictions in detail. Column 1 describes public hospitals that privatized during the sample period. These comprise the treated units. Column 2 describes the primary comparison group: public hospitals that did not experience a change in ownership during this period. Column 3 describes all privately-owned, non-profit and for-profit hospitals that were not converted to public ownership during this period. Column 4 presents the corresponding values on the full sample. 'Other' admissions refers to hospital admissions not covered by Medicaid or Medicare and mostly comprises privately insured and uninsured patients. Standard deviations are shown in parentheses.

Table 3: Effects on (log) finances

	(1) Total revenue	(2) Total expenses	(3) Personnel expenses	(4) Remaining expenses
A: No controls				
DD	0.057	-0.033	-0.086	0.024
	(0.026)	(0.024)	(0.024)	(0.036)
Obs	16,662			
B: Market controls				
DD	0.064	-0.027	-0.080	0.031
	(0.026)	(0.024)	(0.024)	(0.036)
Obs	16,651			
Mean outcome (t-1)	8,109	8,444	4,604	3,840

Notes: The table presents effects on revenue and expenses at the privatized hospitals, obtained by estimating Equation 2 on hospital-year level data. All outcomes are normalized by contemporaneous, adjusted admissions and presented in logs. Adjusted admissions are admissions scaled by the ratio of outpatient to inpatient revenue. Column 1 presents results for total revenue (inpatient plus outpatient revenue minus contractual allowances and discounts), obtained from Medicare cost reports. Column 2 presents results for total expenses, which comprises personnel expenses (Column 3) and remaining expenses (Column 4), all of which are obtained from the American Hospital Association survey. Panel A reports coefficients from a two-way fixed effects specification with no covariates. Panel B reports coefficients from a two-way fixed effects specification including time-varying, county-level controls as described in Section 4. Panel B has fewer observations since the market-level covariates are not available for 1995 and 1996. The mean values pertain to outcomes (in levels) at privatized hospitals in the year prior to privatization. Standard errors are clustered by hospital and are presented in parentheses.

Table 4: Effects on patient (log) volume

	(1) Total	(2) Medicaid	(3) Medicare	(4) Other
A: No controls				
DD	-.084	-.149	-.049	-.138
	(.027)	(.042)	(.030)	(.043)
Obs	20,998			
B: Market controls				
DD	-.090	-.157	-.056	-.142
	(.027)	(.042)	(.030)	(.043)
Obs	19,385			
Mean outcome (t-1)	3,014	617	1,351	1,046

Notes: The table presents estimated effects on patient volume at the privatized hospitals obtained by estimating Equation 2 on hospital-year level data. Columns 1, 2, 3, and 4 present the effects on log total, Medicaid, Medicare, and other admissions, respectively. ‘Other’ admissions refers to hospital admissions not covered by Medicaid or Medicare and mostly comprises privately insured and uninsured patients. Panel A reports coefficients from a two-way fixed effects specification with no covariates. Panel B reports coefficients from a two-way fixed effects specification including time-varying, county-level controls as described in Section 4. Panel B has fewer observations since the market-level covariates are not available for 1995 and 1996. The mean values pertain to patient volume at privatized hospitals in the year prior to privatization. Standard errors are clustered by hospital and are presented in parentheses.

Table 5: Effects on staff

	(1) Total	(2) Physician	(3) Nurse	(4) Other	(5) Contract
A: No controls					
DD	-0.57	-0.03	-0.02	-0.52	-0.01
	(0.26)	(0.01)	(0.06)	(0.19)	(0.01)
Obs	20,998				8,632
B: Market controls					
DD	-0.52	-0.03	-0.01	-0.48	-0.01
	(0.26)	(0.01)	(0.07)	(0.19)	(0.01)
Obs	19,385				8,628
Mean outcome (t-1)	7.40	0.10	1.90	5.30	0.20

Notes: The table presents effects on full-time equivalent (FTE) employed staff at the privatized hospitals, obtained by estimating Equation 2 on hospital-year level data. Column 1 presents results for total FTEs, which comprises of physician, nurse, and all others, presented in columns 2, 3, and 4, respectively. We normalize staff FTEs in each column by contemporaneous, adjusted admissions; we then multiply outcomes by 100 for ease of exposition. Adjusted admissions are admissions scaled by the ratio of outpatient to inpatient revenue. Column 5 presents results for contract FTEs, which come from Medicare cost reports and include management and patient care staff. Panel A reports coefficients from a two-way fixed effects specification with no covariates. Panel B reports coefficients from a two-way fixed effects specification including time-varying, county-level controls as described in Section 4. Panel B has fewer observations since the market-level covariates are not available for 1995 and 1996. The mean values pertain to outcomes (in levels) at privatized hospitals in the year prior to privatization. Standard errors are clustered by hospital and are presented in parentheses.

Table 6: Robustness checks

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
		Finances			Volume				Staff	
	Revenue	Tot. Exp.	Pers. Exp.	Total	Medicaid	Medicare	Other	Total	Physician	Other
A. Baseline	0.057	-0.033	-0.086	-0.084	-0.149	-0.049	-0.138	-0.57	-0.03	-0.52
	(0.026)	(0.024)	(0.024)	(0.027)	(0.042)	(0.030)	(0.043)	(0.26)	(0.01)	(0.19)
B. Weighted by beds	0.036	-0.079	-0.122	-0.091	-0.165	-0.080	-0.111	-0.77	-0.04	-0.61
	(0.026)	(0.032)	(0.024)	(0.029)	(0.044)	(0.034)	(0.048)	(0.22)	(0.02)	(0.16)
C. Treated group trend	0.067	-0.018	-0.047	-0.061	-0.119	-0.034	-0.062	-0.29	-0.00	-0.27
	(0.037)	(0.030)	(0.031)	(0.030)	(0.058)	(0.047)	(0.066)	(0.33)	(0.01)	(0.25)
D. CS estimator	0.066	-0.017	-0.063	-0.064	-0.146	-0.015	-0.124	-0.52	-0.03	-0.45
	(0.028)	(0.026)	(0.027)	(0.026)	(0.053)	(0.044)	(0.048)	(0.27)	(0.01)	(0.20)
Obs (panels A-D)	16,655	16,655	16,655	20,998	20,997	20,997	20,997	20,998	20,998	20,998
E. Balanced panel	0.071	-0.020	-0.068	-0.059	-0.146	-0.027	-0.096	-0.56	-0.02	-0.51
	(0.029)	(0.029)	(0.029)	(0.031)	(0.046)	(0.034)	(0.048)	(0.28)	(0.01)	(0.21)
Obs	16,099	16,099	16,099	20,522	20,521	20,521	20,521	20,522	20,522	20,522
F. Matched sample	0.043	-0.056	-0.110	-0.027	-0.122	-0.018	-0.085	-0.82	-0.03	-0.69
	(0.031)	(0.028)	(0.028)	(0.029)	(0.051)	(0.040)	(0.049)	(0.31)	(0.01)	(0.23)
Obs	4,175	4,175	4,175	4,904	4,904	4,904	4,904	4,904	4,904	4,904
G. All treated obs	0.028	-0.044	-0.103	-0.076	-0.143	-0.058	-0.148	-0.68	-0.03	-0.56
	(0.026)	(0.022)	(0.023)	(0.032)	(0.048)	(0.035)	(0.043)	(0.25)	(0.01)	(0.18)
Obs	19,503	19,503	19,503	24,844	24,843	24,843	24,843	24,844	24,844	24,844
H. Switchers included	0.061	-0.032	-0.085	-0.083	-0.147	-0.047	-0.138	-0.59	-0.02	-0.53
	(0.026)	(0.024)	(0.024)	(0.027)	(0.042)	(0.030)	(0.043)	(0.26)	(0.01)	(0.20)
Obs	18,988	18,988	18,988	23,620	23,619	23,619	23,619	23,620	23,620	23,620

Notes: The table shows the results of robustness checks for the effects on hospital finances, patient volume, and employees estimated for the privatized hospitals, presented in Tables 3, 4, and 5, respectively. For brevity, we do not present results for outcomes where we do not detect effects, such as non-personnel expenses and nurse employment. For each outcome we present the baseline estimates in Panel A. Panel B includes static hospital beds to weight hospitals. Panel C uses the baseline specification including a linear trend interacted with an indicator for privatized hospitals. Panel D presents coefficients using the Callaway and Sant'Anna (2020) estimator. Panel E drops treated hospitals privatized after 2014 to ensure we observe each privatized hospital for five years before and after the transition. Panel F presents results estimated on a matched subsample, identified using propensity score matching. Panel G uses all treated observations, including those from the year of privatization and those beyond the five-year window around privatization (if available). Panel H includes hospitals as additional control units that are most commonly labeled as public, switch between public and private, but were not captured in our manual validation of conversions. Please see Section 5.5 for additional details. Standard errors are clustered by hospital.

Table 7: Descriptive statistics (market-level)

	(1) Treated HSAs	(2) Control HSAs	(3) Total
# treated hospitals	1.3	0.0	0.3
	(0.6)	(0.0)	(0.6)
Total hospitals	5.4	4.3	4.6
	(5.1)	(6.4)	(6.1)
Total admissions	34,152	30,927	31,650
	(54,842)	(79,635)	(74,780)
Total beds	875	778	800
	(1,313)	(1,902)	(1,786)
% Medicaid adm	15.4	13.9	14.2
	(6.3)	(7.1)	(6.9)
% Medicare adm	45.2	47.1	46.7
	(10.2)	(9.6)	(9.8)
% other adm	39.3	39.0	39.1
	(11.5)	(10.0)	(10.4)
Total FTEs/100 adj adm	7.1	7.2	7.2
	(2.2)	(3.4)	(3.2)
% in poverty	14.1	12.9	13.2
	(5.0)	(4.7)	(4.8)
% unemployment	4.9	4.7	4.8
	(2.3)	(2.4)	(2.4)
% uninsurance	20.6	19.0	19.3
	(6.0)	(5.6)	(5.7)
HHI (admissions)	4,981	5,814	5,627
	(2,607)	(2,930)	(2,881)
# HSAs	204	706	910

Notes: The table presents descriptive statistics for the market-level sample, where markets are defined by Health Service Areas (HSAs). We use values from 1999 for most HSAs. In rare instances in which we do not observe an HSA in 1999, we use values from that HSA's first year in the data. Treated HSAs have at least one hospital that undergoes public to private conversion during 2000–18. Control HSAs do not have any conversions during our sample period. All rows present means and standard deviations (in parentheses).

Table 8: Effects on aggregate patient (log) volume

	(1) Total	(2) Medicaid	(3) Medicare	(4) Other
A: No controls				
DD	-.009	-.042	-.001	.004
	(.014)	(.026)	(.016)	(.022)
Obs	19,404			
B: Market controls				
DD	-.021	-.051	-.016	-.008
	(.014)	(.026)	(.015)	(.022)
Obs	17,983			
C: Heterogeneity by market HHI				
DD	.043	.034	.042	.068
	(.016)	(.024)	(.017)	(.019)
x 1(> med. HHI)	-.105	-.154	-.088	-.128
	(.026)	(.048)	(.029)	(.041)
D: Heterogeneity by market poverty				
DD	.016	.032	.022	.019
	(.019)	(.031)	(.019)	(.027)
x 1(> med. poverty)	-.050	-.148	-.045	-.029
	(.027)	(.048)	(.029)	(.042)
Mean outcome (t-1)	36,550	6,796	15,282	14,473

Notes: The table presents estimated effects on patient volume at the market-level obtained by estimating the market-level equivalent of Equation 2 on market-year level data. We define markets using Health Service Areas (HSAs), as described in Section 6. Columns 1, 2, 3, and 4 present the effects on log total, Medicaid, Medicare, and other admissions, respectively. 'Other' admissions refers to hospital admissions not covered by Medicaid or Medicare and mostly comprises privately insured and uninsured patients. Panel A reports coefficients from a two-way fixed effects specification with no covariates. Panel B reports coefficients from a two-way fixed effects specification including time-varying, HSA-level controls: population, unemployment, uninsurance, and poverty rates. Panel B has fewer observations since the covariates are not available for 1995 and 1996. Panel C presents the corresponding results from a triple difference specification including an interaction term with an indicator for the market having a Herfindahl-Hirschman Index (based on admission shares) in 1999 greater than the median among treated markets. Panel D is analogous to Panel C but instead includes an interaction term with an indicator for the market having a poverty rate in 1999 greater than the median. The mean values pertain to patient volume (in levels) in treated markets in the year prior to privatization. Standard errors are clustered by HSA and are presented in parentheses.

A Additional figures and tables

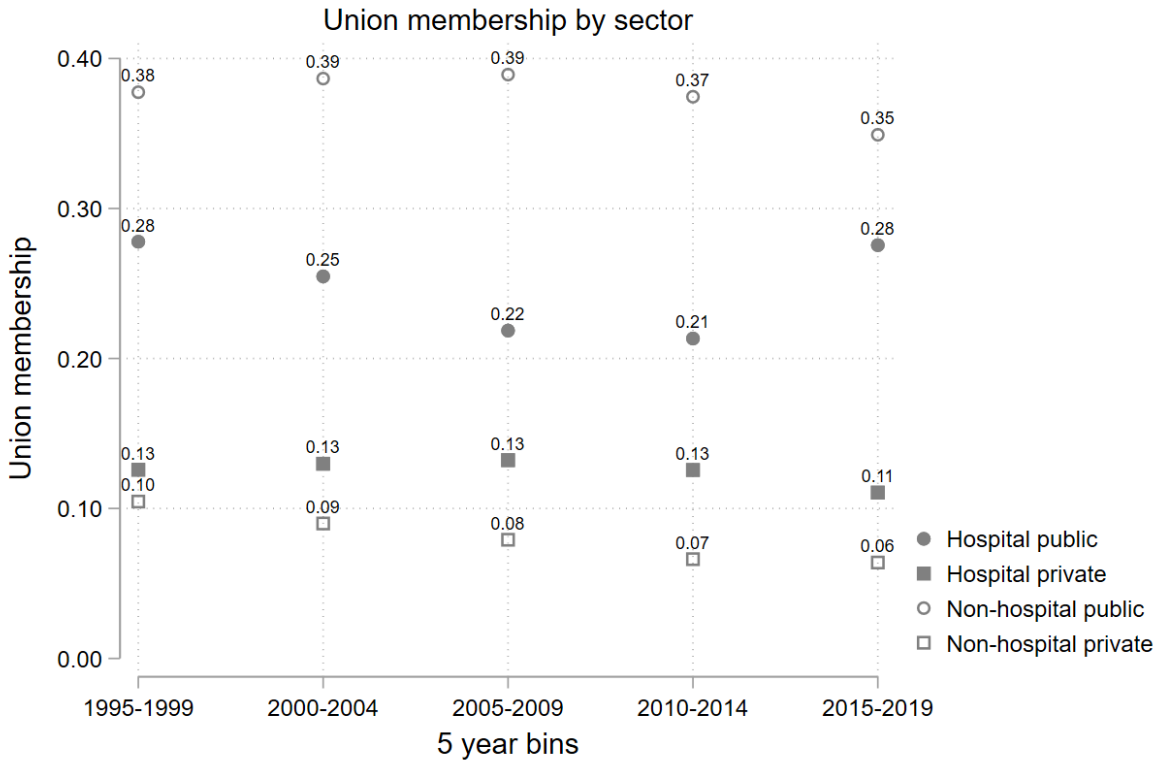


Figure A.1: Unionization among public and private employees

Note: The figure presents the mean share of unionized hospital employees in public and privately owned firms over time. We plot values separately for employees of hospitals and all other firms within each group. The underlying data was sourced from the Annual Social and Economics Supplement of the Current Population Survey (ASC) over 1995–2019. ASC includes individuals who are surveyed in the "earner study," which includes approximately one quarter of the CPS sample. We restricted the sample to include: individuals who are part of the earner study; employed: at work or has a job but not at work; older than 17 and younger than 66. Due to small sample sizes, we pool data from five years to compute means. Individuals are weighted by the corresponding population weights provided in the survey. We identified the hospital industry using NAICS1990 code 831 and NAICS1997 code 622.

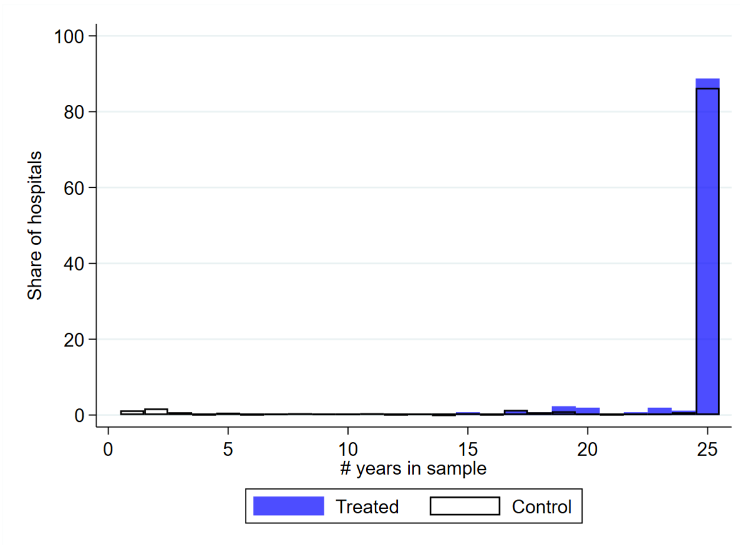


Figure A.2: Balance of hospital panel

Note: The figure presents a frequency distribution of the number of years a hospital is observed in the sample, separately for privatized (treated) and control hospitals. The maximum number of years possible is 25 (1995–2019).

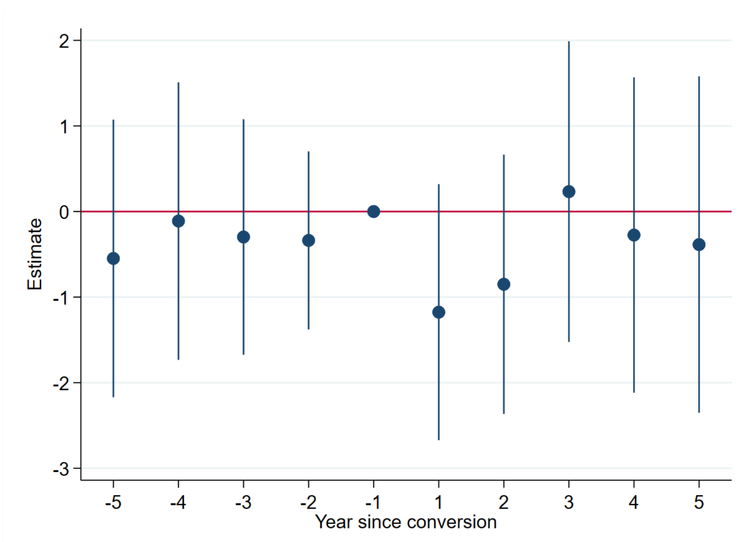


Figure A.3: Effect on total patient volume per bed

Note: The figure presents dynamic effects on total volume per bed obtained by estimating Equation 3 on hospital-year level data. Total patient volume is normalized by contemporaneous hospital beds. Year zero is the year of privatization and is excluded for treated hospitals since it represents partial treatment. The error bars denote 95% confidence intervals. Standard errors are clustered by hospital.

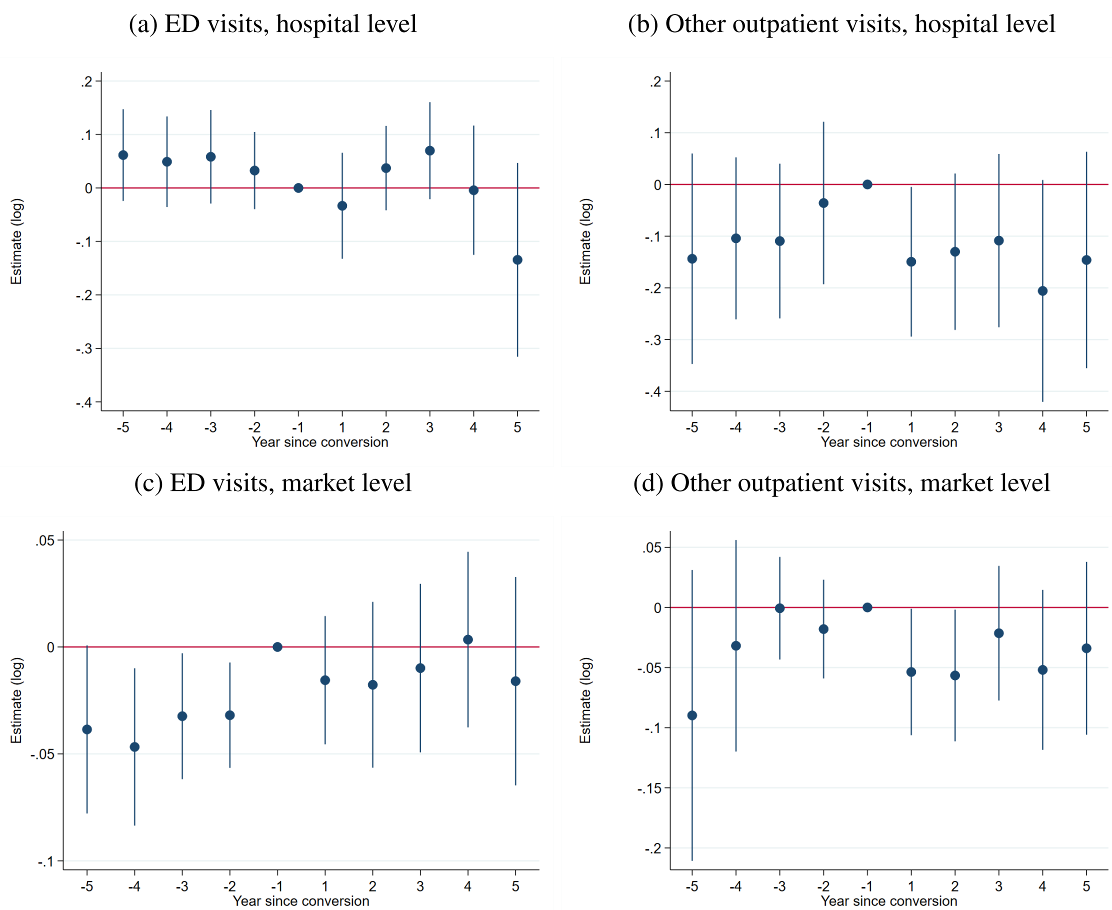


Figure A.4: Effects on ED and other outpatient visits

Note: The figure presents event study plots obtained by estimating Equation 3 on the hospital-year sample (panels (a) and (b)) and market-year sample (panels (c) and (d)). The outcomes are log emergency department (ED) visits and other (non-ED) outpatient visits. Year zero is the year of privatization and is excluded for treated hospitals since it represents partial treatment. The error bars denote 95% confidence intervals. Standard errors are clustered by hospital in panels (a) and (b) and by market in panels (c) and (d).

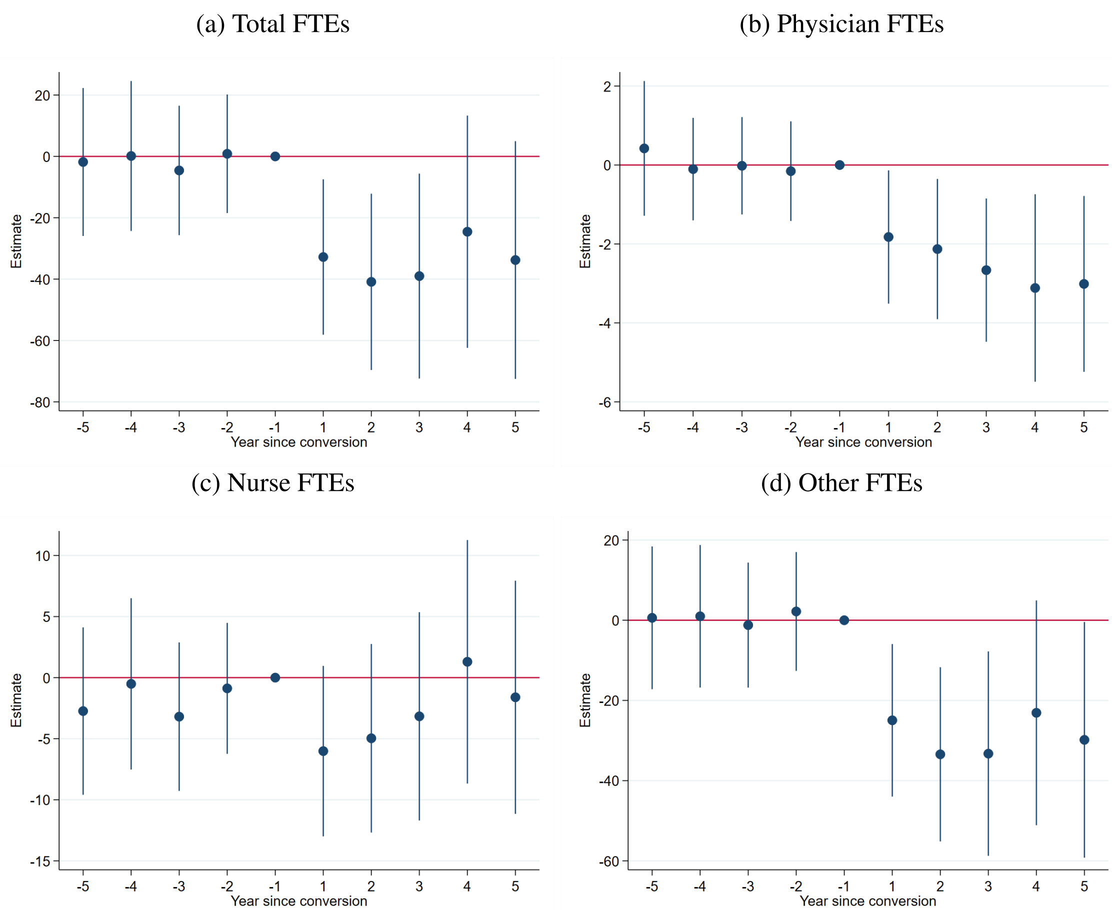


Figure A.5: Effects on staffing (FTE per 100 beds)

Note: The figure presents event study plots obtained by estimating Equation 3 on hospital-year level data. The comparison group is comprised of hospitals that remain public throughout our sample period. Outcomes are total full-time equivalent employees (FTEs), physician FTEs, nurse FTEs, and other FTEs in panels (a), (b), (c), and (d), respectively. All outcomes are normalized by the contemporaneous number of hospital beds and presented per 100 beds. Year zero is the year of privatization and is excluded for the treated hospitals since it represents partial treatment. The error bars present 95% confidence intervals. Standard errors are clustered by hospital.

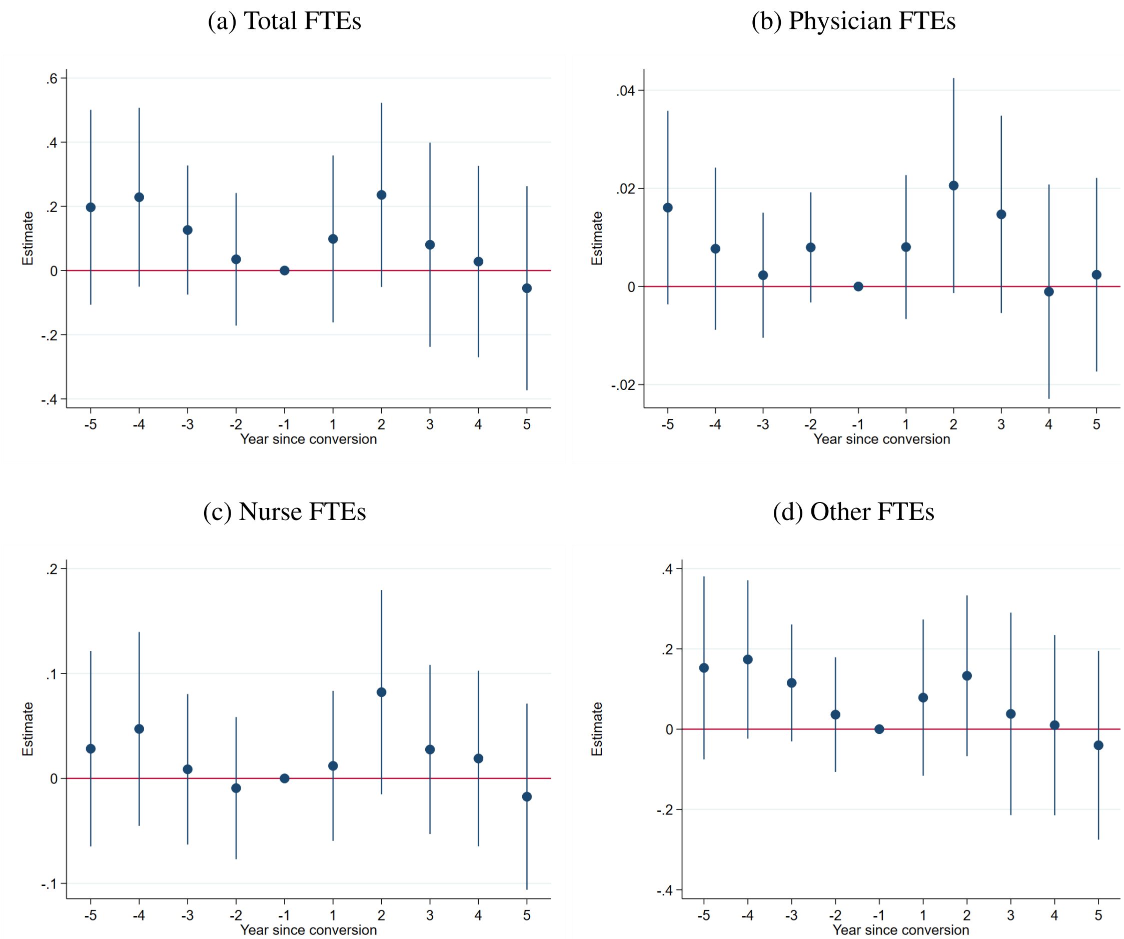


Figure A.6: Effects on market staff (per 100 adjusted admissions)

Note: The figure presents event study plots obtained by estimating the market-level equivalent of Equation 3 on market-year level data. We define hospital markets using Health Service Areas (HSAs), as described in Section 6. The outcomes are as indicated in the figure and are normalized by contemporaneous, adjusted admissions. We then multiply outcomes by 100 for ease of exposition. Year zero is the first year a privatization occurs in a given market and is excluded for treated markets since it represents partial treatment. The error bars present 95% confidence intervals. Standard errors are clustered by HSA.

Table A.1: Public (non-federal) hospital share of beds by state in 2019

State	Share	# Hospitals	State	Share	# Hospitals
Wyoming	70.8	32	Nevada	14.1	58
Alabama	44.4	116	Kentucky	13.7	121
Mississippi	40.7	112	Nebraska	13.5	99
Kansas	36.8	152	New Jersey	12.9	99
South Carolina	32.9	88	Georgia	11.7	172
North Carolina	31.8	135	Ohio	11.3	224
Iowa	29.8	123	Arkansas	10.4	102
Washington	27.0	107	Rhode Island	10.3	15
Louisiana	26.1	200	Montana	10.1	66
Idaho	25.2	52	Connecticut	9.9	42
New York	23.6	210	West Virginia	9.3	61
Colorado	23.5	106	Maryland	8.5	62
California	22.9	419	Massachusetts	8.2	102
New Mexico	22.2	55	Illinois	8.0	208
Hawaii	22.1	28	District Of Columbia	7.4	14
Virginia	20.1	123	Delaware	6.3	13
Oregon	19.8	65	Wisconsin	6.3	149
Oklahoma	19.4	146	Arizona	6.2	110
Tennessee	19.0	132	Michigan	6.2	165
Utah	18.6	59	New Hampshire	5.5	31
Missouri	18.2	143	Maine	5.4	39
Indiana	17.5	161	South Dakota	4.4	64
Florida	16.8	253	Pennsylvania	3.8	235
Texas	15.8	588	North Dakota	2.6	50
Alaska	14.6	26	Vermont	1.7	17
Minnesota	14.4	141			

Notes: The table presents public (non-federal) shares of hospital beds for all 50 states and DC using American Hospital Association survey data from 2019. All hospitals, including non-general-acute-care hospitals, were included in share calculations. States are listed in decreasing order of public shares. The total number of hospitals for each state is given in the third column.

Table A.2: Types of privatization deals

	(1) Non-profit	(2) For-profit	(3) Total
A. Less control	119	25	144
- Contract Management	70	10	80
- Miscellaneous	49	15	64
B. More control	66	48	114
- Sale	36	33	69
- Lease/Joint venture	30	15	45
Total	185	73	258

Notes: This table presents a breakdown of the privatization deals in our main analysis sample. These occur between 2000-2018. Columns 1 and 2 present the number of hospitals that converted to private non-profit and for-profit, respectively. Panel A lists the modes that permit the private firm less control over hospital operations. In contract management, the private firm operates the hospital under a short-term contract. "Miscellaneous" includes cases where a new private firm was incorporated–subject to oversight by the previous government owners–specifically to operate the hospital and cases where the modality could not be identified. Panel B lists modes of transfer that allowed the private firm more control over hospital operations. These include sale, lease, and joint ventures. Appendix B.1 describes these categories in more detail with examples.

Table A.3: Effects on ED and other outpatient (log) volume

	(1)	(2)	(3)	(4)
	Hospital		Market	
	ED	Other Outpt	ED	Other Outpt
A: No controls				
DD	-0.048	-0.069	0.018	-0.016
	(0.032)	(0.063)	(0.016)	(0.028)
Obs	20,998	20,998	19,404	19,404
B: Market controls				
DD	-0.043	-0.067	0.016	-0.010
	(0.032)	(0.063)	(0.016)	(0.028)
Obs	19,385	19,385	17,983	17,983
Mean outcome (t-1)	15,424	53,766	136,340	455,672

Notes: The table presents estimated effects on log emergency department (ED) and non-ED outpatient volume at the privatized hospital (cols. 1 and 2) and on the market experiencing privatization (cols. 3 and 4). Panels A and B presents coefficients obtained by estimating Equation 2 without and with time-varying covariates, respectively. The mean values pertain to outcomes (in levels) at treated hospitals or markets in the year prior to privatization. Standard errors are clustered by hospital or market, depending on the level of treatment.

Table A.4: Effects on Traditional Medicare patient volume

	(1) All	(2) Duals	(3) Non duals
A: Full sample			
1: Baseline	-0.083	-0.077	-0.091
	(0.030)	(0.035)	(0.029)
2: C-S	-0.042	-0.056	-0.038
	(0.031)	(0.034)	(0.033)
Mean	6.09	4.82	5.69
Observations	13,824	13,824	13,824
B: Matched sample			
1: Baseline	-0.023	-0.028	-0.028
	(0.038)	(0.044)	(0.037)
2: C-S	-0.002	-0.023	0.004
	(0.035)	(0.039)	(0.037)
Mean	6.20	4.96	5.79
Observations	3,893	3,893	3,893

Notes: This table presents effects on Traditional Medicare (TM) patient volume at the privatized hospitals, estimated using 100% Medicare fee-for-service inpatient claims data over 2000–17. The outcomes are logs of total TM volume, dual eligibles, and non-duals. Panels A and B present results on the full and matched samples, respectively. In each panel, rows 1 and 2 present results from the baseline two-way fixed effects and Callaway and Sant'Anna (2020) models, respectively. These models have fewer observations since the claims data spans a shorter period than the American Hospital Association sample used in the main analysis. To ensure we have two years before and after every privatization, we limit treated units to hospitals privatized during 2002–15. Hence, these models include 215 privatized hospitals instead of the 258 used in the main analysis.

Table A.5: Effects on staff (per 100 beds)

	(1) Total	(2) Physician	(3) Nurse	(4) Other	(5) Contract
A: No controls					
DD	-33.4	-2.5	-1.7	-29.5	0.16
	(12.7)	(0.8)	(3.2)	(9.5)	(1.35)
Obs	20,998				8,632
	.				
B: Market controls					
DD	-33.3	-2.6	-1.9	-29.1	0.12
	(12.7)	(0.8)	(3.2)	(9.5)	(1.35)
Obs	19,385				8,628
Mean outcome (t-1)	511.2	10.2	138.6	362.1	13.60

Notes: The table presents effects on personnel expenses and full-time equivalent (FTE) employed staff at the privatized hospitals, obtained by estimating Equation 2 on hospital-year level data. Column 1 presents results for total FTEs, which comprises of physician, nurse, and other staff, presented in columns 2, 3, and 4, respectively. We normalize staff FTEs in each column by 100 contemporaneous hospital beds, which is approximately the size of a public hospital in our sample. Column 5 presents results for contract FTEs, which come from Medicare cost reports and include management and patient care staff. Panel A reports coefficients from a two-way fixed effects specification with no covariates. Panel B reports coefficients from a two-way fixed effects specification including time-varying, county-level controls as described in Section 4. Panel B has fewer observations since the market-level covariates are not available for 1995 and 1996. The mean values pertain to outcomes at privatized hospitals in the year prior to privatization. Standard errors are clustered by hospital and presented in parentheses.

Table A.6: Heterogeneity by type of privatization

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
		Finances			Volume				Staff	
	Revenue	Tot. Exp.	Pers. Exp.	Total	Medicaid	Medicare	Other	Total	Physician	Other
A: Baseline										
DD	0.057	-0.033	-0.086	-0.084	-0.149	-0.049	-0.138	-0.57	-0.03	-0.52
	(0.026)	(0.024)	(0.024)	(0.027)	(0.042)	(0.030)	(0.043)	(0.26)	(0.01)	(0.19)
B: Heterogeneity by extent of private control										
DD	0.058	-0.016	-0.060	-0.104	-0.184	-0.047	-0.168	-0.37	-0.02	-0.35
	(0.033)	(0.030)	(0.030)	(0.035)	(0.054)	(0.040)	(0.053)	(0.32)	(0.01)	(0.24)
x 1(more private control)	-0.002	-0.038	-0.059	0.046	0.083	-0.003	0.070	-0.45	-0.01	-0.38
	(0.051)	(0.047)	(0.047)	(0.054)	(0.082)	(0.058)	(0.088)	(0.51)	(0.02)	(0.38)
C: Heterogeneity by for-profit conversion										
DD	0.046	-0.037	-0.076	-0.146	-0.186	-0.084	-0.210	-0.67	-0.03	-0.58
	(0.031)	(0.029)	(0.030)	(0.032)	(0.051)	(0.036)	(0.048)	(0.32)	(0.01)	(0.24)
x 1(for-profit)	0.040	0.017	-0.037	0.212	0.128	0.120	0.244	0.34	0.03	0.20
	(0.052)	(0.048)	(0.045)	(0.057)	(0.084)	(0.060)	(0.098)	(0.48)	(0.02)	(0.36)
Obs	16,662	16,662	16,662	20,998	20,997	20,997	20,997	20,998	20,998	20,998

Notes: The table presents heterogeneous effects by type of hospital privatization on hospital finances, patient volume, and employees estimated for the privatized hospitals. For brevity, we do not present results for outcomes where we do not detect effects, such as non-personnel expenses and nurse employment. Panel A reproduces our baseline estimates from Tables 3, 4, and 5. Panel B presents coefficients from a triple differences specification, in which the post-deal indicator is interacted with an indicator for deals in which the private firm has more operational control over the hospital (see Section 2 for more details). Panel C presents coefficients from a triple differences specification using an indicator for hospitals that converted to for-profit ownership. All models are estimated without covariates. Standard errors are clustered by hospital and presented in parentheses.

Table A.7: Effects on aggregate staff

	(1) Total	(2) Physician	(3) Nurse	(4) Other
A: No controls				
DD	-0.03	0.003	0.01	-0.05
	(0.12)	(0.007)	(0.03)	(0.09)
Obs	19,404			
B: Market controls				
DD	0.01	0.004	0.02	-0.01
	(0.11)	(0.007)	(0.03)	(0.08)
Obs	17,983			
C: Heterogeneity by union membership				
DD	0.13	0.01	0.02	0.09
	(0.16)	(0.01)	(0.04)	(0.12)
x 1(< med. union membership)	-0.33	-0.02	-0.03	-0.29
	(0.18)	(0.01)	(0.05)	(0.13)
Mean outcome (t-1)	6.59	0.13	1.92	4.54

Notes: The table presents effects on full-time equivalent (FTE) employed staff at the market-level obtained by estimating the market-level equivalent of Equation 2 on market-year level data. We define markets using Health Service Areas (HSAs), as described in Section 6. Column 1 presents results for total FTEs, which comprises of physician, nurse, and other staff, presented in columns 2, 3, and 4, respectively. We normalize staff FTEs in each column by 100 contemporaneous, adjusted admissions. Adjusted admissions are admissions scaled by the ratio of outpatient to inpatient revenue. Panel A reports coefficients from a two-way fixed effects specification with no covariates. Panel B reports coefficients from a two-way fixed effects specification including time-varying, HSA-level controls: population, unemployment, uninsurance, and poverty rates. Panel B has fewer observations since the covariates are not available for 1995 and 1996. Panel C presents the corresponding results from a triple difference specification including an interaction term with an indicator for markets located in states with rates of union membership (among private workers) in 1999 less than the median among states with treated markets. The mean values pertain to outcomes in treated markets in the year prior to privatization. Standard errors are clustered by HSA and presented in parentheses.

Table A.8: Robustness checks (market-level)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Volume				Staff			
	Total	Medicaid	Medicare	Other	Total	Physician	Nurse	Other
A. Baseline	-0.009	-0.042	-0.001	0.004	-0.03	0.003	0.01	-0.05
	(0.014)	(0.026)	(0.016)	(0.022)	(0.12)	(0.007)	(0.03)	(0.09)
B. Weighted by beds	0.026	0.0003	0.030	0.051	-0.09	-0.01	-0.02	-0.06
	(0.010)	(0.0185)	(0.012)	(0.014)	(0.07)	(0.01)	(0.03)	(0.05)
C. Treated group trend	-0.037	-0.035	-0.033	-0.051	0.29	0.02	0.08	0.19
	(0.016)	(0.031)	(0.020)	(0.040)	(0.13)	(0.01)	(0.04)	(0.10)
D. CS estimator	-0.005	-0.036	0.001	0.004	0.05	0.01	0.01	0.03
	(0.012)	(0.025)	(0.016)	(0.023)	(0.13)	(0.01)	(0.04)	(0.09)
Obs (panels A-D)	19,403	19,403	19,403	19,403	19,403	19,403	19,403	19,403
E. Balanced panel	-0.002	-0.043	0.006	0.012	-0.02	0.005	0.01	-0.03
	(0.015)	(0.027)	(0.017)	(0.022)	(0.14)	(0.008)	(0.04)	(0.10)
Obs	19,122	19,122	19,122	19,122	19,122	19,122	19,122	19,122
F. Matched sample	0.000	-0.002	0.007	-0.001	0.05	0.01	0.07	-0.03
	(0.017)	(0.029)	(0.018)	(0.028)	(0.21)	(0.01)	(0.04)	(0.18)
Obs	4,120	4,120	4,120	4,120	4,120	4,120	4,120	4,120
G. All treated obs	0.016	-0.035	0.025	0.037	-0.20	-0.004	-0.02	-0.17
	(0.022)	(0.033)	(0.024)	(0.026)	(0.17)	(0.008)	(0.04)	(0.13)
Obs	22,535	22,535	22,535	22,535	22,535	22,535	22,535	22,535
H. Switchers included	-0.004	-0.038	0.008	0.009	0.02	0.002	0.03	-0.004
	(0.014)	(0.024)	(0.016)	(0.022)	(0.09)	(0.007)	(0.03)	(0.070)
Obs	19,989	19,989	19,989	19,989	19,989	19,989	19,989	19,989

Notes: The table shows the results of robustness checks for the effects on market-level patient volume and full-time equivalent (FTE) employed staff per 100 adjusted admissions, presented in Tables 8 and A.7, respectively. It follows the same format and presents the same checks as in Table 6. For each outcome we present the baseline estimates in Panel A. Panel B includes static hospital beds to weight markets. Panel C uses the baseline specification including a linear trend interacted with an indicator for treated markets. Panel D presents coefficients from the Callaway and Sant'Anna (2020) estimator. Panel E drops markets treated after 2014 to ensure we observe each treated market for five years before and after the transition. Panel F presents results estimated using a matched subsample identified using propensity score matching. Panel G uses all treated observations, including those from the year of privatization and those beyond the five-year window around privatization (if available). Panel H includes hospitals (when aggregating to markets) that switch between public to private (or vice versa) but were not captured in our manual validation of conversions. Standard errors are clustered by market and presented in parentheses.

B Data Appendix

B.1 Privatization taxonomy

We first identify cases of public hospitals that were converted to private control during our study period of 2000–18. There is no official source of such events, and thus we utilized the AHA annual survey files over this period. We infer a conversion when we observe a change in management control type from public (state, county, or city) to private (for-profit or non-profit). We validate conversions using information recorded in the annual AHA summary of changes files, which explain each change in the AHA survey from the previous year. This naive approach yielded 354 public to private conversions. Next, we devoted hundreds of hours to manually verify each conversion by combing through hospital websites, news articles, and third-party sites such as the American Hospital Directory. Manual validation help us identify non-trivial numbers of false positive and false negative conversions. Our final number of conversions is 258.

Through these detailed reviews we classified privatizations into five groups, described below. We consider the first two as transitions where the private operator has less control over hospital operations, while the latter three afford greater control. We provide counts for each group in Table A.2. We provide an example for each type to help illustrate the differences across these deals.

- **Contract management:** Occurs when a private (corporation or health system) firm takes over the day-to-day management of a hospital. Government maintains control over the hospital's property, assets, and debts.

  Example: Mercy Hospital Lincoln (Troy, MO) recorded a conversion in the AHA in 2015 from "County" to "other not-for-profit." Manual validation noted that Mercy signed an agreement to lease and manage the facility beginning March 1, 2015.

- **Public hospital incorporating as a private firm:** Occurs when a public health system files for 501c3 nonprofit status ("incorporating").

  Example: Hutchinson Area Health Care (Hutchinson, MN) recorded a conversion in 2008 from "city" to "other not-for-profit." Manual validation noted that in January 2008 Hutchinson Area Health Care became its own private, nonprofit corporation and was no longer a part of the city of Hutchinson.

- **Sale:** Occurs when there is a permanent transfer in the ownership and control of the property, assets, and debts of a hospital, from government to a private corporation or hospital.

  Example: Glenwood Regional Medical Center (West Monroe, LA) recorded a conversion in the AHA in 2006 from "hospital district or authority" to "other not-for-profit." Manual validation noted that IASIS Healthcare LLC announced the signing of a definitive agreement to acquire Glenwood Regional Medical Center from the Hospital Service District for approximately $82.5 million.

- **Long-term lease:** Occurs when a private (corporation or health system) authority takes control over day-to-day management of a hospital for an extended period of time (more than 15 years). The government entity maintains control over the hospital's property, assets, and debts.

  Example: Mercy McCune-Brooks Hospital (Joplin, MO) recorded a conversion in the AHA in 2012 from "city" to "church operated." Manual validation noted that Mercy's 50-year lease of the city-owned hospital was approved by the Carthage City Council in 2012.

- **Joint venture:** Occurs when one or more private (corporations or health systems) firms agree to enter into a joint venture with the local government authority, which results in a newly formed private firm to take over management of the hospital.

Example: Rice Memorial Hospital (Willmar, MN) recorded a conversion in the AHA in 2018 from "city" to "other not-for-profit." Manual validation noted that Rice Memorial Hospital, ACMC Health and CentraCare Health signed the final agreement to establish Carris Health, a subsidiary of CentraCare Health, which is a not-for-profit health care system. Carris Health committed to make a capital investment of $32 million in Rice Memorial Hospital over the next 10 years. The hospital's assets would continue to be owned by the City.

C Methodology

C.1 Sample selection

To construct our analytic sample of control hospitals, we start with American Hospital Association (AHA) survey data for the years 1995 to 2019. In the raw data there are ∼6,200 hospitals per year and ∼8,400 unique hospitals over the sample period. We make the following sample restrictions:

- Drop hospitals whose most common AHA service code is not "general medical and surgical" (2,457 hospitals)
- Drop hospitals that on average report fewer than 10 beds (42 hospitals)
- Drop hospitals that are ever classified as federal government by the AHA (293 hospitals). These include military, Veterans Affairs, Indian Health Service, and Department of Justice hospitals
- Drop hospitals that are only classified as public (state and local) in some years of the sample period but not all. This group includes hospitals that are most commonly labeled as private (290 hospitals) and hospitals that are most commonly labeled as public (122 hospitals). This is a conservative restriction to ensure that our comparison group is comprised of non-converting, public hospitals
- Drop hospitals that are within 15 miles of at least one treated hospital (32 hospitals)

Our final, analytic sample consists of 802 control hospitals.

As discussed previously, we created our list of public to private conversions by starting with conversions implied by changes in the AHA's control variable and then manually validating each conversion. From this process we identified 269 total conversions. From our manual validation, we found that two treated hospitals experience more than two conversions (i.e. public to private or private to public) over our sample period; we dropped these hospitals. Five hospitals were dropped that convert from private to public and back to private within our sample period. Finally, we dropped four treated hospitals whose most common AHA primary service code was not "general medical and surgical." Our final set of treated hospitals consists of 258 public to private conversions. We note that two treated hospitals experience a second conversion in which they convert back from private to public. For these two hospitals we drop observations on or after the second conversion.

C.2 Propensity score matching

In one of our robustness checks reported in Table 6, we apply propensity score matching (PSM) to our analytic sample to identify treated and control hospitals that are similar on pre-period observables. Specifically, we conduct one-to-one, nearest neighbor matching without replacement and estimate logit models to predict privatization with the following explanatory variables from t-1 to t-3 (where t denotes the year of privatization for a given treated hospital):

- # hospital beds
- Total admissions
- Medicaid admissions
- Total expenses
- % in poverty (measured at the county-year level)
- % unemployment (measured at the county-year level)
- Health Service Area population (only t-1; calculated by aggregating county-year population estimates)

We impose the restriction that propensity scores of matched pairs be in the same decile of the propensity score distribution. We apply PSM sequentially by first searching for similar control hospitals for hospitals that privatize in 2000, the first year of conversions in our data. Control hospitals that match to these privatizing hospitals are removed from the donor (control hospital) pool prior to searching for matches for hospitals that privatize in 2001. We continue this process for all 19 years of privatizations (2000–2018) and are able to match all 258 treated hospitals.

We also apply PSM to our market-level (HSA) sample using an analogous approach. The only difference is that we match on the total number of hospitals in the market from t-1 to t-3, rather than the total number of hospital beds.