

# Owner Incentives and Performance in Healthcare: Private Equity Investment in Nursing Homes

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Amid an aging population and a growing role for private equity (PE) in the care of older adults, this paper studies how PE ownership affects U.S. nursing homes using patient-level Medicare data. We show that PE ownership leads to a patient cohort with lower health risk. However, after instrumenting for the patient-nursing home match, we find that PE ownership increases mortality by 11%. Declines in measures of patient well-being, nurse staffing, and compliance with care standards help to explain the mortality effect. Overall, we conclude that PE has nuanced effects with adverse outcomes for a subset of patients. (*JEL* G14, G23, G28, I11)

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The U.S. population, like that of many advanced economies, is aging rapidly. This has created increasing demand for the care of older adults. Private equity (PE)-owned firms are playing a growing role in meeting this need. In this paper, we examine how PE ownership of nursing homes affects patients and taxpayers. Relative to independent private firms, PE ownership brings short-term, high-powered incentives to maximize profits. Existing literature and the policy debate provide opposing predictions.

On the one hand, evidence indicates that for-profit healthcare firms can maintain long-term implicit contracts with stakeholders (Duggan 2000; Adelino, Lewellen, and Sundaram 2015). Voices from the private sector suggest this may apply to PE; for example, a 2019 report from consulting firm EY concluded that “Not only is PE perceived to have a beneficial overall impact on health care businesses, it is also considered to positively influence the focus on quality and clinical services” (EY 2019). Finally, PE has been found to have positive effects in other industries (Kaplan 1989; Kaplan and Weisbach 1992; Davis et al. 2014; Bloom et al. 2015; Bernstein and Sheen 2016; Hochberg and Rauh 2013).

On the other hand, nursing home customers are particularly vulnerable and face severe information frictions (Carlin, Umar, and Yi 2020). Consumers may make choices that are not in their best interest if financial literacy is lacking or their decision-making is impaired by cognitive decline, for example (Carlin and Robinson 2012). This could lead to different outcomes than in other parts of healthcare and the economy more broadly. Theories of firm behavior suggest that information frictions and noncontractible quality can weaken the natural ability of a market to align firm incentives with consumer welfare (Arrow 1963; Hansmann 1980; Hart, Shleifer, and Vishny 1997). In 2019, U.S. Senators asked about “the role of PE firms in the nursing home care industry, and the extent to which these firms’ emphasis on profits and short-term return is responsible for declines in quality of care.”<sup>1</sup>

In this paper, we present the first national study on the causal effects of PE ownership of nursing homes and relax key limitations of the prior literature. Existing studies on the role of PE in healthcare have faced challenges of limited geographies, a short sample period, a lack of patient-level data, or rely on a small number of deals (Grabowski and Stevenson 2008; Harrington et al. 2012; Pradhan et al. 2013; Bos and Harrington 2017; Huang and Bowblis 2019). Our paper is the first to employ a national sample of PE acquisitions spanning nearly two decades, to address both patient and (partially) facility-level selection, and to study mortality, an unambiguous measure of patient welfare.

Within healthcare, nursing homes represent an extreme example of reliance on subsidy and are characterized by severe market frictions. First, the average

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<sup>1</sup> The letter from Sherrod Brown, Elizabeth Warren, and Mark Pocan is available online (U.S. Congress 2019).

nursing home receives 75% of its revenue from the government. Second, patients are especially vulnerable and exhibit a strong tendency to go to the closest facility (Grabowski et al. 2013). The sector is also independently important, with spending at \$166 billion in 2017 and projected to grow to \$240 billion by 2025 (Martin et al. 2018). PE firms have acquired both large chains and independent facilities, making it possible to isolate the effects of PE ownership from corporatization (Eliason et al. 2020).

We use patient- and facility-level administrative data from the Centers for Medicare & Medicaid Services (CMS), which we match to PE deal data. The data include about 12,400 unique for-profit nursing homes between 2000 and 2017. Of these, 1,674 were acquired by PE firms in 128 unique deals. Our analysis sample contains 4.2 million unique short-stay patients. We focus on Medicare, which accounts for about 60% of the unique patients that enter a nursing home during our sample period.

There are two empirical challenges to estimating the causal effects of PE ownership. The first is nonrandom selection of acquisition targets. We partially address this by including facility fixed effects in estimation, which eliminates time invariant differences across facilities and their local markets. We also include patient market-by-year fixed effects to mitigate concerns about unobserved differential trends in market structure across locations. Finally, we present event studies and assess pre-trends for all outcomes. The results point to causal effects on treated firms, though these are not necessarily externally valid to a random firm in the economy.<sup>2</sup> This interpretation is nonetheless important for social welfare as private equity has a significant and growing footprint.

The second challenge is that the patient composition changes after PE buyouts. We find that patient risk declines, which could reflect an effort to pursue more financially attractive patients. While Medicare compensates nursing homes for the higher costs of serving more complex patients by adjusting payments, these adjustments account for only a fraction of the variance in costs (White, Pizer, and White 2002). Medicare also rewards physical therapy, which favors healthier patients (Carter, Garrett, and Wissoker 2012). Following PE buyouts, we find declines in measures associated with costly care, such as cognitive impairments and inability to perform activities of daily living (Hackmann, Pohl, and Ziebarth 2021). To address potential unobserved selection, we control for the patient-facility match with a differential distance instrumental variables (IV) strategy (McClellan, McNeil, and Newhouse 1994; Grabowski et al. 2013; Card, Fenizia, and Silver 2023), exploiting patient preference for a nursing facility close to their home (the median distance is 4.8 miles). The distance-based instrument strongly predicts facility choice and is uncorrelated with observed patient risk. It controls for selection within the

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<sup>2</sup> As we will show below, prior to the buyout, PE target facilities were larger, located in urban markets, served more patients per bed, and had a more lucrative payer mix.

population of patients who go to a PE-owned facility because it is closer to their home.

We use both ordinary least squares (OLS) and IV differences-in-differences models to examine the effects of PE buyouts on patient welfare. The most important and objective measure in our context is short-term survival, which we define as the probability of death during the stay and the following 90 days (McClellan and Staiger 1999; Hull 2018). In OLS models, we show that PE ownership leads to a 0.3-pp increase in mortality, about 2% of the mean. The IV approach finds that going to a PE-owned nursing home has an increase in mortality of 11% of the mean. This effect is detectable as early as 15 days following discharge and the magnitude is stable out to 365 days.

We take a step toward assessing the external validity of the IV results using a marginal treatment effects (MTE) analysis. Unlike the LATE, the MTE analysis estimates parameters that are not specific to the complier group and allows us to make more general statements regarding the causal effects of PE ownership within the treated sample of facilities. The MTE analysis recovers an average treatment effect similar to the LATE. This implies that the average Medicare patient in our sample would also experience an 11% increase in the chance of short-term mortality if she goes to a PE-owned nursing home. The MTE analysis reveals substantial heterogeneity in treatment effects, including small beneficial effects for some patients.

We assess whether the results are driven by the related but distinct phenomenon of corporatization. The coefficients remain intact when we restrict our attention to PE acquisitions of the largest chains, in which chain size remained constant over the sample period, implying that the effect captures the nature of ownership rather than consolidation or corporatization. We also conduct standard robustness tests, including a placebo analysis, where we show there are no prebuyout effects. Together with the absence of pre-trends in event studies, this suggests that the results do not reflect the targeted facilities being on track to experience these effects regardless of the buyout.

We examine three channels to explain and corroborate the effects on mortality. The first is nurse availability, which is the most important determinant of quality of care (Zhang and Grabowski 2004; Lin 2014). PE ownership leads to a 3% decline in hours per patient-day supplied by the frontline nursing assistants who provide the vast majority of caregiving hours and perform crucial well-being services such as mobility assistance, personal interaction, and cleaning to minimize infection risk. We also find that relatively lower-risk patients drive the negative average effects on mortality, which may reflect less availability of frontline nurses. PE-owned nursing homes also keep low-risk patients longer, which would maximize Medicare revenue but may be worse for the patient. Among high-risk cohorts, PE-owned nursing homes appear to maintain quality, as they increase the number of RNs, who are responsible for the most medicalized aspects of treatment.

The second channel is facility Five-Star ratings, which are constructed by CMS to provide summary information on quality of care. We find negative effects on these ratings. A disconnect between demand and quality of care may reflect information frictions in nursing home quality transparency. Existing work finds weak or no demand response to information about nursing home care quality, including Five-Star ratings (Grabowski and Town 2011; Werner et al. 2012). Consistent with an important role for subsidies, which separate revenue from the consumer, as a mechanism for the negative effects, we find that quality declines are driven by nursing homes with above-median Medicare funding as a share of total revenue.

If PE ownership affects mortality by leading to a lower quality of care, we expect negative effects on measures of patient well-being. To investigate this third channel, we consider three measures of patient well-being that are key standards for CMS. In OLS (IV) models, we find a decrease in mobility of 6.2% (3%), increase in ulcer development of 8.5% (0%), and increase in pain intensity of 10.5% (8.3%). Event studies indicate no pre-trends and show discontinuous changes after the buyouts. This third channel corroborates the effect on mortality, even though there are differences between the OLS and IV models.

Taken together, our results indicate nuanced effects of PE ownership. Patients become less risky after PE buyouts and thus it is unsurprising to see a smaller effect on mortality in OLS relative to IV analysis. The baseline OLS results and MTE analysis show that for some patients, there is no appreciable effect on mortality. However, we do find a large increase in mortality on average. Overall, it seems likely that PE ownership either does not affect or benefits more sophisticated patients, but adversely affects those who face more information frictions, since we find the greatest effect for patients most likely to go to a PE facility due to distance.

Finally, to understand implications for the taxpayer and to shed light on how PE firms create value, we explore changing financial strategies. Consistent with patients being on average lower risk, OLS models find small declines after buyouts in the amount billed to Medicare per stay (note profits may increase if caring for these patients is less costly). In contrast, the IV estimate is in the opposite direction, indicating an 8% increase in the amount billed. Facility finances shed light on why nursing homes are attractive targets for PE buyouts given their low and regulated profit margins, often cited at just 1-2%. Using CMS cost reports, we find that there is no effect of buyouts on net income, which points to strategies maximizing longer term profitability. Three types of expenditures are associated with PE profits and tax strategies: “monitoring fees” charged to portfolio companies, lease payments after real estate is sold to generate cash flows, and interest payments reflecting the importance of leverage in the PE business model (Metrick and Yasuda 2010; Phalippou, Rauch, and UMBER 2018). We show that all three increase after buyouts, with interest payments rising by over 200%. Finally, the negative

effects on quality-of-care measures are driven by facilities with higher levels of financial liabilities and by those acquired by healthcare-focused PE funds, consistent with a role for PE's unique operational model in explaining the changes in quality. While many aspects of facility finances, including labor costs and overall revenue, are either sparsely populated or ambiguously documented in the cost reports, the elements that we can analyze point to changing financial strategies that could enable attractive returns for the PE fund without increasing reported net income of the facility.

In terms of policy implications, our results suggest that, in partial equilibrium, restricting PE transactions would save lives. However, several important caveats imply a need for further study. First, it is possible that in the longer term, restricting acquisitions could affect the incentives of providers to create new facilities, which could affect long term health outcomes. Second, our empirical strategy does not fully address nonrandom targeting of facilities. For example, some facilities not currently being acquired by PE funds could benefit from such an acquisition. Finally, a large literature shows that PE firms increase efficiency in terms of profit maximization. If payments were designed to better align incentives between firm owners and patients (as well as taxpayers), it seems likely that PE-owned nursing homes would behave differently, leading to better outcomes for patients.

This paper contributes to multiple strands of the literature. Most broadly, our results imply that high-powered incentives to maximize profits are not unambiguously beneficial in contexts with market frictions and government subsidy, which may be helpful for policy makers considering actions to improve transparency and accountability (Rose-Ackerman 1996; Picone, Chou, and Sloan 2002; Bénabou and Tirole 2006; Curto et al. 2019). In this way, we expand on the literature describing how PE ownership affects target firm operations (Boucly, Sraer, and Thesmar 2011), product quality (Lerner, Sorensen, and Strömberg 2011; Eaton, Howell, and Yannelis 2020; Fracassi, Previtro, and Sheen 2022), and value (Bernstein, Lerner, and Mezzanotti 2019; Biesinger, Bircan, and Ljungqvist 2020; Gupta and Van Nieuwerburgh 2021).

We also contribute to the healthcare economics literature, including how firm ownership interacts with price incentives and regulation in healthcare (Dafny, Duggan, and Ramanarayanan 2012; Ho and Pakes 2014; Eliason, Grieco, McDevitt, and Roberts 2018; Ho and Lee 2019; Curto, Einav, Levin, and Bhattacharya 2021).<sup>3</sup> Our paper joins others within the literature on nursing homes, which grow more economically important as a population ages (Grabowski, Gruber, and Angelelli 2008; Grabowski et al. 2013; Lin 2015; Hackmann 2019; Hackmann, Pohl, and Ziebarth 2021). It is also related to Liu (2021), who shows how PE ownership of hospitals affects

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<sup>3</sup> See also Grabowski and Hirth (2003), Jones, Propper, and Smith (2017), Hill, Slusky, and Ginther (2019), Kunz et al. (2020), and Capps, Carlton, and David (2020).

price negotiations with insurance companies. Our results imply that owner incentives are of first-order importance, pointing to possible benefits from government reimbursements that target patient outcomes.

To our knowledge, [Stevenson and Grabowski \(2008\)](#) were the first to study PE acquisitions in healthcare using survey data. They find little correlation between PE ownership and quality changes. A number of subsequent studies focus on case studies of PE acquisitions in health care, including [Bos and Harrington \(2017\)](#). [Gondi and Song \(2019\)](#) provide a summary of issues related to PE acquisitions of healthcare facilities. [Harrington et al. \(2012\)](#) study nurse staffing following PE acquisitions among the largest national for-profit chains. They find that PE-owned facilities have higher staffing deficiencies, but that this does not change following buyouts. [Pradhan et al. \(2013\)](#) and [Pradhan et al. \(2014\)](#) use survey data to study five acquisitions in Florida between 2000 and 2007, and explore financial performance, staffing and quality. They find higher operating margins and lower staffing levels. [Cadigan et al. \(2015\)](#) study how investor acquisition of nursing homes affects revenues and costs, and find negligible effects. [Casalino \(2020\)](#) documents that PE has increasingly acquired obstetrician-gynecological medical groups.

Two closely related studies are [Huang and Bowblis \(2019\)](#) and [Gandhi, Song, and Upadrashta \(2020b\)](#). [Huang and Bowblis \(2019\)](#) study PE acquisitions of five nursing home chains in Ohio, focusing on the health status of long-term stay residents between 2005 and 2010. They use a distance-based IV design similar to the one used in our paper to address patient selection. They find little evidence of quality declines but do not explore patient mortality. [Gandhi, Song, and Upadrashta \(2020b\)](#) study how market structure affects the impact of PE acquisitions in the nursing home sector. They find that PE has positive effects on nurse availability in more competitive markets, but negative effects in concentrated markets.

Relative to these studies we make three contributions. First, we comprehensively examine the effect on patient mortality using a national sample of PE acquisitions, demonstrating the importance of accounting for patient composition changes in this setting. We also show that there is considerable heterogeneity in the mortality effect across different types of facilities and patients, which may help guide future studies in this area. Second, we identify channels that help explain the effects on patient health, such as reductions in nurse availability and adherence to standards following PE ownership. Third, we test and confirm the link between these channels and specific aspects of PE ownership, such as specialization.

The economics of nursing homes garnered national attention when the COVID-19 pandemic exposed systemic flaws at long-term care facilities, which accounted for approximately 20% of U.S. deaths from the virus.<sup>4</sup>

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<sup>4</sup> See [Chidambaram and Burns \(2022\)](#).

Braun et al. (2020) and Gandhi, Song, and Upadrashta (2020a) find that PE-owned facilities fared as well or better under the COVID-19 pandemic. Some papers more generally tackle the challenges at nursing homes during COVID-19 (Shen et al. 2022). We do not study performance during COVID-19 for two reasons. First, we do not have Medicare claims data during this period, and second, it is difficult to control for PE selecting homes that would subsequently experience systematically different COVID-19 intensities.

## 1. Institutional Background

### 1.1 The economics of nursing homes

Nursing homes provide both short-term rehabilitative stays, usually following a hospital procedure, and long-term custodial stays for patients unable to live independently. There are two unique features of the long-term care market in the United States relative to other healthcare subsectors. First, government payers (Medicaid and Medicare) account for 75% of revenue, while private insurance plays a much larger role in other subsectors (Johnson 2016).<sup>5</sup> Second, about 70% of nursing homes are for-profit, which is a much larger share than other subsectors. For example, fewer than one-third of hospitals are for-profit. Policy makers have long been concerned about low-quality care at nursing homes in the United States, and for-profit ownership has often been proposed as a causal factor (Institute of Medicine 1986; Grabowski et al. 2013).<sup>6</sup>

As with any business, the economics of nursing homes are shaped by the nature of demand, the cost structure, and the regulatory environment. On the demand side, nursing homes serve older adult patients but are paid by third-party, largely, government payers. Over 95% of facilities treat both Medicare and Medicaid patients (Harrington et al. 2018). Both programs pay a prospectively set amount per day of care for each covered patient (“per diem”), which does not incorporate quality of care, reputation, or other determinants that would be considered by a well-functioning market. These rates are nonnegotiable, and facilities simply choose whether they will accept the beneficiaries of these programs. Medicare fee-for-service pays much more, at roughly \$515 per patient-day relative to \$209 per patient day from Medicaid.<sup>7</sup> Medicaid still pays more than the marginal cost of treatment per

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<sup>5</sup> Medicare is an entitlement health insurance program for Americans above age 65. It covers short-term rehab care following hospital inpatient care, and accounts for about 60% of the unique patients that enter a nursing home, and 15% of overall patient-days in our data. Medicaid is a means-tested insurance program targeted at low-income and nonelderly individuals with disabilities, accounting for about 60% of nursing home patient-days.

<sup>6</sup> This concern is frequently reflected in the popular media, including as a reason for high death rates from COVID-19 in nursing homes. For example, a *New York Times* article in December 2020 asserted that: “Long-term care continues to be understaffed, poorly regulated and vulnerable to predation by for-profit conglomerates and private-equity firms. The nursing aides who provide the bulk of bedside assistance still earn poverty wages, and lockdown policies have forced patients into dangerous solitude” (Kim 2020).

<sup>7</sup> See Spanko (2019).



day. [Hackmann \(2019\)](#) calculates that the marginal cost of treatment per-day is about \$160 on average. Overall profit margins are in the low single digits ([MedPAC 2017](#)), a topic we return to at the end of the paper.

Nursing homes provide institutional care and so have high fixed costs, making the occupancy rate an important driver of profitability. Nursing staff represent the largest component of operating costs, at about 50% of the budget ([Dummit 2002](#)). Broadly speaking, there are three types of nurses. Entry-level Certified Nurse Assistants (CNAs) account for about 60% of total staff hours and provide most of the direct patient care. Licensed Practical Nurses (LPNs) have more training and experience than CNAs but cannot manage patients independently. Registered Nurses (RNs) have the highest skill and experience levels, and can independently determine care plans for patients. LPNs and RNs each account for about 20% of nurse hours. Nurse availability is crucial to the quality of care, and there is a consensus that a low ratio of nursing staff to residents is associated with higher patient mortality and other adverse clinical outcomes ([Tong 2011](#); [Lin 2014](#); [Friedrich and Hackmann 2021](#)). Staffing ratios are therefore a standard metric used to examine nursing home quality.

There is information asymmetry between patients and healthcare providers ([McGuire 2000](#)). As comparing nursing homes on quality is difficult and price is not a helpful signal, reputation may play a large role in nursing home demand ([Arrow 1963](#)). Profit-maximizing facilities might invest in high-quality care to build and sustain their reputation, yet face a dynamic incentive problem because they can generate higher profits in the short-term by cutting patient care costs. It is unclear which inputs affect nursing home reputation, but prior studies suggest that patient demand does not respond to poor quality scores on government mandated report cards, potentially leaving short-term incentives to prevail ([Grabowski and Town 2011](#); [Werner et al. 2012](#)).

## 1.2 The economics of private equity control

PE ownership has different financial incentives and business strategies than other types of for-profit ownership, such as independent or publicly traded firms. Compared to preexisting for-profit owners, private equity owners have higher-powered incentives to maximize firm value because fund managers are compensated through a call option-like share of the profits, employ substantial amounts of leverage, usually aim to liquidate investments within a short time frame, and do not have existing relationships with target firm stakeholders.

A central deal type in PE, which composes the transactions we study, is the leveraged buyout (LBO). In an LBO, the target firm is acquired primarily with debt financing, which is placed on the target firm's balance sheet, and a small portion of equity.<sup>8</sup> One way that PE creates value, sometimes placed under the header of "financial engineering", is to exploit the favorable tax treatment of

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<sup>8</sup> See [Metrick and Yasuda \(2010\)](#) and [Kaplan and Strömberg \(2009\)](#), who provide a detailed discussion of the PE business model and review the academic evidence on their effects.

debt (Spaenjers and Steiner 2020). The reliance on debt means that PE-owned companies have much higher leverage ratios (i.e., debt relative to equity or firm value) than other types of companies, which structurally creates incentives to take risks and requires the company to dedicate a large portion of its cash flows to interest payments (Metrick and Yasuda 2010).

PE is also associated with particularly high-powered incentives to maximize profits because the General Partners (GPs) who manage PE funds are compensated through a call option-like share of the profits (Kaplan and Strömberg 2009). Specifically, their compensation stems primarily from the right to 20% of profits from increasing portfolio company value between the time of the buyout and an exit, when the company is sold to another firm or taken public. Since most funds have 10-year time horizons to return cash to investors, assets are typically held for 3 to 7 years. A modern PE deal is typically not successful if the business continues as-is, motivating aggressive and short-term value-creation strategies. In contrast, a traditional business owner running the firm as a long-term going concern with less leverage may prefer lower but more stable profits.

A large literature in finance shows that PE buyouts increase productivity, increase operational efficiency, and generate high returns.<sup>9</sup> Boucly, Sraer, and Thesmar (2011) show how PE ownership can alleviate credit constraints, enabling more investment. Governance engineering, in the parlance of Kaplan and Strömberg (2009), includes changes to the compensation, benefits, and composition of the management team at the target firm to align their incentives with those of the PE owners, for example, instituting equity-based compensation (Gompers, Kaplan, and Mukharlyamov 2016). Bloom et al. (2015) show that PE-owned firms are better managed than similar firms that are not PE-owned. In operations engineering, GPs apply their business expertise to add value to their investments. For example, they might invest in new technology, expand to new markets, and cut costs (Gadiesh and MacArthur 2008; Acharya et al. 2013; Bernstein and Sheen 2016). Davis et al. (2014) show that after PE buyouts, manufacturing firms expand efficient operations while closing inefficient ones. Work has also found positive effects on product quality (Bernstein and Sheen 2016), workplace safety (Cohn, Nestoriak, and Wardlaw 2021), and product breadth (Fracassi, Previtro, and Sheen 2022), among other metrics.

Considering these changes in the context of nursing homes, the effects of PE ownership on patients are theoretically ambiguous. On the one hand, better management, stronger incentives, and access to credit may lead to improvements in care quality. On the other hand, the literature finding positive effects has primarily studied settings with low information frictions and little government subsidy. In contrast, nursing homes feature severe

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<sup>9</sup> See Kaplan (1989), Kaplan and Schoar (2005), Guo et al. (2011), Acharya et al. (2013), Harris et al. (2014), Robinson and Sensoy (2016), Korteweg and Sorensen (2017), Eaton et al. (2020).

information frictions and misaligned incentives. The intensive government subsidy separates revenue from the consumer. There is also low price elasticity of demand; cost is not salient because Medicare or Medicaid shoulder much of the payment burden. Care quality is opaque, leading to benefits from reallocating care resources to marketing.

Previous owners may have had to commit to implicit contracts with stakeholders, for example, promising that in exchange for government revenue, they would provide quality care at a reasonable cost. They may have been unable or unwilling to take advantage of new opportunities for value creation that would violate these implicit contracts. As a new owner with higher-powered incentives to maximize profits, superior management capability, and a shorter time frame for ownership, the PE investor may be well-positioned to take advantage of these opportunities for value creation.

The higher debt load and incentive misalignment discussed above could act via three dimensions to adversely affect quality. First, cost-cutting measures and a focus on capturing subsidies could come at the expense of quality improvement. Second, large interest payments stemming from the new debt obligations may reduce cash available for care. Relatedly, since PE owners often sell real estate assets shortly after the buyout to generate cash that can be returned to investors, the nursing home may also take on the additional cost of rent. Such cash flows early in the deal's lifecycle boost ultimate discounted returns. For example, in one of the largest PE deals in our sample, the Carlyle Group bought HCR Manorcare for about \$6.3 billion in 2007, of which roughly one quarter was equity and three-quarters were debt. Four years later, Carlyle sold the real estate assets for \$6.1 billion, offering investors a substantial return on equity (Keating and Whoriskey 2018). Afterward, HCR Manorcare rented its facilities; the monthly lease payments are essentially another debt obligation, and a *Washington Post* investigation found that quality of care deteriorated following the real estate sale (Keating and Whoriskey 2018). The third force is the relatively short-term time horizon, which could push managers to maximize short-term profits at the expense of long-term performance. In the case of HCR Manorcare, the nursing home chain was ultimately unable to make its interest and lease payments and entered bankruptcy proceedings in the spring of 2018. Carlyle sold its stake to the landlord.

## 2. Data and Descriptive Statistics

In this section, we briefly summarize our data sources and provide descriptive statistics on the sample, including an analysis of PE targeting. In [Internet Appendix A](#), we describe these elements in comprehensive detail. Since nonprofit nursing homes may have other objectives in addition to profit maximization, comparing their behavior to that of for-profit (and PE-owned) facilities may be misleading. We therefore limit our analysis sample only to

for-profit facilities. Our results should accordingly be interpreted as the differential effects of PE ownership relative to other for-profit owners.

## **2.1 Data**

We obtain facility-level annual data between 2000 and 2017 from publicly available CMS sources. In each year we observe about 12,400 unique skilled nursing facilities (we use the term “nursing home” interchangeably), for a total of approximately 227,000 observations. These data include patient volume, nurse availability, and various components of the five-star ratings, which first appear in 2009. Approximately half of the PE deals in our sample occurred after 2009.

Our second data source consists of patient-level data for Medicare beneficiaries from 2004 to 2016. We use the Medicare enrollment and claims files (hospital inpatient, outpatient, and nursing homes) for the universe of fee-for-service Medicare beneficiaries. We merge these files with detailed patient assessments recorded in the Minimum Data Set (MDS). These data are confidential and were accessed under a data use agreement with CMS. They include patient enrollment details, demographics, mortality, and information about nursing home and hospital care during this period.

In the patient-level analysis, the unit of observation is a nursing home stay that begins during our sample period, which starts in 2005 in order to have at least one look-back year. We consider only the patient’s first nursing home stay in our entire sample period so that we can unambiguously attribute outcomes to one facility and make our patient sample more homogeneous. This produces a sample of about 4.2 million patients over 12 years. We are most interested in the effect on mortality, which is an unambiguously bad outcome, has little measurement error, and is difficult to “game” on the part of a facility or government agency. For these and other reasons, short-term mortality (with suitable risk adjustment) has become the gold-standard measure of provider quality in the health economics and policy literature (McClellan and Staiger 1999; Hull 2018). We use an indicator for death within 90 days following discharge, including deaths that occur in the nursing home. A high level of short-term mortality indicates the general morbidity of this patient cohort: 1 in 6 patients die in the facility or within 3 months of discharge.

We use two measures of spending: the amount billed to Medicare just for the nursing home stay, and the total amount including the stay and all care received in the following 90 days, in case better quality care leads to lower subsequent spending (expressed in 2016 dollars). Medicare covers all costs until the 21<sup>st</sup> day of stay, when the patient begins paying a coinsurance. About 90% of total payments in our data are from Medicare, with patients bearing the remainder. We complement the mortality analysis with three ancillary measures of patient well-being, which CMS uses when computing the Five-Star quality ratings for nursing homes. The first is an indicator for the patient’s self-reported mobility score declining during the stay. The second is an indicator for developing a

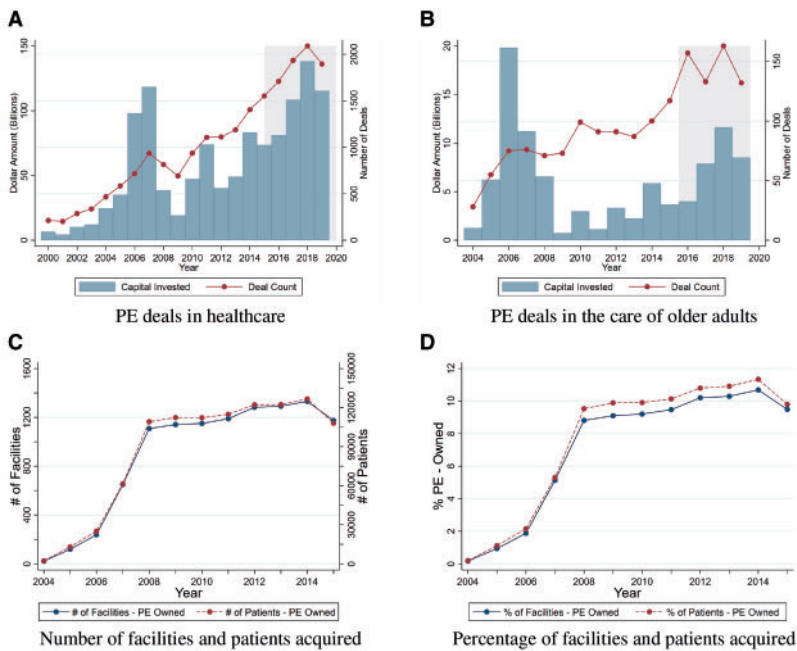
pressure ulcer. The third is an indicator for the patient's self-reported pain intensity score increasing during the stay. To identify nursing homes acquired in PE deals, we make use of a proprietary list of transactions in the "elder and disabled care" sector compiled by Pitchbook Inc., a leading market intelligence firm in this space. The deals span 2004 to 2015, so that we will have sufficient time to evaluate outcomes. We match the target names to individual nursing facilities using name (facility or corporate owner) and address as recorded in the CMS data. This process yields 128 deals, which correspond to a change in ownership to PE for 1,674 facilities. The vast majority of deals in Pitchbook are not at hazard of matching, as they concern assisted living or other care facilities for older adults that are not Medicare-accepting skilled nursing facilities. (See [Internet Appendix A](#) for details.)

[Figure C.1](#) shows the number of deals by year. We observe about 90 unique PE firms that acquired nursing homes. Most deals are syndicated and involve multiple PE firms. [Table C.1](#) presents the top-five deals by number of facilities acquired. Deal sizes are skewed, with the top-five deals accounting more than half of the facilities acquired. On average, we observe PE-owned facilities for 8 years post-acquisition, so the results should be interpreted as medium to long-term effects of PE ownership. While we likely underestimate PE's presence in this sector, our sample size is similar to an estimate of 1,876 nursing homes reportedly acquired by PE firms over a similar duration, 1998–2008 ([GAO 2010](#)). The PE investors in our sample include very large funds, smaller funds, and specialized healthcare PE investment funds. The Carlyle Group and Formation Capital account for the most deals.

## 2.2 Descriptive statistics

Overall, PE investment in healthcare has increased dramatically in recent decades, as shown using Pitchbook data in panel A of [Figure 1](#). Panel B focuses on the Older Adult and Disabled Care subsector, which includes the nursing homes that we study as well as assisted living and other types of care. The shaded areas in the graphs correspond to years after our sample ends, and indicate that deal activity continued to accelerate beyond 2015. The bottom-two panels describe the skilled nursing facilities in our CMS data that are PE-owned. As of 2015, PE-owned facilities represented about 10% of all for-profit nursing facilities, corresponding to an annual flow of about 100,000 patients. Note that the large spike in the mid-2000s seen in all the graphs reflects an economy-wide PE boom during this period, and is not specific to healthcare. Similarly, the flat lining in panels C and D starting in 2010 reflects the lull in deal activity due to the Great Recession. Given the patterns in panel B, the share of facilities that are PE-owned is likely substantially higher today.

[Table 1](#), panel A, presents summary statistics on key variables observed at the facility-year level, where a facility is a single nursing home. Panel B presents summary statistics at the unique patient level on the final Medicare patient sample (recall we focus on a patient's first stay). PE targets are slightly



**Figure 1**  
Private equity ownership in healthcare

This figure shows PE deals in healthcare over time. Panels A and B present the total capital invested (left axis) and number of transactions (right axis) by PE firms in healthcare and care of older adults, by year. Panels C and D focus on the number of active nursing homes owned by PE firms in each year. Panel C presents the number of PE-owned facilities (left axis) and patients admitted at these facilities (right axis). Note that the total number of facilities ever bought by PE firms is larger (1,674) than what is plotted here since some of these facilities closed or went back to non-PE ownership over time. Panel D presents these trends as a percentage of total number of facilities and patients admitted, respectively.

larger, have fewer staff hours per resident, and a lower Overall Five-Star rating. At the sector level, ratings and staffing have secularly increased over time. For staffing, this reflects more stringent regulatory standards. As the PE deals occur later in the sample on average, it is remarkable that they have lower average ratings. Panel B shows that demographic measures are similar across the types of facilities, such as patient age and a high-risk indicator.<sup>10</sup> PE-owned facilities bill about 3% more per stay than non-PE facilities. [Internet Appendix Figure C.2](#), panel A, presents the CDF of stay lengths for Medicare patients in our sample. Medicare stays are relatively short, with a median length of 27 days. We limit the sample to stays fewer than 100 days because Medicare does not pay for longer stays. [Table A.1](#) describes characteristics, measured in the year prior to the deal, associated with buyouts. Facilities in

<sup>10</sup> We use the Charlson Comorbidity index, a standard measure of patient mortality risk based on comorbidities. We create a high-risk indicator equal to one if the Charlson score for stays in the past three months is greater than two.

**Table 1**  
**Descriptive statistics**

	All			Not PE owned		PE owned	
	Mean	SD	Count	Mean	Count	Mean	Count
<b>A. Facility-level attributes</b>							
Overall Five-Star rating	3.05	1.29	111,807	3.07	101,044	2.83	10,763
Deficiency Five-Star rating	2.76	1.24	111,807	2.77	101,044	2.62	10,763
Staff hours per pat. day	3.44	1.32	227,446	3.44	214,456	3.38	12,990
Nurse assistant hours per pat. day	2.21	0.73	227,446	2.22	214,456	2.06	12,990
Licensed nurse hours per pat. day	0.81	0.42	227,446	0.81	214,456	0.82	12,990
Registered nurse hours per pat. day	0.40	0.46	227,446	0.39	214,456	0.49	12,990
Number of beds	108.00	53.41	227,446	107.74	214,456	112.34	12,990
Admissions	186.54	162.71	227,446	181.94	214,456	262.47	12,990
Ratio Black	0.12	0.18	227,446	0.12	214,456	0.12	12,990
Ratio Medicaid	0.63	0.22	227,445	0.64	214,455	0.60	12,990
Ratio Medicare	0.14	0.15	227,445	0.14	214,455	0.18	12,990
Ratio private	0.22	0.17	227,445	0.22	214,455	0.22	12,990
Management fees (2016\$)	7,527	116,603	197,062	6,280	184,498	25,833	12,564
Building lease (2016\$)	6,404	80,743	197,085	5,211	184,521	23,919	12,564
Interest expense (2016\$)	14,663	165,595	197,103	6,079	184,539	140,733	12,564
<b>B. Medicare patient attributes</b>							
Age	81.13	8.09	4,231,460	81.18	3,641,924	80.81	589,536
Female	0.63	0.48	4,231,460	0.64	3,641,924	0.62	589,536
Black	0.09	0.29	4,231,460	0.09	3,641,924	0.09	589,536
White	0.87	0.33	4,231,460	0.87	3,641,924	0.88	589,536
Married	0.35	0.48	4,231,460	0.35	3,641,924	0.36	589,536
Charlson score (previous) > 2	0.27	0.45	4,231,460	0.27	3,641,924	0.29	589,536
Cardio-vascular disease	0.18	0.39	4,231,460	0.18	3,641,924	0.18	589,536
Injury	0.19	0.39	4,231,460	0.19	3,641,924	0.19	589,536
Other	0.63	0.48	4,231,460	0.63	3,641,924	0.63	589,536
Dual eligible	0.17	0.37	4,231,460	0.17	3,641,924	0.15	589,536
Differential distance (miles)	9.77	12.47	4,231,460	11.07	3,641,924	1.75	589,536
Mortality (stay + 90 days)	0.19	0.39	4,231,460	0.18	3,641,924	0.19	589,536
Uses antipsychotics	0.10	0.29	4,231,460	0.10	3,641,924	0.09	589,536
Mobility reduces	0.53	0.50	4,231,460	0.52	3,641,924	0.62	589,536
Develops ulcers	0.09	0.28	4,231,460	0.08	3,641,924	0.09	589,536
Pain intensity increases	0.26	0.44	4,231,460	0.26	3,641,924	0.30	589,536
Amount billed per patient stay (2016\$)	13,100	11,400	4,231,460	13,000	3,641,924	13,400	589,536
Amount billed per patient stay + 90 days (2016\$)	20,800	20,100	4,231,460	20,700	3,641,924	21,100	589,536

This table presents descriptive statistics for key variables used in the analysis. Panel A presents descriptives on facility-level data for all for-profit nursing homes over the years 2000–2017, while panel B presents patient-level data for Medicare patients with index stays over the years 2005–2016. A unit of observation is a facility-year in panel A and a unique patient in panel B (since we retain only the first stay per patient). Columns 1, 2, and 3 present means, standard deviations, and number of observations for the full sample. We categorize facilities into two groups. Columns 4 and 5 present means and number of observations at facilities that never experienced a PE acquisition or before PE acquisition during our sample period. Columns 6 and 7 present corresponding values for facilities in the post-buyout period. For most variables, about 10% of the observations pertain to facilities that experienced a PE acquisition. Sample sizes differ across variables in panel A since they were sourced from multiple sources or in some cases were reported only for more recent years. In panel A, all continuously varying variables were winsorized at the 1% and 99% levels. We compute the Charlson Co-morbidity index using comorbidities recorded in hospital inpatient and outpatient claims (first 10 dx codes) over the 3 months prior to, but not including, the index stay. Spending values in panel B are winsorized at the 99% level and deflated to be in 2016 dollars. ‘Amount billed per patient stay + 90 days’ includes hospital inpatient, outpatient including emergency department, and nursing home stay spending over the 90 days following discharge from the index stay and includes the index stay. The following patient-level variables were sourced from the Minimum Data Set (MDS): marriage, mobility, pressure ulcers, and pain intensity. Medicare patients that could not be merged into the MDS (94% match rate) were dropped from the sample. Facilities with less than 100 Medicare patients over the entire period were omitted from the patient-level sample. If any of the MDS variables was missing, then we set the respective indicator to zero. We exclude patients facing an absolute magnitude of differential distance of greater than 20 miles.

more urban counties and in states with higher older adult population shares are more likely to be targeted.<sup>11</sup> County-level percent Black does not predict buyouts, nor do income and homeownership (not presented). Larger, chain- and hospital-owned facilities are more likely to be acquired than independent facilities, likely reflecting the fixed costs of executing a PE deal. Finally, the Five-Star overall rating has a negative relationship with buyouts, indicating that PE firms target relatively low-performing nursing homes. These factors remain statistically significant predictors when included simultaneously in the same model, shown in column 5. These results highlight the need to estimate the effects of PE ownership within-facility.

### 3. Empirical Strategy for Patient-Level Analysis

There are two primary concerns related to measuring the causal effects of PE ownership on patient-level outcomes. First, PE funds may target facilities that are different in ways the econometrician cannot observe. To partly address this concern, we include facility fixed effects, eliminating time-invariant differences across facilities. We also include market-by-year fixed effects, identifying PE effects off of variation among patients in the same market and in the same year. This common design does not fully account for unobservables driving PE targeting. Therefore, we focus our causality argument on treatment effects for the treated, rather than external validity to a random firm in the economy. Our evidence suggests that treated firms were not on track to the outcomes we observe and would have continued, at least in the medium term, on their preexisting path in the absence of the LBO. This interpretation is important for social welfare as private equity has a significant and growing footprint in the economy. The second concern is that following a PE buyout, the composition of patients may change, further confounding the analysis. Differential customer selection following PE ownership could reflect both supply-side channels, such as changes in advertising and hospital referrals, or patient perceptions about PE ownership. Recent studies have documented that nursing homes selectively admit less costly patients (Gandhi 2022). Hackmann, Pohl, and Ziebarth (2021) find that patients with cognitive impairments and who need help with more activities of daily living (ADL) are the most expensive to serve. Further, CMS payment adjustment emphasized rehabilitation therapy, which favors healthier patients who can tolerate therapy (Carter, Garrett, and Wissoker 2012; Castelluci 2019).<sup>12</sup> We first demonstrate

<sup>11</sup> The map in Figure C.3 shows that deals are not excessively concentrated in particular areas of the country.

<sup>12</sup> Medicare's payment adjustments were heavily tied to additional minutes of physical therapy until 2019. Unlike prospective payments for hospitals, there was no provision for outlier payments for the most expensive patients. For more details on payment adjustment, see <https://www.cms.gov/Research-Statistics-Data-and-Systems/Computer-Data-and-Systems/MDS20SWSpecs/Downloads/44-Group-Worksheet.pdf>. Carter, Garrett, and Wissoker (2012) note this approach created incentives for additional therapy and against admitting clinically complex patients.



that compositional changes do occur and then introduce an instrumental variables approach to address this concern.

For patient-level OLS analyses, we use the following difference-in-differences model, which exploits variation in the timing of the PE deals across facilities:

$$Y_{i,j,r,t} = \alpha_j + \alpha_{r,t} + \phi PE_{i,j,r,t} + X'_{i,z} \gamma + \varepsilon_{i,j,r,t}. \quad (1)$$

Here,  $PE_{i,j,r,t}$  is an indicator set to one if patient  $i$  in Hospital Referral Region (HRR),  $r$ , chooses PE-owned facility  $j$  in year  $t$ . Our preferred model includes facility fixed effects ( $\alpha_j$ ) and patient HRR-by-year fixed effects ( $\alpha_{r,t}$ ). We allow markets to evolve on different trends to mitigate the possibility of differences in market structure confounding the results. The vector  $X_{i,z}$  denotes patient risk controls including age, indicators for gender, marital status, dual eligible, and 17 disease categories.<sup>13</sup> Standard errors are clustered by facility to account for unobserved correlation in outcomes across patients treated at the same nursing home. To address potential concerns with staggered treatment, we use the [Callaway and Sant'Anna \(2021\)](#) and [Sun and Abraham \(2021\)](#) estimators in robustness tests.<sup>14</sup> Consistent with the existing literature cited above, we show that patient risk declines following PE ownership. Table 2, panel A, presents point estimates for patient risk measures. Specifically, we test for changes in *initial* patient risk (assessed at the time of admission) following acquisition. We examine effects on a mix of comorbidities to broadly capture changes in patient risk. The coefficients indicate that admitted patients are less likely to suffer from cognitive impairments (depression, dementia, Alzheimer's) and need help with fewer ADLs following PE ownership, factors which strongly predict treatment costs. Figure 2 presents the corresponding event study plots, which generally suggest flat or increasing trends in patient risk prior to the deal but declining trends following the acquisition. We are concerned that this shift toward a healthier patient composition will lead to downward bias in mortality and spending effects. Therefore, we develop an instrument for the match between patients and nursing homes.

### 3.1 Distance-based instrument

We combine the differences-in-differences model above with a differential distance instrument ([McClellan, McNeil, and Newhouse 1994](#)) to control for endogenous patient selection into nursing homes. The thought experiment we approximate is to randomly draw a patient who goes to a PE facility after the buyout relative to a randomly drawn patient who went to that

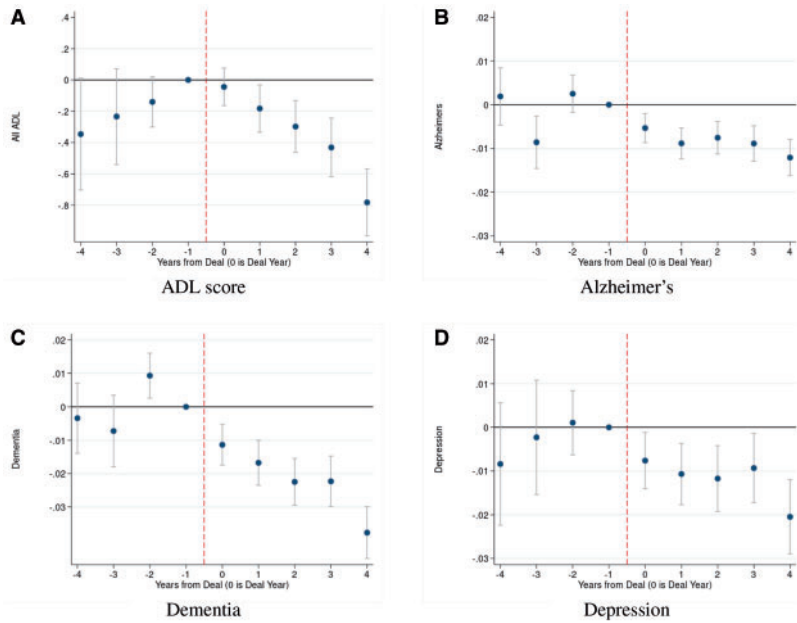
<sup>13</sup> We construct these indicators with diagnostic codes recorded in claims from the 3 months prior to the index nursing home stay (hospital stays, ED visits, and outpatient visits). "Dual eligible" is commonly used to describe patients eligible for both Medicare and Medicaid.

<sup>14</sup> The former compares the outcomes of treated facilities with never-treated facilities, to ensure that using ever-treated facilities as controls does not bias the results. The latter corrects for treatment effect heterogeneity by weighting observations in accordance with the share of facilities treated in each year.

**Table 2**  
**Effect of private equity buyouts on patient outcomes**

<i>A. Initial patient assessments (OLS estimates)</i>				
	(1) ADL score at admission	(2) Alzheimer's at admission	(3) Dementia at admission	(4) Depression at admission
1(PE)	-0.1124* (0.062)	-0.0047*** (0.001)	-0.0157*** (0.002)	-0.0060** (0.002)
Observations	3,962,599	3,962,599	3,962,599	3,953,806
Y-mean	15.90	0.05	0.17	0.25
<i>B. OLS estimates</i>				
	(1) Mortality (Stay + 90 days)	(2) log(Amount billed) Per patient stay	(3) log(Amount billed) Per patient stay + 90 days	
1(PE)	0.0030** (0.001)	-0.0258*** (0.006)	-0.0163*** (0.005)	
Observations	4,231,356	4,231,356	4,231,356	
Y-mean	0.18	9.05	9.56	
<i>C. IV estimates</i>				
	(1) Mortality (Stay + 90 days)	(2) log(Amount billed) Per patient stay	(3) log(Amount billed) Per patient stay + 90 days	
1(PE)	0.0195*** (0.007)	0.0768*** (0.025)	0.0604*** (0.021)	
Observations	4,231,356	4,231,356	4,231,356	
Y-mean	0.18	9.05	9.56	
F-stat	243	243	243	
<i>D. Placebo analysis (IV estimates)</i>				
	(1) Mortality (Stay + 90 days)	(2) log(Amount billed) Per patient stay	(3) log(Amount billed) Per patient stay + 90 days	
1(PE)	0.002 (0.004)	-0.0215 (0.016)	-0.0241* (0.013)	
Observations	3,710,981	3,710,981	3,710,981	
Y-mean	0.20	9.02	9.53	
F-stat	590	590	590	

This table presents estimates of the relationship between PE ownership and patient health and spending. Panel A presents OLS results for initial assessments of patients entering nursing homes, obtained by estimating Equation (1). In panel B, each cell presents the coefficient  $\phi$  obtained by estimating Equation (1). The independent variable is an indicator for the patient being admitted to a PE nursing home. We present effects for claims-based patient quality outcomes, patient death within 90 days of discharge from the index stay, and total amount billed (2016\$). Panel C presents the coefficient  $\phi$  for the same outcomes obtained by estimating Equation (3) by 2SLS, with the indicator for the patient being admitted to a PE nursing home, instrumented by differences in distance to the nearest PE and non-PE facility. Panel D presents these results from a placebo analysis of the relationship between private equity ownership and patient health and spending. We assign placebo PE acquisition to 3 years before the actual acquisition and discard data for any facility starting with the year it actually got acquired. Accordingly, we recompute differential distance values taking into account these placebo acquisitions. All regressions include facility and patient HRR  $\times$  Year fixed effects, and patient risk controls. Patient risk controls include age, race, gender, marital status, indicators for 17 preexisting conditions used to compute the Charlson index, and an indicator if patients are dual eligible. Standard errors are clustered by facility. \*  $p < .1$ , \*\*  $p < .05$ , \*\*\*  $p < .01$ .



**Figure 2**  
**Initial patient assessments**

This figure presents event studies on initial patient assessments around the time a nursing home experiences a PE buyout. We estimate these models on collapsed facility-year-level data. Each point in the figures represents a coefficient obtained by estimating an event study version of Equation (1). The model is fully saturated, including coefficients for all years, except Year = -1, which is the omitted point, but we plot only years -4 to 4. Panel A presents results on activities of daily living (ADL) score for patients where a higher score indicates more independence, panel B on an indicator for Alzheimer's, panel C on an indicator for Dementia, and panel D on an indicator for Depression, respectively, at admission to the index nursing home stay. Standard errors are clustered by facility.

facility before the buyout, and then compare this difference to an analogous one in the same set of years for patients at non-PE facilities. The instrument to simulate randomly drawing patients exploits patient preference for healthcare providers located nearby (Einav, Finkelstein, and Williams 2016; Card, Fenizia, and Silver 2023; Currie and Slusky 2020). This is especially true for nursing homes; for example, Hackmann (2019) finds that the median distance between a senior's residence and her nursing home is under 4.3 miles. In our data, the median and 90th percentile distances between a patient and her nursing home are 4.8 and 18 miles, respectively. About 33% of all patients choose the facility located closest to them (see Figure C.4).<sup>15</sup> As a result of these patterns, this instrument has been used to control for patient selection into nursing homes (Grabowski et al. 2013; Huang and Bowblis 2019).

<sup>15</sup> Distance patterns remain remarkably stable over time in our sample. Figure C.4D shows that mean distance to facility is unaffected by PE buyout.

The differential distance instrument is the difference (in miles) between two distances: from a patient’s home ZIP code to the closest PE-owned facility ZIP code; and from the patient’s residence to the nearest non-PE facility ZIP code. Lower values of the instrument mean the patient is relatively closer to a PE-owned facility. When it is negative, the nearest PE-owned facility is closer than the nearest non-PE-owned facility. PE ownership evolves over time as more deals take place (and some PE funds exit their investments), creating variation across years in differential distance for individuals residing in the same ZIP code. Following convention, we drop patients facing a large differential distance value, specifically, one of more than 20 miles.<sup>16</sup> This also has two benefits: First, included patients plausibly live in markets targeted by PE firms and are thus more homogeneous; Second, the instrument has more power because it excludes inframarginal patients in places where facility choice is not sensitive to differential distance.

We estimate the first and second stages using Equations (2) and (3), respectively. The endogenous regressor of interest  $PE_{i,j,r,t}$  is, as above, an indicator set to one if patient  $i$  in Hospital Referral Region (HRR),  $r$ , chooses PE-owned facility  $j$  in year  $t$ . We instrument with linear and squared differential distance,  $D_i$ , applicable to patient  $i$  based on her ZIP code,  $z$ , in the year the stay begins,  $t$ .

$$PE_{i,j,r,t} = \alpha_j + \alpha_{r,t} + \zeta_1 D_i + \zeta_2 D_i^2 + X'_{i,z} \xi + v_{i,j,r,t}, \quad (2)$$

$$Y_{i,j,r,t} = \alpha_j + \alpha_{r,t} + \phi \hat{P}E_{i,j,r,t} + X'_{i,z} \gamma + \varepsilon_{i,j,r,t}. \quad (3)$$

The other variables are as described for Equation (1). It is crucial to include facility fixed effects here ( $\alpha_j$ ), to control for level characteristics that attract PE ownership but are not caused by it. Our setting departs from that in [McClellan, McNeil, and Newhouse \(1994\)](#), whose canonical paper used differential distance. The authors studied the causal effect of using a particular clinical procedure rather than a facility-level attribute of ownership.

The instrument is strongly predictive of nursing home type. The first stage results are reported in Table 3. Column 2 presents the estimates from our preferred specification. A five-mile decrease in differential distance (0.4 s.d.) increases the probability of going to a PE-owned nursing home by 3.3 percentage points (pp), about 24% of the mean. The F-statistic exceeds 220, well above conventional rule-of-thumb thresholds for weak instruments.<sup>17</sup> We conduct multiple robustness checks, including time-varying socioeconomic

<sup>16</sup> We exclude ZIP codes based on the absolute magnitude of the differential distance, treating patients very close to PE facilities the same as those very far away. In practice, the differential distance is rarely less than  $-20$ . Differential distance values update for some ZIP codes over time as facilities are acquired or sold by PE firms. We exclude such ZIP codes only if their differential distance remains more than 20 miles in magnitude throughout.

<sup>17</sup> An alternative approach to constructing differential distance is to consider the distance from the hospital where the patient was treated prior to the nursing home stay, rather than the home residence. However, this has found to be a weaker instrument ([Rahman, Norton, and Grabowski 2016](#); [Cornell et al. 2019](#)).

**Table 3**  
**Patient-level analysis: First stage**

	(1) 1(PE)	(2) 1(PE)	(3) 1(PE)	(4) 1(PE)
Differential distance (In 10 miles)	-0.0699*** (0.003)	-0.0699*** (0.003)	-0.0698*** (0.003)	-0.0451*** (0.002)
(Differential distance) <sup>2</sup> (In 10 miles)	0.0092*** (0.001)	0.0092*** (0.001)	0.0091*** (0.001)	0.0066*** (0.001)
Facility FEs	Y	Y	Y	Y
Patient controls		Y	Y	Y
ZIP code controls			Y	
Patient FEs level	HRR × Year	HRR × Year	HRR × Year	HSA × Year
Observations	4,231,356	4,231,356	4,227,196	4,230,487
Y-mean	0.14	0.14	0.14	0.14
F-stat	243	243	241	220

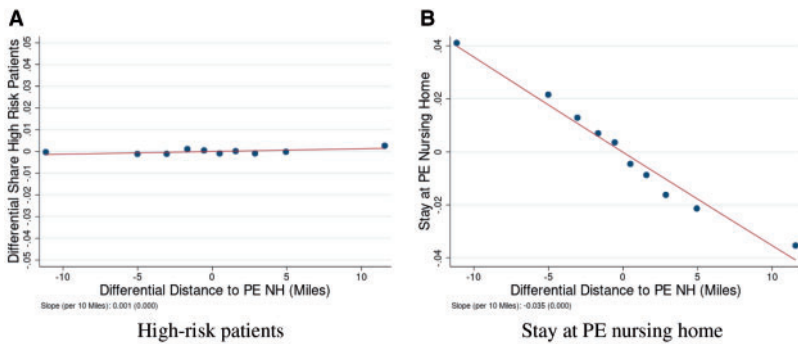
This table presents estimates of the relationship between PE ownership of the nursing home and the patient's differential distance. The cells presents the coefficients  $\zeta_1$  and  $\zeta_2$  obtained by estimating Equation (2). The independent variable is the difference in distance (both linear and quadratic, in 10 miles) to the nearest PE nursing home and the nearest non-PE nursing home for the patient. This is calculated based on distances between the respective ZIP code centroids. The outcome variable is an indicator for whether the nursing home serving the patient is PE-owned (=1 if PE-owned, 0 otherwise). Column 1 controls for facility and patient market (Hospital Referral Region) × Year fixed effects. Column 2 (our preferred specification) adds controls for patient risk controls (indicators for 17 preexisting conditions used to define the Charlson Co-morbidity index inferred from claims over the 3 months prior to admission, and sex, age, race, marital status, and an indicator if patients are dual eligible). Column 3 adds controls for patient ZIP-year characteristics: median household income, the shares of the population that are white, that are renters rather than homeowners, that are below the Federal poverty line, and that are enrolled in the Medicare advantage program. Column 4 uses the same controls as in column 2 but defines patient market using a narrower market definition: Hospital Service Area (HSA) instead of HRR. Standard errors are clustered by facility. \*  $p < .1$ , \*\*  $p < .05$ , \*\*\*  $p < .01$ .

variables at the patient's ZIP-code-year level ( $z$ ) and omitting all controls other than fixed effects.<sup>18</sup>

IV estimation differs from randomized controlled trials because it requires two untestable assumptions. The first is conditional random assignment, under which unobserved characteristics correlated with the outcomes of interest are not correlated with differential distance after conditioning on covariates. This subsumes the exclusion restriction, which is that the patient's differential distance to a PE facility affects outcomes only by influencing her probability of being treated at a PE facility. To provide support for the conditional randomization assumption, we examine the correlation between the instrument and patient observables, particularly covariates which may affect mortality, such as risk. Figure 3, panel A, presents the relationship between patient risk and the instrument and indicates little or no correlation.<sup>19</sup> The figure shows that the probability of being a high-risk patient (Charlson score > 2)

<sup>18</sup> The socioeconomic variables, from the American Community Survey, are annual median household income, the share of the population that is white, that rents versus owns a home, and that is below the Federal poverty line. In unreported analyses, we find similar results if we use a linear model for differential distance rather than a quadratic one.

<sup>19</sup> We project the high-risk indicator (see Section A.2) on the controls we use in our main regression and collapse the residuals into 10 bins. Similarly, we run a regression of differential distance on the controls and collapse the



**Figure 3**  
**Patient characteristics with differential distance**

This figure presents scatter plots of patient characteristics against differential distance to the nearest PE facility. The independent variable is the difference (in miles) of the nearest PE nursing home to the nearest non-PE nursing home for the patient. The dependent variable in panel A is an indicator for the patient to have a Charlson Co-morbidity index (based on diagnoses recorded in hospital inpatient and outpatient claims over the 3 months before admission to nursing home) greater than 2, and in panel B is an indicator for the nursing home being PE-owned. The data were collapsed into 10 equally sized bins, and we plot the means of residuals in each bin that were obtained from models including facility and patient HRR  $\times$  Year fixed effects, and patient demographics: age, race, gender, marital status, and an indicator if patient is dual eligible. Panel B additionally controls for indicators for 17 preexisting conditions used to compute the Charlson index. The figures also present quadratic fitted lines for these plots. Each plot also presents the slope coefficient (per 10 miles of differential distance) with the corresponding standard error. Standard errors are clustered by facility.

increases by 0.1% for a 10-mile increase in differential distance. This is small in absolute terms and negligible compared to the proportion of high-risk patients in the sample (27%). Table 3 columns 1 and 2 show that the coefficients for differential distance are unaffected by including patient-level controls, corroborating this interpretation. One important test adds time-varying ZIP-code-level socioeconomic controls in case PE firms target places on track to different demographic profiles (Table 3, column 3). We also confirm that we find similar patterns when we use a much more granular market definition, to mitigate the concern of within-market targeting of patients (Table 3, column 4).<sup>20</sup>

Under random assignment, characteristics of patients with above- and below-median differential distance should be similar. Table 4, where we summarize patient characteristics for the two groups, suggests this is the case. The top-two rows of the table show that, consistent with a strong instrument, the probability of going to a PE-owned facility declines from 24% for the below-median group to 4% for the above-median group. The patient characteristics

residuals into 10 bins. We plot the means of each bin, with the risk residuals on the y-axis and distance residuals on the x-axis. The figure also presents a fitted line and the slope coefficient.

<sup>20</sup> We use hospital service areas (HSAs) as an alternate, more granular definition of nursing home markets. There are nearly 3,400 HSAs in the United States, while there are only about 300 HRRs. The Dartmouth Atlas group defines both HRRs and HSAs based on healthcare use patterns by Medicare beneficiaries so that they are relatively self-contained.

**Table 4**  
**Balance of patient characteristics**

Patient attribute	(1) DD < Median	(2) DD > Median
Differential distance	1.63	17.91
PE-owned nursing home	0.24	0.04
Age	81.10	81.16
Female	0.63	0.64
Black	0.10	0.08
Married	0.35	0.34
Dual eligible	0.15	0.18
<u>Charlson Score categories:</u>		
AMI	0.08	0.08
Congestive heart failure	0.23	0.24
PVD	0.06	0.05
CEVD	0.13	0.14
Dementia	0.04	0.04
COPD	0.22	0.23
Rheumatoid arthritis	0.03	0.03
Peptic ulcer	0.02	0.02
Mild liver disease	0.01	0.01
Diabetes	0.21	0.22
Diabetes + Complication	0.04	0.04
Paraplegia	0.03	0.03
Renal disease	0.14	0.13
Cancer	0.09	0.09
Severe liver disease	0.01	0.01
Metastatic cancer	0.04	0.04
AIDS	0.00	0.00
<u>Initial assessments:</u>		
ADL score at admission	15.89	15.87
Alzheimer's at admission	0.05	0.06
Dementia at admission	0.17	0.17
Depression at admission	0.25	0.26
Number of patients	2,116,500	2,114,960

This table presents the balance in patient attributes with respect to the instrument: differential distance. We divide patients into two groups based on whether their differential distance is below or above the median value (8.9 miles). Recall that differential distance (DD) is the difference between distance to the nearest PE nursing home and the nearest non-PE nursing home for the patient. Column 1 presents the means of patient characteristics for patients with DD below the median value, while column 2 presents the means for patients with DD greater than the median. We present patient demographics, 17 comorbidity indicators used to compute the Charlson Comorbidity index, and 4 initial assessment characteristics of patients at the time of admission to the index nursing home stay from the Minimum Data Set. Charlson score categories use diagnostic codes on hospital inpatient and outpatient claims over the 3 months prior to, but not including, the index nursing home stay. Paraplegia includes both partial and complete paralysis. We generated indicators for the Charlson score disease categories using the 'charlson' command in Stata, available at <http://fmwww.bc.edu/RePEc/bocode/c/charlson.html>.

in the subsequent rows are extremely similar between the two groups. For example, the two groups have nearly identical mean ages and shares of patients that are female or married. The instrument also appears to balance patients on the same four measures of cognitive impairments and activities of daily living that decline post-buyout. [Internet Appendix B](#) describes evidence on random assignment in more detail. For example, one test in the spirit of [Angrist, Lavy, and Schlosser \(2010\)](#) and [Grennan et al. \(2021\)](#) shows that

differential distance lacks predictive power for inframarginal patients, in the first stage and in the outcome equations.

The second assumption is monotonicity, under which lower differential distance makes all patients more likely to choose a PE-owned facility. This is true on average, but the assumption is at the patient-level which is untestable. Monotonicity is necessary to interpret the IV estimate as a well-defined local average treatment effect (LATE). Figure 3, panel B, contains a binscatter plot of the first stage, showing that the likelihood of going to a PE-owned facility decreases nearly linearly with differential distance. It is estimated in the same way as panel A described above, except that the outcome is an indicator for the facility being PE-owned. The linear pattern is consistent with monotonicity. [Internet Appendix B](#) describes further evidence supporting this assumption.

#### 4. Patient-Level Effects

This section presents the main results of the paper, which are the effects of PE ownership on short-term mortality and spending per patient. It then considers observed and unobserved heterogeneity in these effects. Finally, it examines measures of patient well-being.

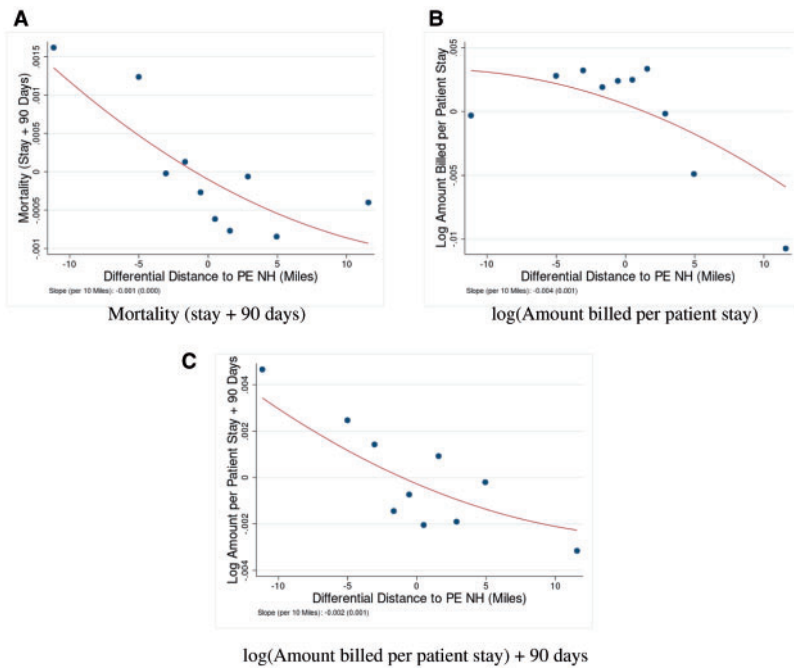
##### 4.1 Effects on mortality and spending

We begin with OLS models using Equation (1), which includes patient-level controls, facility fixed effects, and patient HRR-by-year fixed effects. Table 2, panel B, contains the results. We find an OLS effect on mortality of 0.3 pp, which is about 2% of the mean. There are small, negative effects on spending of 1%–3% (columns 2 and 3). We expect that the selection on unobservedly lower risk should bias these OLS results down, as discussed in Section 3 and as indicated by the risk measures in Figure 2. The corresponding event studies, in Figure C.6, suggest no pre-trends, supporting the parallel trends assumption that underlies our empirical model (i.e., targeted and control facilities would continue on parallel trends absent the buyout).<sup>21</sup>

Table 2, panel C, reports the IV effects from Equation (3). Consistent with downward bias in OLS, we see much larger effects in the IV analysis. However, it is important to emphasize that the IV effects are for compliers with the instrument who go to a PE facility because it is closer, and thus could experience larger effects than a randomly selected patient. Column 1 shows that going to a PE-owned nursing home increases the probability of death during the stay and the following 90 days by about 2 pp, or 11% of the mean. In the context of the health economics literature, this is a very large effect. Next, the amount billed per nursing home stay per patient—which is almost all paid by Medicare—increases by 8% (column 2). As Table 1 shows, on

<sup>21</sup> These figures are constructed by collapsing data to the facility-year level and estimating an event study version of Equation (1).





**Figure 4**  
**Patient outcomes with differential distance**

This figure presents scatter plots of patient characteristics against differential distance to the nearest PE facility. The independent variable is the difference in distance (in miles) of the nearest PE nursing home to the nearest non-PE nursing home for the patient. The dependent variable in panel A is an indicator for patient death during or 90 days post-nursing home stay, in panel B refers to the logarithm of the total payment for the index nursing home stay, and in panel C is the logarithm of the total payment for index stay and 90 days post-nursing home stay. The data were collapsed into 10 equally sized bins, and we plot the means of residuals in each bin that were obtained from models including facility and patient HRR  $\times$  Year fixed effects, and patient demographics: age, race, gender, marital status, indicators for 17 preexisting conditions used to compute the Charlson index, and an indicator if patient is dual eligible. The panels also present quadratic fitted lines for these plots. Each plot also presents the slope coefficient (per 10 miles of differential distance) with the corresponding standard error. Standard errors are clustered by facility.

average PE-owned nursing homes bill \$13,400 per stay, while non-PE nursing homes bill \$13,000. Higher costs do not seem to reflect additional preventive care that enables lower costs later, because the total amount billed for the stay and the subsequent 90 days increases by 6%. The IV estimates imply that the reduced form effects should decline as differential distance grows larger (i.e., relative to the nearest alternative, a PE facility is farther away). Figure 4 provides nonparametric evidence of such a pattern, using the same approach as in Figure 3, except that the y-axis variables are the patient outcomes. Consistent with the IV results, mortality and spending are highest among patients with the lowest differential distance and decline as patients are relatively farther away from PE facilities.

We calculate the implied cost in statistical value of life-years in [Table C.3](#), panel C. The IV coefficients are mapped to lives and life-years lost using the number of index stays at PE-owned nursing homes during our sample period. This calculation implies that about 22,500 additional deaths occurred due to PE ownership over the 12-year sample period. To estimate life-years lost, we rely on observed survival rates for Medicare patients at all nursing homes. This leads to an estimate of about 172,400 lost life-years.<sup>22</sup> Applying a standard estimate of statistical value of a life-year of \$100,000 ([Cutler and McClellan 2001](#)) inflated to 2016 dollars, we arrive at a mortality cost of \$22.4 billion.

We present a range of robustness tests in [Internet Appendix B](#), and summarize the most important ones here. First, we use a placebo analysis to assess whether preexisting trends might explain the results. We artificially set the PE dummy to turn on before the deal and drop observations after the true deal year. [Table 2](#), panel D, finds economically small and insignificant or marginally significant placebo effects, consistent with no differential trends prior to acquisition. [Table 5](#), row 2, reports specification checks that vary the controls and market definition. Importantly, we expect that if the instrument does not randomly assign patient risk, including patient controls should substantially affect the results. In row 2.A, we omit all patient controls. In row 2.B., we include ZIP-year socioeconomic controls.<sup>23</sup> Across all of the tests, the coefficients remain robust and similar in magnitude to the baseline estimates in row 1. [Internet Appendix B](#) presents other checks that validate the instrument, such as showing that it is very weak among patients who are relatively very far or very close to PE-owned facilities or equally close to PE and non-PE facilities, and does not explain mortality for these patient groups ([Table C.4](#)). The effect on mortality is robust to using alternative durations to define the metric. [Table C.5](#) shows that the effect varies between 9% and 12% of the mean mortality rate for horizons ranging from 15 to 365 days. Finally, we present the alternative OLS [Callaway and Sant'Anna \(2021\)](#) and [Sun and Abraham \(2021\)](#) estimators in [Figure C.5](#), panel A.

The results so far point to nuanced effects of PE ownership. Since patients become less risky after PE buyouts, two things could come into play: the facility may be cream-skimming or patients may be changing how they select a nursing

<sup>22</sup> As life expectancy differs substantially between men and women, we estimate the effect separately by gender. We calculate the average life expectancy at discharge by gender by observing the actual life span for each patient in our data. For patients still alive at the end of our sample period, we approximate the year of death based on patient gender and age using Social Security actuarial tables. We adjust this downward to account for the fact that decedents tend to be older on average (by about 2 years). We then apply this mean life expectancy to the number of deaths computed above and obtained the number of life-years lost. This approach may overstate the true value if the *incremental* deaths at PE facilities are of older patients. This approach also understates the true value since we don't account for the loss in longevity not resulting in death during our sample.

<sup>23</sup> The other tests are as follows. In row 2.C, we use more granular HSAs instead of HRRs to define patient markets. In rows 3A and 3B, we test sensitivity to the varying the maximum threshold of differential distance for sample inclusion. The coefficients are robust to using a narrower (15-mile) or wider (25-mile) threshold than the baseline value of 20 miles. In row 4, we use an indicator for above-median differential distance rather than the continuously varying value. In row 5, we cluster standard errors by deal rather than by facility.

**Table 5**  
**Patient-level analysis: Robustness**

	(1) Mortality (Stay + 90 days)	(2) log(Amount billed) Per patient stay	(3) log(Amount billed) Per patient stay + 90 days
<b>1. Base specification</b>			
1(PE)	0.0195*** (0.007)	0.0768*** (0.025)	0.0604*** (0.021)
<b>2. Varying controls</b>			
<b>A. No controls</b>			
1(PE)	0.0285*** (0.008)	0.1162*** (0.026)	0.0634*** (0.022)
<b>B. ZIP-year controls</b>			
1(PE)	0.0187*** (0.007)	0.0776*** (0.025)	0.0599*** (0.021)
<b>C. HSA-year FEs</b>			
1(PE)	0.0402*** (0.015)	0.1113*** (0.043)	0.1286*** (0.038)
<b>3. Changed distance threshold</b>			
<b>A. Narrower distance threshold</b>			
1(PE)	0.0177** (0.007)	0.0508** (0.025)	0.0406* (0.021)
<b>B. Wider distance threshold</b>			
1(PE)	0.0181** (0.007)	0.0986*** (0.025)	0.0729*** (0.021)
<b>4. Alternate functional form</b>			
1(PE)	0.0297*** (0.010)	0.1423*** (0.032)	0.1060*** (0.028)
<b>5. Clustering by deals</b>			
1(PE)	0.0195** (0.009)	0.0769*** (0.028)	0.0604** (0.028)
<b>6. Corporatization</b>			
<b>A. Include chain controls</b>			
1(PE)	0.0195*** (0.007)	0.0769*** (0.025)	0.0604*** (0.021)
<b>B. Top-5 deals only</b>			
1(PE)	0.0408*** (0.012)	0.1017** (0.040)	0.0866** (0.035)
<b>7. Excluding select deals</b>			
<b>A. W/O top-2 deals</b>			
1(PE)	0.0330*** (0.011)	0.1034*** (0.038)	0.0929*** (0.032)
<b>B. Excluding Formation and Fillmore Capital</b>			
1(PE)	0.0426*** (0.013)	0.1088** (0.044)	0.1300*** (0.038)
<b>8. Excluding top-5 states with retirement inflows and outflows</b>			
1(PE)	0.0188** (0.008)	0.0891*** (0.027)	0.0739*** (0.023)
Observations	4,231,356	4,231,356	4,231,356
Y-mean	0.18	9.05	9.56

This table presents results from specification checks on the relationship between PE ownership and patient health and spending, corresponding to results in panel C in Table 2. The first panel presents the base specification. The second panel presents results by varying controls: row 2A presents coefficients from a model with fixed effects only, row 2B includes patient ZIP controls: median household income, the shares of the population that are white, that are renters rather than homeowners, that are below the federal poverty level, and that are enrolled in Medicare Advantage program, and while all other rows include HRR  $\times$  year fixed effects, row 2C uses Hospital Service Areas (HSA) instead of HRR. The third panel tests robustness to varying the distance threshold to include patients. Rows 3A and 3B exclude all patients whose ZIP code has a minimum differential distance of more than 15 miles and 25 miles, respectively. The fourth panel uses an indicator of differential distance greater than median value rather than actual differential distance for the estimates. The fifth panel clusters estimates by deals instead of facilities. The sixth panel checks if results are driven by PE ownership or corporatization: row 6A controls for facility being part of a chain, and row 6B limits the PE group to only the facilities bought in the 5 largest PE deals. The seventh panel presents various sample cuts: row 7A calculates the results excluding all data for chains involved in the 2 largest PE deals, and row 7B excludes all deals with Formation and Fillmore Capital. The eighth panel excludes five states with largest retiree inflows and outflows. All rows (except 2A) include patient risk controls: age, race, gender, marital status, indicators for 17 preexisting conditions used to compute the Charlson score, and an indicator if patients are dual eligible. Standard errors are clustered by facility (except for fifth panel). \*  $p < .1$ , \*\*  $p < .05$ , \*\*\*  $p < .01$ .

home. We find a very large IV effect on mortality and a small OLS effect, consistent with the OLS capturing some compositional shifts toward less risky patients. The IV result represents a causal effect within the subset of patients who comply with the instrument by going to the closest facility; these patients may be more vulnerable, in the sense that they do not opt to travel farther to find the best match, and may face more information frictions. In this subset, we find a very large, positive effect on mortality. We discuss the IV effects further below in a marginal treatment effects analysis.

**Alternative Explanations** We assess the plausibility of several alternative interpretations of our main results. First, PE ownership could bring economies of scale or corporatization, which [Eliason et al. \(2020\)](#) propose to explain the negative effects of dialysis center mergers. We conduct two tests for this in Table 5. The first adds to our main model a control for being a chain versus an independent facility (about 15% of PE-owned facilities remain independent post-buyout). If our effects are explained by the “rolling-up” of independent facilities into more efficient chains, the estimates should attenuate. Instead, they are essentially unchanged (row 6.A). Second, we use only the top-five deals to define PE ownership. In these deals, the target chains already owned more than 100 facilities and stayed nearly the same size over the sample period. Therefore, in this model chain size is held constant and we evaluate the effect of a change in ownership. The effects remain large and significant (row 6.B). In sum, neither chain corporate structures nor synergies in large firms seems to explain our results.

Second, perhaps only select large deals or PE firms drive the results. However, the estimates remain large and significant when we exclude facilities bought in the top-two deals, each involving more than 300 facilities (row 7.A). Similarly, the results are robust to excluding facilities bought by Formation Capital and Fillmore Capital, which may not be representative of the average PE firm given their specialization in nursing homes (row 7.B).

Third, we test whether the results are limited to states with rapid changes in the aged share of the population during our sample period due to heavy inflow or outflow of retirees. PE firms may disproportionately target or avoid facilities in such states if they expect them to experience rapid growth or decline in nursing home demand. Therefore, we exclude the five states with highest net inflows and outflows of retirees during this period, respectively, from our sample and estimate our main results.<sup>24</sup> The results, in row 8, remain very similar to our main estimates.

Finally, if PE firms also acquire hospitals in the same market along with nursing homes, they might also affect hospital quality and the share of patients

<sup>24</sup> The states with highest net inflows are Arizona, Florida, Idaho, Nevada, and South Carolina. The states with highest net outflows are Wyoming, Vermont, New York, New Jersey, and Illinois. The U.S. Census identified these states in its Current Population Report as having the highest and lowest net migration rates for the population over 65 years ([Mateyka and He 2022](#), figure 1).

discharged to nursing homes. In other words, changes in the upstream hospital market could bias our estimates. To assess this, we conduct two tests. The first examines how PE entry into a market affects hospital quality, measured as changes in 90-day mortality rates for all Medicare heart attack patients admitted to hospitals, a standard hospital quality metric used by CMS and the economics literature (Chandra et al. 2016). Table C.6, panel A, presents OLS estimates from patient-level models in which the outcome is an indicator for 90-day mortality of Medicare heart attack patients. We detect small, statistically insignificant, but precisely estimated effects, regardless of how we define the market. Second, panel B presents the corresponding effect on the proportion of hospital patients that are discharged to a nursing home. These coefficients are also small and statistically insignificant. Hence, we do not find evidence to support this concern.

## 4.2 Heterogeneity in the IV mortality effect

The selection on risk we document above raises the question of whether the effects are larger for some groups than for others. This section explores heterogeneity both on observed attributes and on unobserved resistance to treatment, using a marginal treatment effects (MTE) framework.

**4.2.1 Observed attributes** To assess heterogeneity in the IV analysis, we split the sample based on observed characteristics in Table 6. Panel A presents results for different patient groups, while panel B presents a companion analysis for different groups based on market or facility characteristics. We also report the mean mortality rates for each subgroup to help interpret the magnitudes. We begin by describing heterogeneity across patients reported in panel A. First, higher risk as measured by disease burden should be associated with more need for high-skill, medicalized RN care. Lower-risk patients might be more sensitive to changes in staff attentiveness (e.g., helping them to use the toilet or minimizing infection risk). Therefore, we split the sample into two groups around the high-risk indicator (Charlson score above two), which isolates patients with higher mortality (29% vs. 14%). The results indicate that relative to the mean mortality rate, lower-risk patients experience a much greater increase in mortality ( $0.02/0.14 \approx 14\%$  vs.  $0.02/0.29 \approx 8\%$ , row 1).

We consider gender in panel A, row 2, and find similar effects among men and women. Third, we divide the sample by the median length of stay. Note that length of stay could be affected by PE ownership, so this analysis should be thought of as relevant to understanding the mechanism. The mortality effect is driven by patients with below median stays, consistent with the previous result that lower-risk patients experience worse mortality effects. It also contradicts a potential concern that PE facilities appear worse on mortality because of sicker patients who require long stays and are independently more likely to die.

We explore this further by studying the impact of PE ownership on length of stay at different parts of the length of stay distribution. We estimate a series of

**Table 6**  
**Heterogeneity in patient mortality**

	(1) Observations	(2) Mean	(3) Coefficient	(4) (Std. Errors)
<i>A. Patient-Level characteristics</i>				
<b>1. Risk</b>				
Low risk	3,073,198	0.14	0.0197**	(0.008)
High risk	1,157,836	0.29	0.0216	(0.014)
<b>2. Gender</b>				
Male	1,547,356	0.23	0.0202*	(0.012)
Female	2,683,713	0.16	0.0186**	(0.008)
<b>3. Length of Stay</b>				
Length of stay < median	2,159,020	0.22	0.0342***	(0.011)
Length of stay > median	2,072,060	0.15	0.0094	(0.009)
<b>4. Discharge location</b>				
Home	2,420,039	0.06	0.0070	(0.006)
Facility	1,098,419	0.35	0.0826***	(0.018)
Other	712,213	0.34	0.0042	(0.016)
<b>5. Health diagnosis for discharge to facility</b>				
Injury and infection	210,194	0.47	0.1194***	(0.043)
Cardiovascular	297,548	0.45	0.0399	(0.033)
Other	320,423	0.41	0.0582*	(0.031)
<i>B. Facility-level characteristics</i>				
<b>1. Urban indicator</b>				
Rural	870,421	0.20	0.0139	(0.012)
Urban	3,360,790	0.18	0.0260***	(0.010)
<b>2. log(Beds)</b>				
log(Beds) < median	1,322,962	0.17	0.0324**	(0.013)
log(Beds) >= median	2,908,277	0.19	0.0134	(0.009)
<b>3. Admits per bed</b>				
Admits per bed < median	604,084	0.23	0.0323	(0.022)
Admits per bed >= median	3,627,022	0.18	0.0156**	(0.008)

This table presents heterogeneity in the effects of PE ownership on patient mortality. Column 1 presents the sample size, and column 2 presents the corresponding mean mortality. Columns 3 and 4 present the corresponding coefficient  $\phi$  and its standard error obtained by estimating Equation (3) by 2SLS. The independent variable is the indicator for a patient being admitted to a PE nursing home, instrumented by differences in distance to the nearest non-PE and PE nursing home. The outcome variable is an indicator for patient death within 90 days of discharge from the index stay. Panel A explores the heterogeneity on several patient-level factors by dividing patients into 2 groups based on severity of preexisting comorbidities (high risk = Charlson index greater than 2) in row 1, gender in row 2, length of stay (above and below median) in row 3, patient discharge location in row 4, and the reason for discharge to facility in row 5. Panel B explores the heterogeneity on several facility-level factors by dividing facilities based on facilities being in Urban counties in row 1, facility size in row 2, facility admits per bed in row 3. All models include facility and patient HRR  $\times$  year fixed effects. We additionally control for the usual patient risk controls as in the main regression. Standard errors are clustered by facility. \*  $p < .1$ , \*\*  $p < .05$ , \*\*\*  $p < .01$ .

IV regressions (using Equation (3)) in which we adjust the dependent variable to be an indicator for the stay being longer than  $X$  days. Figure C.2, panel B, presents the estimated effects. We plot the  $X$  values on the  $x$ -axis and the coefficient from the model is plotted on the  $y$ -axis. For example, the coefficient for  $x = 15$  days implies a 4-pp increase in the probability of stays becoming

longer than 15 days following PE ownership. The figure documents that PE-owned facilities keep very short stayers longer, with little or no effect on stays becoming longer than 35 or more days. This would maximize Medicare revenue, because Medicare pays fully for only the first 20 days of care and then tapers off. Tying this to the mortality results above, it may be that extending short stays plays an important role in elevating mortality.<sup>25</sup>

Panel A, row 4, of Table 6 examines mortality effects for patients discharged to different destinations. The mortality effect is driven by patients discharged to facilities (predominantly hospitals), though the coefficients are positive for all destinations. Note that we do not find significant changes in the proportions of patients sent to the different destinations (see Table C.6, panel C). However, there could be changes in the composition of patients sent to different locations. In panel A, row 5, we disaggregate patients discharged to hospitals by the category of the reason for hospitalization and find the largest effect on mortality occurs for patients with an injury or infection, a finding that may be consistent with lower-quality care.

Table 6, panel B, further explores heterogeneity on three dimensions. First, we test for differences in the mortality effect between urban and rural counties. Row 1 shows that the mortality effect is substantially larger at facilities in urban counties, even though baseline mortality levels are a bit lower. This pattern is consistent with the finding above that the mortality effect is greater among less risky patients. Second, we examine heterogeneity by facility bed size. The results in row 2 show that the mortality effect is substantially larger at facilities below the median bed size (129 beds), implying that these facilities experience greater disruption to patient care due to PE ownership. Larger facilities may benefit from having access to better managerial personnel. Row 3 explores heterogeneity on a related dimension, the facility's patient throughput, which may partly reflect operational excellence. Indeed, we find that facilities admitting more patients per bed than the median on average (4 patients per bed) have a substantially smaller mortality effect.<sup>26</sup>

Apart from offering some insight into the variation in health effects, the analysis in panel B also helps assess the generalizability of the results beyond our sample to the population of U.S. nursing homes. As discussed in Section 2.2, the facilities in our sample are larger, admit more patients per bed, and are more likely to be in urban counties than the average nursing home in the United States. However, the heterogeneity results suggest that the nonrandom targeting of facilities by PE firms doesn't push the mortality effect consistently in one direction. While the disproportionate urban focus likely pushes up the

<sup>25</sup> We find an increase in length of stay only for low-risk patients (results not reported). Since we also find the mortality effect was greater among low-risk patients, this further finding lends support to the interpretation that extending length of stay is linked to elevated mortality.

<sup>26</sup> Intuitively, facilities above the median bed size or with higher occupancy rates treat many more patients in aggregate than those below the respective median values. This explains why the samples are not evenly split in panel B, rows 2 and 3.

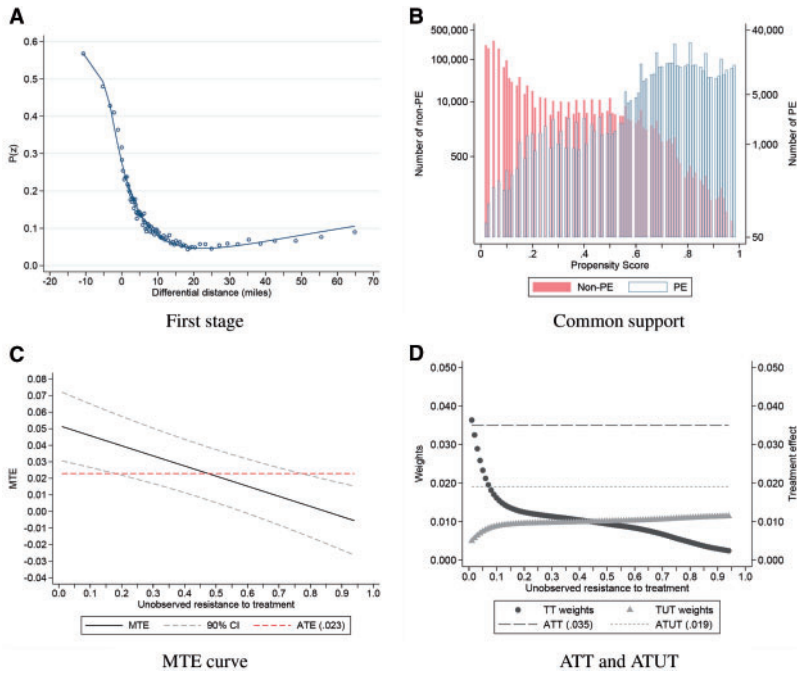
estimated mortality effect relative to the population average, the size and patient throughput characteristics go in the opposite direction.

**4.2.2 MTE approach** We explore whether the effect on mortality varies across complier patients with their propensity for treatment at a PE-owned facility, using the framework of a standard selection model and by estimating Marginal Treatment Effects (MTE). This allows us to test whether compliers experience a constant treatment effect or whether they sort into PE facilities based on treatment gains. Another benefit is that the MTE enables us to estimate parameters that are not specific to the complier group and make more general statements regarding the causal effects of PE ownership within the treated sample of facilities. Specifically, we can estimate parameters such as the Average Treatment Effect (ATE) and Average Treatment on the Treated (ATT) (Heckman and Vytlačil 2005).

Our MTE approach follows the estimation approach in Cornelissen et al. (2018). See Internet Appendix Section B.3 for the model and estimation details. Here, we note the two additional functional form assumptions required to extrapolate beyond complier patients and obtain treatment effects for other patient groups. Both assumptions are standard in the recent MTE literature (Cornelissen et al. 2016; Brinch et al. 2017). First, we assume that the MTE is additively separable into an observed and unobserved component. The observed component varies linearly with the covariates, and therefore by market, year, facility, and patient risk. The unobserved component depends only on the predicted probability of going to a PE facility—also known as the propensity of treatment,  $p$ —obtained from the first stage model. This assumption therefore restricts the unobserved component to be identical for all individuals at the same propensity of treatment, primarily based on their relative distance to a PE facility. Individuals who go to a PE facility despite having a low value of  $p$  are revealed to have low resistance to treatment, and vice-versa. Second, to further aid identification, we assume that the unobserved component has a polynomial functional form in  $p$ . Using a polynomial allows interpolation between different values of  $p$ . In our baseline model we assume the unobserved component of the MTE is linear in  $p$ , but we relax this restriction in robustness checks and find similar results. We obtain standard errors by bootstrap, clustering them by facility as in the LATE analysis.

Figure 5, panel A, presents the variation in the estimated propensity score. We collapse the data to percentiles of differential distance,  $D$ , and plot a nonparametric fit of  $p$  values against the corresponding percentile means of  $D$ . This shows a similar pattern first observed in Figure 3, namely, the probability of going to a PE-owned facility declines nearly monotonically as differential distance increases. However, this figure masks the full support of the distribution of  $p$ , which extends over the entire unit interval. Figure 5, panel B, highlights the overlap in the distributions of the propensity scores for treated (PE) and untreated (non-PE) patients by plotting histograms for the





**Figure 5**  
Marginal treatment effects

This figure presents results pertaining to a marginal treatment effects (MTE) analysis using the Medicare patient-level data. Panel A presents the ‘first-stage’ fit of predicted probability of treatment or propensity score, w.r.t the instrument. Panel B presents the overlap in distributions of PE and non-PE groups by propensity score. This plot uses a log scale due to the large number of non-PE patients with low propensity. Panel C presents the weights for the IV and corresponding estimates. Panel D presents the weights for the Average Treatment on the Treated (ATT) and Average Treatment on the Untreated (ATUT) and the corresponding estimates. Section 4.2 describes the MTE estimation.

two groups against  $p$  on the  $x$ -axis. The figure confirms that the treated and untreated groups overlap in distributions over nearly the entire unit interval.<sup>27</sup>

Figure 5, panel C, presents the estimated MTE curve and its 90% confidence intervals on the  $y$ -axis against the unobserved resistance to treatment,  $u$ , on the  $x$ -axis. The MTE curve slopes downward, implying reverse selection on treatment gains; that is, individuals with the least resistance to going to a PE facility experience the worst mortality effects. The slope is highly statistically significant ( $p < 0.01$ ), confirming reverse selection on gains. In contrast, individuals with the highest resistance are slightly better off at PE facilities. The figure also plots the ATE, which is 2.3 pp (SE = 1.0 pp).<sup>28</sup>

<sup>27</sup> We have at least 50 patients from both the treated and control groups at every percentile value of the propensity score over the range 0–0.94. We trim the sample to this range when estimating the MTE curve.

<sup>28</sup> Figure C.7, panel A, shows a specification check on the assumption of a linear MTE curve. We compute a weighted average MTE equivalent to the main 2SLS estimate and obtain a similar value (2.3 pp vs. 2 pp).

We aggregate the MTEs using the appropriate weights, shown in Figure 5, panel D, to obtain various treatment effect parameters. Given the downward sloping nature of the MTE curve, we expect the ATT to be higher than the ATUT, which is what we find. The ATE and ATT are statistically significant, while the ATUT is not.

In sum, the ATE implies that a randomly chosen Medicare patient from our sample would experience a 2.3-pp increase in the chance of short-term mortality if she goes to a PE-owned nursing home. This is comparable to the estimated LATE. It implies the estimated LATE is representative of the treatment effect for the average short-stay Medicare patient in the sample. Second, the MTE curve implies reverse selection on gains and that some patients – those with greater resistance to treatment – are unaffected and may even experience small mortality gains if they choose a PE-owned facility.

### 4.3 Patient well-being

If the effect on short-term mortality is related to lower patient welfare, we expect to see consistent evidence using other well-being measures. We focus on three clinical measures of well-being that CMS uses as outcomes for short-stays when computing five-star ratings (surprisingly, mortality is not one of them). These are patient mobility, developing ulcers, and increasing pain intensity. The OLS models find positive effects on all three outcomes, reported in Table 7, panel A. Relative to their respective means, there is a decrease in mobility of 6.2% (column 1), increase in ulcer development of 8.5% (column 2), and increase in pain intensity of 10.5% (column 3). The corresponding event studies, in Figure C.6, indicate no differential pre-trends and large, gradual increases following acquisition.

The IV models, in panel B of Table 7, show positive effects on two of the three outcomes. Mobility decreases by 3% of the mean (column 1), while pain increases by 8.3% of the mean (column 3). However, there is no effect on developing ulcers in the IV model (column 2). Overall, the evidence of harmful effects on other measures of patient well-being help confirm that the adverse effect on mortality is not a spurious finding.

## 5. Operational Changes

This section uses facility-level data to explore operational changes that could help explain the adverse patient welfare effects described above.

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suggesting little mis-specification error. We also estimate the MTE curve using higher-order polynomials. Figure C.7, panel B, shows that the curve remains downward sloping regardless of the polynomial.

**Table 7**  
**Patient well-being**

	A. OLS estimates		
	(1) 1(Mobility decreases)	(2) 1(Develops ulcers)	(3) 1(Pain intensity increases)
1(PE)	0.0321*** (0.003)	0.0068** (0.003)	0.0272*** (0.005)
Observations	4,231,356	4,231,356	4,231,356
Y-mean	0.52	0.08	0.26
	B. IV estimates		
	(1) 1(Mobility decreases)	(2) 1(Develops ulcers)	(3) 1(Pain intensity increases)
1(PE)	0.0152* (0.009)	-0.005 (0.007)	0.0216 (0.014)
Observations	4,231,356	4,231,356	4,231,356
Y-mean	0.52	0.08	0.26

This table presents estimates of the relationship between PE ownership and measures of patient well-being obtained from clinical assessments. Each cell in panel A presents the  $\phi$  obtained by estimating Equation (1), and each cell in panel B presents the IV version of the same results, obtained by estimating Equation (3). The independent variable is an indicator for the patient being admitted to a PE nursing home in panel A, and is instrumented by differences in distance to the nearest PE and non-PE facility in panel B. We present results for patient-level outcomes, an indicator for decrease in patient mobility, developing/worsening pressure ulcers, and increase in pain intensity. All models include facility and patient HRR  $\times$  Year fixed effects. We additionally control for the usual patient risk controls as in the main regression. Standard errors are clustered by facility. \*  $p < .1$ , \*\*  $p < .05$ , \*\*\*  $p < .01$ .

### 5.1 Empirical strategy

For outcomes available only at the nursing home level, we cannot instrument for patient selection so we use a facility-level version of the OLS differences-in-differences model, presented in Equation (4).

$$Y_{j,t} = \alpha_j + \alpha_t + \beta PE_{j,t} + P'_{j,t} \gamma_1 + M'_{j,t} \gamma_2 + \varepsilon_{j,t}. \tag{4}$$

$PE_{j,t}$  takes a value of one if facility  $j$  is PE-owned in year  $t$ . The coefficient of interest is  $\beta$ , which captures the relationship between PE ownership and the outcome  $Y_{j,t}$ . We include facility ( $\alpha_j$ ) and year fixed effects ( $\alpha_t$ ). The vector  $P_{j,t}$  includes three controls for facility-level patient mix and  $M_{j,t}$  includes five county-level controls for time-varying market attributes.<sup>29</sup> As there may be a concern that control variables could be affected by PE ownership, we also present results without any controls.

<sup>29</sup> The patient-level controls are as follows: The Case Mix index (CMI) is a composite measure of patient risk based on medical history of diagnosis or treatment for a large number of conditions. Second, the Acuity index is a measure of patient risk computed using the patient's assessed activities of daily living (ADL) scores. In both cases, a greater value indicates a riskier patient cohort for the nursing home. We winsorize both the CMI and Acuity index at the 1% and 99% levels in each year. The third control is the share of the facility's patients who are Black. The county-level controls are as follows: The Herfindahl Hirschman index (HHI) is based on the shares of beds, the number of for-profits, the number of chain-owned, the number of hospital-based, and the number of overall facilities. These are calculated using a leave-one-out procedure from the facility-level data.

The identifying assumption is that PE targets and control facilities would continue on parallel trends in the absence of the acquisition. We test for differential pre-trends using event study figures, which plot the coefficients  $\beta_s$  from Equation (5).

$$Y_{j,t} = \alpha_j + \alpha_t + \sum_{s \neq -1} \beta_s \text{DealYear}_{j,s} + P'_{j,t} \gamma_1 + M'_{j,t} \gamma_2 + \varepsilon_{j,t} \quad (5)$$

$\text{DealYear}_{j,s}$  is an indicator that is one in year  $s$  relative to the buyout year for facility  $j$ , and zero otherwise. The remaining terms are as defined above for Equation (4). Finally, as above, we show robustness to the Callaway and Sant'Anna (2021) and Sun and Abraham (2021) estimators for these facility-level models.

## 5.2 Results

We consider three types of operational channels. The first two explicitly concern facility quality, while the last pertains to financial strategies particular to the PE industry. The main results are presented in Table 8. For each outcome, the top row of coefficients is from specifications with only facility and year fixed effects, while the bottom row adds the full set of patient and market controls. Figures 6 and 7 handle the corresponding event studies.

**5.2.1 Compliance with standards and staff availability** First, we consider compliance with care protocols in panel A of Table 8. Our outcome of interest is the facility-level Five-Star rating, which varies from one (worst) to five (best). After PE buyouts, the Deficiency rating declines by 0.08 points (column 1), which is about 3% of the mean and 7% of the standard deviation (the most relevant measure given how this variable is constructed). This rating reflects whether the facility is satisfying care protocols such as storing and labeling drugs properly, disinfecting surfaces, as well as other aspects of care such as ensuring resident rights and avoiding patient abuse. The Overall rating similarly declines (column 2). Figure 6 presents event studies for each outcome. There are no pre-trends, consistent with the identifying assumption, and the negative effects appear immediately after the change in ownership and persist for at least 5 years.<sup>30</sup>

Intensive subsidy is one reason high-powered profit maximizing incentives may lead to adverse effects in the nursing home context. While we cannot randomly allocate subsidy intensity across facilities, for these compliance outcomes we can assess whether facilities that rely more on Medicare for revenue experience more adverse effects. Indeed, Table C.7 shows that the

<sup>30</sup> The Overall rating has three components: the Deficiency rating, a Quality rating based on metrics computed using claims data and clinical assessments, and a Staffing rating, which is based on staffing measures evaluated in panel B. Since we assess quality and staffing changes more granularly, we do not present the effects on these components, but we find negative, significant effects of equal or larger magnitudes there as well.

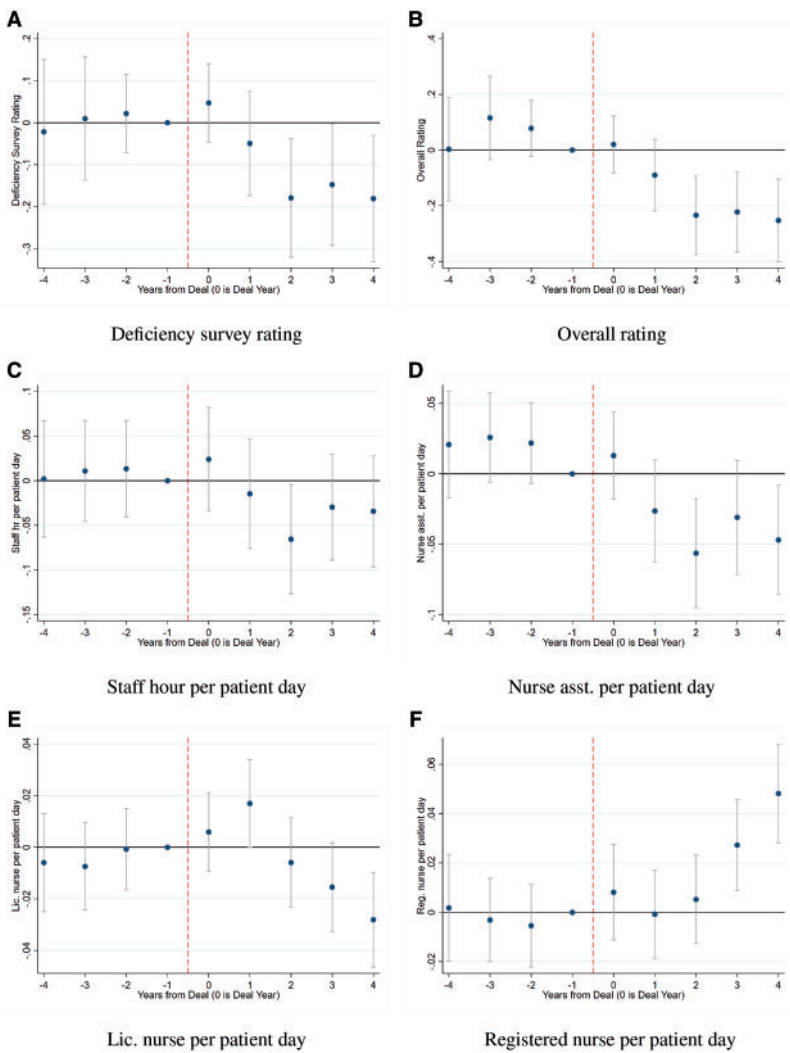
**Table 8**  
**Mechanisms and operational changes**

A. Five-star rating				
	(1) Deficiency rating	(2) Overall rating		
1(PE) (No control)	-0.077** (0.037)	-0.083** (0.036)		
1(PE) (With control)	-0.080** (0.037)	-0.086** (0.036)		
Observations	111,728	111,728		
Y-mean	2.8	3.1		
B. Staff per patient day				
	(1) All staff	(2) Nurse assistant	(3) Licensed nurse	(4) Registered nurse
1(PE) (No control)	-0.046*** (0.017)	-0.062*** (0.010)	-0.024*** (0.006)	0.039*** (0.005)
1(PE) (With control)	-0.042** (0.017)	-0.060*** (0.010)	-0.023*** (0.006)	0.040*** (0.005)
Observations	227,254	227,254	227,254	227,254
Y-mean	3.4	2.2	0.8	0.4
C. log(Financials)				
	(1) Management fee	(2) Building lease	(3) Interest expense	
1(PE) (No control)	0.056* (0.032)	0.570*** (0.061)	1.179*** (0.096)	
1(PE) (With control)	0.056* (0.032)	0.570*** (0.061)	1.174*** (0.096)	
Observations	196,880	196,900	196,918	
Y-mean	0.2	0.4	0.3	

This table presents estimates of the relationship between PE ownership and nursing home outcomes. Each cell presents the coefficient  $\beta$  obtained by estimating equation (4) with a different outcome. The independent variable is an indicator for whether a nursing home is PE-owned (=1 if PE-owned, 0 otherwise) starting in the next year from the deal announcement date. Panel A presents results for quality outcomes as measured by the Five-Star rating awarded by CMS on overall rating and deficiencies identified by independent contractors in audits, respectively. A negative effect on ratings implies a decline in quality. Panel B presents results on per patient nurse availability for all nurses, nurse assistants, licensed nurses, and registered nurses. Panel C presents results on the logarithm of management fees, building lease cost, and interest expenses. The top row presents results with no controls. The bottom row presents the results including controls, which consist of market-level and patient mix controls, as described in Section 5.1. All models include facility and year fixed effects. All variables are winsorized at the 1% and 99% levels. Standard errors are clustered by facility. \*  $p < .1$ , \*\*  $p < .05$ , \*\*\*  $p < .01$ .

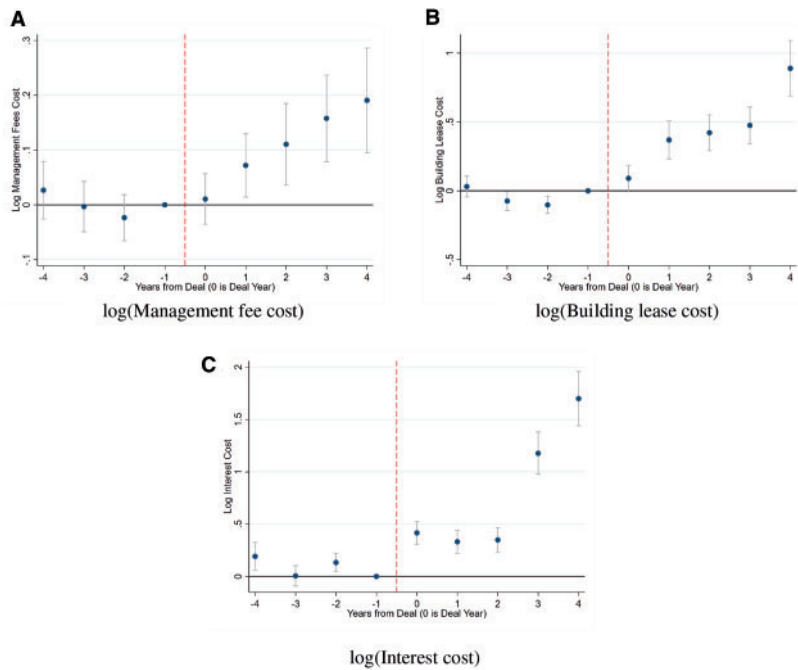
negative effects on the two rating measures are *entirely* driven by facilities with above-median Medicare revenue. For example, the negative effect on the Deficiency rating is  $-0.12$  for this group, or about 4.3% of the mean and 9.7% of a standard deviation.<sup>31</sup>

<sup>31</sup> We restrict this analysis for these outcomes because they are facility-wide and do not depend on the type of patient, while the other splits (staff, mortality) reflect the composition of patients and the requirements of different payers. For example, Medicare patients require more RNs. In unreported analysis, we do not find differential effects in our other outcome measures.



**Figure 6**  
**Aggregate quality and staffing outcomes**

This figure presents event studies on quality of care measures (Five-Star ratings) and Staffing around the time a nursing home experiences a PE buyout. Each point in the figures represents the coefficient  $\beta_s$  obtained by estimating Equation (5) as discussed in Section 5. The model is fully saturated, including coefficients for all years, except year  $-1$ , which is the omitted point, but we plot only years  $-4$  to  $4$ . Panels A and B present effects on the Five-Star ratings awarded by CMS - deficiencies identified by independent contractors in audits and overall rating, respectively. A negative effect on ratings implies a decline in quality. Panels C to F present results on nurse staffing per-patient for all staff, nurse assistants, licensed nurses, and registered nurses respectively. All models include facility and year fixed effects, as well as patient mix and market controls, which are described in Section 5.1. All dependent variables are winsorized at the 1% and 99% levels. Standard errors are clustered by facility.



**Figure 7**  
**Facility finances**

This figure presents event studies on facility finances around the time a nursing home experiences a PE buyout. Each point in the figures represents the coefficient  $\beta_s$  obtained by estimating Equation (5) as discussed in Section 5. The model is fully saturated, including coefficients for all years, except Year = -1, which is the omitted point, but we plot only years -4 to 4. Panels A to C present results on the logarithm of the management fee cost, building lease cost, and interest cost, respectively. All models include facility and year fixed effects, patient mix, and market controls, as described in Section 5.1. All dependent variables are winsorized at the 1% and 99% levels. Standard errors are clustered by facility.

In Table 8, panel B, we assess effects on nursing staff hours per patient-day, a well-established measure of nursing home quality that accounts for changes in patient volume. Column 1 shows a modest decline of 0.04 hours in aggregate staff hours (1.2% of the mean). This masks larger changes for different types of nurses that offset each other. A decrease in “frontline” caregivers (CNAs and LPNs) is shown in columns 2 and 3, respectively. Together, there is a decline of around 0.08 hours for these two groups (2.8% of the mean). In contrast, there is an increase in the use of Registered Nurses (RNs) by about 0.04 hours (10%). The event studies in Figure 6 again reveal no pre-trends and show immediate declines after the deal in frontline staffing, while the increase in RN staffing appears in the third year after the buyout. The increase in RN staff hours does not compensate for the decline in lower-skilled nurse hours because RNs account for a small fraction of all staff hours. Medicare cost reports indicate that CNAs and LPNs receive an hourly wage that is about 40% and 70% respectively of the wage paid to RNs, which is around \$35 per

hour. Data on nurse salary costs is available only 2011 onward. Using this information, we estimate a statistically significant decline in total nurse cost of 2–2.5% (SE, 0.8%).<sup>32</sup>

We perform multiple robustness tests such as including controls for chains and excluding the top-two deals (Table C.8). The coefficients are similar. The alternative OLS Callaway and Sant’Anna (2021) and Sun and Abraham (2021) estimators are in Figure C.5, panel B. Following Goodman-Bacon (2021), we show that most of the DD estimator’s weight is on the comparison between treated and never-PE facilities (Table C.9).

Using estimates from the literature for the effect of frontline nurse availability on mortality, we calculate that the estimated decline in frontline nurse staffing predicts an increase in mortality of 0.24 pp.<sup>33</sup> It is intuitive that lower staffing—in particular low-skill staffing—would be associated with increases in adverse conditions related to lack of attention, such as lower mobility and higher pain intensity. Higher RN availability is consistent with less complex patients driving the mortality effects. RNs are responsible for supervision and treatment decisions, while frontline nurses support activities of daily living, such as preventing infections. Managers may have looked for ways to cut overall labor costs while changing the mix of nursing staff capability to maintain quality and patient experience, as RNs are crucial to nursing home quality (Zhang and Grabowski 2004; Lin 2014). An alternative explanation is the regulatory focus on RNs. For example, CMS uses the availability of RNs to determine eligibility for Medicare reimbursement.<sup>34</sup> Given the tight regulatory scrutiny of RN availability, it is difficult to reduce staffing levels in this category.

To explore the relationship between declines in staff availability and quality, we compare changes in nurse availability and Five-Star ratings within target facilities around the PE buyout event. This recovers correlations and does not imply causality, so we present the raw data in bin-scatter plots. Figure C.8 shows the change in Five-Star rating over the 3 years around PE acquisition on the  $y$ -axis against the change in aggregate staff hours per patient day during the same period on the  $x$ -axis. The plots show that facilities which experienced larger declines in nurse availability also experienced meaningfully greater

<sup>32</sup> This includes salaries and payments to contracted nurses. Given the data limitation, we cannot perform a more thorough analysis on hourly wages and total nurse wages comparable to the analysis on nurse hours. We also do not observe whether facilities take other cost reduction steps such as using more part-time labor and reducing individual shifts. So this reduction in total wages may understate the true change.

<sup>33</sup> Tong (2011) exploits an increase in minimum nurse staffing regulation in California and finds a decline in on-site patient mortality due to greater availability of frontline nurses. Specifically, Tong (2011) reports a 15% decline in mortality due to an increase in nurse availability of one hour per resident-day. Since we estimate a decline of 0.083 hours, this predicts an increase of  $0.083 \times 15 = 1.25\%$  of the mean, or 0.24 pp. More recently, Ruffini (2022) exploits variation in minimum wage requirements to isolate the effects of nurse staffing changes on quality and also finds mortality effects.

<sup>34</sup> Specifically, such facilities are defined by having “an RN for 8 consecutive hours a day, 7 days a week (more than 40 hours a week), and that there be an RN designated as Director of Nursing on a full time basis” (see <https://www.cms.gov/Regulations-and-Guidance/Guidance/Manuals/Downloads/som107c07pdf.pdf>).



declines in ratings. The patterns are consistent across rating types and suggest that cuts to nursing staff may be an important channel to explain the quality declines.

**5.2.2 Finances and operations** Our final analysis uses CMS cost reports to analyze key sources of expenditure related to the PE business model. We begin by noting that nursing homes are widely known to have relatively low and regulated profit margins, often cited at just 1%–2%.<sup>35</sup> Our data on nursing home cost reports submitted to CMS indicate that nursing homes report negative operating margins on average, and PE-owned nursing homes are not on average more profitable. In unreported analysis, we find no effect of buyouts on net income or overall revenue. This raises the question of how PE firms create value from nursing home investments.

The academic literature and popular press associate three types of firm expenditures with the PE playbook. The first is often termed “monitoring fees”, which are charged to portfolio companies (in the CMS cost reports, these are listed as “management fees”).<sup>36</sup> *Metrick and Yasuda (2010)* note that these are thought to be between 1% and 5% of EBITDA. Our data suggest that they increase over time after buyouts, as shown in Figure 7, panel A, where the fees are flat before the buyout, and then rise dramatically afterward. Table 8, panel C, column 1, indicates that on average, management fees increase by 5.7% after acquisition (we exponentiate large coefficients since the outcome is in log dollars).

The second type of expenditure is lease payments. The value of real estate is one reason that nursing homes and other typically low-margin assets can be profitable investments; the investor can sell the real estate to a related company or to a third party (*Dixon 2007; Keating and Whoriskey 2018; Brown 2019*). Cash from the real estate sale can be disbursed as profits to the PE fund. A cash inflow early in the life of the investment is especially beneficial to the fund’s Internal Rate of Return, a key industry performance metric. The nursing home assumes the obligation of future rent payments. As an example, a *New York Times* report on the nursing home industry notes that:

“[PE] investors created new companies to hold the real estate assets because the buildings were more valuable than the businesses themselves, especially with fewer nursing homes being built. Sometimes, investors would buy a nursing home from an operator only to lease back the building and

<sup>35</sup> [https://www.medpac.gov/wp-content/uploads/import\\_data/scrape\\_files/docs/default-source/reports/mar17\\_medpac\\_ch8.pdf](https://www.medpac.gov/wp-content/uploads/import_data/scrape_files/docs/default-source/reports/mar17_medpac_ch8.pdf)

<sup>36</sup> In their summary of buyout fund economics, *Metrick and Yasuda (2010)* write that “we think of monitoring fees as just another way for BO funds to earn a revenue stream.” These fees should not be confused with the usual 2% of fund value that General Partners earn each year for managing Limited Partners’ capital, before profits from investments.

charge the operator hefty management and consulting fees” (Goldstein, Silver-Greenberg, and Gebeloff 2020).<sup>37</sup>

Consistent with this strategy, column 2 shows that facility building lease payments increase dramatically by about 77% after PE acquisitions. Figure 7, panel B, confirms the lack of pre-trends and the increase post-buyout.

The third type of expenditure is interest on debt. While not a direct source of PE profits, debt is tightly related to the overall PE model for creating value. [Metrick and Yasuda \(2010\)](#) note that the ratio of debt to equity in a buyout deal is typically around 5:1. The interest payments become a cost to the portfolio company. In Figure 7, panel C, we see that like the previous two outcomes, interest payments are flat before the buyout and then rise dramatically afterward. Column 3 indicates that the increase is about 224%.

Taking the results on nurse availability together with the estimated effects on interest, lease, and management fees payments, we infer that PE ownership shifts operating costs away from staffing inputs toward costs that are profit drivers for the PE fund. To our knowledge, this paper offers the first instance in the literature on PE in which these three profit drivers have been documented systematically.

The final outcome we explore is patient capacity and volume. [Table C.10](#), column 1, shows that beds remain nearly unchanged, which may partly reflect state regulations restricting expansions. Admissions increase by 3.3%, or about six patients per year, for the average facility (column 2). The event study, in [Figure C.9B](#), shows an increasing trend post-acquisition. The apparent disconnect between demand and quality of care may reflect information frictions in observing nursing home quality, as discussed earlier ([Arrow 1963](#); [Grabowski and Town 2011](#); [Werner et al. 2012](#)). Higher admissions raise the question of whether PE ownership increases overall access to nursing home care, providing care for individuals who would not otherwise have gone to a nursing home. To test whether this is the case, we assess the effects of PE entry into a nursing home market, using the HRR definition. [Table C.10](#), column 3, shows no effect of initial PE entry on admissions at the market level, corroborated by flat patterns in the event study ([Figure C.9C](#)). Hence, the increase in facility-level admissions likely reflects business stealing.

To explore what makes PE special relative to other types of private ownership, we conduct three heterogeneity tests that focus on particular elements of the PE business model. First, we consider whether funds that specialize in healthcare lead to larger effects on mortality. If an important channel for our results is the application of the General Partner’s business

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<sup>37</sup> Two examples further illuminate these types of transaction. First, the HCR Manorcare deal discussed in Section 1.2, where the chain’s real estate assets were spun off and sold shortly after the acquisition by the Carlyle Group. Second, at a Congressional hearing the executive director of the Long-Term Care Community Coalition said “more and more with entities buying up nursing homes, they have no experience in the business, they sell out the underlying property” ([Brown 2019](#)).

expertise to their investments, then we expect specialist funds to have more operational impact, for example, conducting cost reduction behaviors more aggressively. Using Pitchbook's classification of each fund's industry focus, we separate the sample into Generalist and Specialist funds. Of 124 funds, 58 are specialists. We interact the indicator for PE ownership with an indicator for being a specialist in healthcare. The results, in [Table C.11](#), show that the negative effects on both the overall quality rating and nurse availability are driven by specialist funds. This is consistent with them having greater industry-specific expertise and imposing larger operational changes than generalist funds.

Second, we ask whether deals in which the facility had greater financial liabilities at baseline, that is, prior to buyout, are also the ones with more negative staffing/ratings changes.<sup>38</sup> This would be the case if the reductions in nurses and other inputs are part of a response to greater financial stress. We interact the indicator for PE ownership with reported liabilities per bed, to control for differences in facility size. The interaction, in the second group of coefficients in [Table C.11](#), again loads negatively for both rating and staffing. This is consistent with financial burden as an important channel through which PE ownership affects nursing home inputs and quality of care.

Last, we explore whether the fund age is related to the changes in quality. If a short time horizon is a primary driver for the negative effect, we might expect to find larger results when the fund is older and the pressure to deliver returns to investors quickly is more intense. This test comes with a caveat, since in general PE ownership is substantially shorter-term than other types of private ownership, so this extensive margin could drive the results regardless of the point in the fund lifecycle at which the investment occurs. Here, we interact the indicator for PE ownership with the fund age at the time of the deal. While the interaction coefficients are negative for both rating and staffing, they are not statistically significant (bottom group of [Table C.11](#)). This suggests that a short time horizon may help explain PE effects in general but does not lead to a strong intensive margin result among PE funds.

## 6. Conclusion

This paper studies PE buyouts in healthcare, an important sector where PE activity has increased and generated policy debate. We find that going to a PE-owned nursing home significantly increases short-term mortality. The amount depends on patient composition. OLS results suggest an effect of about 2%, while an IV approach suggests a much larger effect of 11% within the relatively vulnerable population that goes to the closest facility. A marginal treatment effects analysis suggests that the LATE is representative of the effect for the

<sup>38</sup> Liabilities are reported in the nursing home cost report files in worksheet G, line 60. These include all current liabilities, long-term liabilities, and total fund balances.

average Medicare patient in our sample. We document multiple changes in clinical and operational factors that help explain the increased mortality and spending. We also find a corresponding increase in operating costs that tend to drive profits for PE funds.

Our results suggest that, on average, PE acquisitions of nursing facilities lead to adverse health outcomes for some patients. While outcomes deteriorate post-acquisition, we caution that possible policy recommendations such as restricting acquisitions in healthcare require careful analysis and consideration of factors beyond the scope of our study. The overall incentive structure of an industry likely affects how acquisitions affect patient outcomes. Also, the fact that buyouts are not random may lead to heterogeneous effects for different types of nursing homes. A large body of work has shown that PE acquisitions lead to profitability improvements, and if policy makers can align the incentives of investors and consumers, it is possible that this will lead to better patient outcomes. Moreover, restrictions on PE buyouts may affect the incentives that owners have to start nursing home companies in the first place. Future work should focus on these trade-offs and possible solutions, and optimal regulation and incentive schemes for investors in healthcare.

Many additional channels could be investigated by future research. Although our results imply that PE ownership reduces productivity in the nursing home context, it may well have positive effects in other healthcare sectors with better-functioning markets (La Forgia and Bodner 2023). Beyond healthcare, significant PE investment has been injected in education, defense, and infrastructure, among other sectors, that also feature high levels of government subsidy and opaque product quality. Further work can be done to help design government programs to align the interests of PE-owned firms with those of taxpayers and consumers.

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