# Patient vs. Provider Incentives in Long Term Care*

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**Abstract**

We study how patient and provider incentives affect the mode and cost of long term care. Our analysis of 1 million nursing home stays yields four main insights. First, Medicaid-covered residents prolong their stays instead of transitioning to community-based care due to limited cost-sharing. Second, nursing homes shorten Medicaid stays when capacity binds to admit more profitable residents who pay out-of-pocket. Third, longer stays for marginal Medicaid beneficiaries do not improve health outcomes on average, pointing to annual overspending of $0.9bn. Fourth, transitioning from per-diem to episode-based provider reimbursement is more effective than increasing resident cost-sharing in shortening Medicaid stays.

**Keywords:** Long Term Care, Nursing Homes, Patient Incentives, Episode Based Reimbursement, Medicaid.

**JEL Codes:** H51, H75, I11, I13, I18, J14.

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# 1 Introduction

Long term care (LTC) expenditures are high and rising. In 2013, U.S. LTC spending accounted for $310 billion or 1.8% of U.S. GDP. As the population share aged 85 and older will more than double by 2050, LTC expenditures may increase twofold (Congressional Budget Office, 2013). This spending outlook critically depends on whether public policies can align incentives for patients and providers to efficiently utilize different forms of LTC services. Since more than 50% of LTC expenditures are covered by Medicaid, developing and expanding cost-effective home and community alternatives to expensive nursing home care is of high policy priority for many state Medicaid programs (Kaiser Family Foundation, 2015). Despite their significance for public spending and public health, empirical evidence on the effect of Medicaid policies on LTC utilization remains mixed and incomplete. In particular, we are not aware of any work that *jointly* considers how Medicaid policies affect LTC utilization through patient and provider incentives. Separating these two channels is clearly important for the design of policies that target patient or provider incentives.

In this paper, we shed new light on these questions by assessing whether patient or provider incentives play a larger role in determining LTC utilization in skilled nursing facilities (SNF). To this end, we develop a theoretical model of nursing home discharges to the community that nests patient and provider incentives. We test its reduced-form predictions using plausibly exogenous variation in cost-sharing and profit incentives for nursing home utilization. We then estimate a structural model to evaluate the relative importance of patient and provider incentives and to conduct policy counterfactuals.

Medicaid policies affect financial incentives for LTC utilization on both market sides. Medicaid beneficiaries are incentivized to prolong their SNF stays, which are covered indefinitely without requiring a co-payment. In contrast, Medicaid support for community based LTC, which includes services such as home health aides and adult day care, is less generous, particularly in our sample period from 2000 to 2005. On the provider side, Medicaid reimburses nursing homes on a daily 'per-diem' basis. This gives providers an incentive to prolong the stays of these residents, which increases Medicaid revenues. It is therefore crucial to account for provider and patient incentives when analyzing what drives SNF utilization among Medicaid beneficiaries.

Our analysis focusses on understanding the effect of patient and provider incentives on the timing of nursing home discharges to the community. This particular margin is relevant for three reasons. First, more than 40% of nursing home stays end with a discharge to the community, suggesting that community based care is a feasible alternative for a significant fraction of residents.[1] The precise timing of discharges is largely at the discretion of the nursing home discharge manager and the patient (or her relatives), so it is plausible that economic incentives affect the length of stay at this margin.

Second, we can exploit detailed resident micro data from the Long Term Care Minimum Data Set (MDS) combined with Medicaid and Medicare claims data. Our data provide detailed information on the universe of nursing home residents in California, New Jersey, Ohio, and Pennsylvania from 2000 to 2005. This includes information on admission, discharge dates and discharge reason, the payer source at any time during the stay, as well as health status information at admission, discharge, and different times during the stay. To isolate the role of financial incentives on resident discharges, we focus on a relatively homogenous population of residents who pay out-of-pocket at the beginning of their nursing home stay. Our sample contains information on one million stays and about 15.4 million resident-week observations. About 10% of these residents spend down their assets during their stay and become eligible for Medicaid, and 45% return to the community.

Third, we are able to exploit plausibly exogenous variation in patient and provider incentives in this context. To assess the role of provider incentives, we exploit a novel source of short-term variation in occupancy, i.e. the fraction of filled beds, which affects the nursing home's incentive to prolong the stays of Medicaid beneficiaries. At low occupancy rates, nursing homes benefit from longer Medicaid stays if the per-diem reimbursement rate exceeds the marginal cost. At high occupancy rates, when the capacity constraint is binding, nursing homes can benefit from discharging Medicaid beneficiaries in order to admit more profitable new residents who pay out-of-pocket. These residents pay the private rate set by the nursing home, which exceeds the Medicaid reimbursement rate by about 20% on average. By using week-to-week changes in occupancy, we control for other factors, such as the private rate or the quality of care, that may affect length of stay but vary on a less frequent basis. Specifically, our rich data allow us to control for nursing home-year and week-of-stay fixed effects and differences in observable health at admission. Therefore, we implicitly compare residents in the same nursing home, year, week of stay, and with similar LTC needs. On the patient side, we exploit variation in out-of-pocket prices among residents who spend down their assets and transition from out-of-pocket pay to Medicaid during their nursing home stay. Using the detailed Medicaid claims data, we can identify the exact timing of the Medicaid transitions and test for a reduction in the weekly discharge rate immediately after the start of Medicaid coverage.

Our empirical findings confirm the role of both patient and provider financial incentives

[1] Residents who are discharged to the community have shorter stays on average. Weighted by the length of stay, only about 8% of seniors are discharged to the community.

for discharges to the community. At low occupancies, weekly Medicaid discharge rates are less than half as large as private discharge rates, suggesting that patient financial incentives affect the length of stay. As the occupancy rate increases towards full capacity, we see an economically and statistically significant increase in the discharge probability for Medicaid beneficiaries but no effect for residents who pay out-of-pocket, suggesting that provider incentives also influence the length of stay.

These results pass an extensive list of robustness checks. To provide further evidence for the role of patient incentives, we refine our analysis using a propensity score weighting approach and a border analysis, which exploits differences in Medicaid rules between states. Both approaches yield qualitatively similar results. We also find that reductions in the discharge rates for Medicaid beneficiaries are largest immediately after a resident qualifies for Medicaid and in nursing homes with higher private rates. To corroborate the role of provider incentives, we show that the increases in Medicaid discharge rates at high occupancies are more pronounced in nursing homes with larger differences between the private and the Medicaid rate.

Concerning the effects of nursing home utilization on patient health outcomes, we find that the marginal residents are relatively healthy and have low LTC needs. Further, we find no evidence for increases in hospitalization or mortality rates or worsened health status at discharge due to shorter nursing home stays. These findings suggest that Medicaid beneficiaries are not discharged "too early" at high occupancy rates but point towards substantial overspending. Comparing Medicaid spending on SNF care to the cost of home and community based services (including opportunity costs of informal caregivers), we find that overall LTC spending could be reduced by at least $0.9 billion per year if Medicaid-covered stays were reduced to the average length of stay among residents who pay out-of-pocket, or by about 5.4 weeks on average.

To quantify the relative importance of patient and provider incentives we estimate a dynamic model of nursing homes' discharge behavior, which nests residents' cost sharing incentives. Our parameter estimates imply that both residents' and nursing homes' incentives lead to prolonged stays among Medicaid residents, but that providers react about twice as elastically to discharge incentives. Since shorter SNF stays do not worsen the health of residents who are at the home discharge margin, policymakers can lower LTC spending by changing provider incentives without harming nursing home residents.

We revisit the role of patient and provider incentives in several policy counterfactuals. In contrast to existing policies that predominantly affect patient incentives, we are able to simulate counterfactuals that also affect provider incentives. Our simulation results show that a LTC voucher system, under which Medicaid beneficiaries pay the full private SNF rate and providers receive the same fee for all payer types, reduces length of stay by about 19%. This policy only lowers overall LTC expenditures by \$0.03 billion per year, however, due to the lump-sum transfer to seniors that pays for the voucher. We also simulate the effect of transitioning from per-diem Medicaid reimbursement to an episode based reimbursement model. Specifically, we reduce the daily Medicaid rate by a relatively modest 1% and compensate nursing homes through a fixed up-front payment. This policy does not change the profitability of Medicaid residents but shortens their length of stay by about 5% and yields annual LTC cost savings of at least \$0.09 billion. We also find that transitioning only 3.2% of per-diem reimbursements to an episode-based (up-front) reimbursement is as effective as the voucher program in shortening the length of Medicaid stays but yields substantially larger annual total cost savings of about \$0.34 billion. Hence, we conclude that transitioning to an episode-based reimbursement model is more effective than increasing resident cost-sharing in shortening the length of Medicaid stays.

**Related Literature:** Our findings contribute to several literatures. First, we contribute to the literature on patient moral hazard in LTC, which is surprisingly sparse given the large amount of public funding of nursing home stays through Medicaid. Among the few existing studies, McKnight (2006) and Grabowski and Gruber (2007) find that the decision to enter a nursing home is relatively inelastic with respect to Medicaid cost-sharing incentives, whereas Mommaerts (2017) finds heterogeneous effects of spend-down requirements on SNF entry.[2] In contrast, Konetzka et al. (2014) find that private LTC insurance increases the propensity of nursing home stays. Importantly, very little is known about how patient incentives affect the length of nursing home stays. Our findings provide new evidence on this economically and politically important margin.[3] In addition, we provide new evidence on the effect of public health insurance on access to health care and health outcomes among an elderly and vulnerable population (see, e.g., Card, Dobkin, and Maestas, 2009).

Second, to the best of our knowledge, we provide the first evidence on the causal effect of provider profit incentives on the length of nursing home stays.[4] In two descriptive studies,

[2]In addition, an older literature evaluated different long term care interventions from the early 1980s (known as the Channeling demonstration), which sought to substitute community care for nursing home care. Most studies found that the community care interventions had relatively small effects on nursing home utilization suggesting very little substitutability among community and nursing home care, see Rabiner, Stearns, and Mutran (1994) for a review

[3]Our findings are supported by previous descriptive studies, which document that Medicaid beneficiaries are on average less likely to be discharge to the community, see Weissert and Scanlon (1985), Chapin et al. (1998), and Gassoumis et al. (2013). These studies lack a clean source of identifying variation and cannot explore the mechanisms reconciling this observation.

[4]Previous studies have investigated the effect of Medicaid bed-hold policies on the hospital discharges, see (Intrator et al., 2007). In contrast to the community discharges in our study, most residents return from the hospital to the SNF, so the implications for the length of nursing home stays are unclear in these studies.

Arling et al. (2011) and Holup et al. (2016) find that facilities with higher average occupancy rates are more likely to discharge residents to the community, but do not differentiate by payer type. Our findings are consistent with their observations and can also reconcile the negative cross-sectional relationship between nursing home occupancy and access to care for Medicaid beneficiaries, see Ching, Hayashi, and Wang (2015). The authors argue that capacity constrained nursing homes cream-skim against Medicaid beneficiaries at admission. In contrast, we document that nursing homes can manage their payer mix through endogenous discharges. While we cannot completely rule out potential cream-skimming at admission, we are less concerned about this channel in our context as we study seniors who pay initially out-of-pocket and are therefore relatively homogenous at admission.

Our findings complement existing evidence on the relationship between Medicaid reimbursement rates and the quality of nursing home care. More recent studies find that higher Medicaid reimbursement rates give nursing homes an incentive to increase their nurse-to-resident staffing ratios, which is a key input to the quality of nursing home care (Grabowski, 2001; Hackmann, 2017). Our analysis abstracts away from quality considerations and shifts the focus to the quantity of nursing home care. Our key result concerning the role of provider incentives is that holding the generosity of Medicaid reimbursement rates (and hence quality incentives) fixed, transitioning from per-diem to episode-based provider reimbursements provides significant provider incentives to shorten Medicaid stays.

This broader point also relates to the earlier supplier-induced demand literature (Fuchs, 1978; Green, 1978; Cromwell and Mitchell, 1986), which focused on other health care sectors and was subject to some criticism concerning the empirical methodology (Auster and Oaxaca, 1981; Dranove and Wehner, 1994; Gruber and Owings, 1996). We contribute to this literature by combining plausibly exogenous high-frequency occupancy variation with detailed micro data which allows us to identify the effect of provider incentives on health care utilization. Our source of variation is similar to the short-term variation in occupancy employed by Freedman (2016) who studies neonatal intensive care admissions.

Most closely related to our supply side analysis are two recent studies by Eliason et al. (2016) and Einav, Finkelstein, and Mahoney (2017). These studies find that a discrete jump in Medicare reimbursement changes provider incentives to discharge post-acute care patients from LTC hospitals. We deviate from these studies in several important ways. First, we focus on the design of Medicaid insurance in nursing home care. Second, we leverage detailed health records in the MDS, which allows us to construct a homogenous sample of residents and to identify residents at the margin for home discharges. We also present compelling evidence against negative health effects of shorter SNF stays by analyzing a wide range of measures of well-being in addition to hospitalization and mortality rates. Our findings point to no or even adverse health effects from longer stays, which is important for the normative implications. Third, we compare Medicaid spending for nursing home care to the cost of informal care in the community instead of focussing on institutional health care expenditures. Fourth, we exploit variation in occupancy instead of reimbursement rules as a source of financial provider incentives. The incentives originating from binding capacity constraints apply to other health care industries as well and have implications for the effectiveness of Certificate of Need (CON) laws, which restrict nursing home entry and capacity investment decisions. Finally, and perhaps most importantly, our framework incorporates both provider *and* patient incentives. This allows us to investigate how demand and supply-side cost sharing incentives jointly determine health care utilization.

Distinguishing between these two channels is particularly important for the design of public insurance programs, which affect the incentives for patients and providers, sometimes in opposing directions. While the significant implications for both market sides has been noted for a long time (see e.g., Ellis and McGuire (1993)), the existing empirical literature has largely studied the role of demand and supply-side cost sharing in isolation, see McGuire (2011) for an overview.[5]

The remainder of this paper is organized as follows. We describe institutional details on Medicaid and SNFs discharges in the next section. We then develop a theoretical model of provider and patient incentives and nursing home discharges in Section 3. We test the model predictions using the data described in Section 4 and the empirical strategy developed in Section 5. Section 6 presents first direct evidence on the role of patient and provider incentives. Motivated by this evidence, we then develop the structural model of resident discharges and discuss its identification and estimation in Section 7. Section 8 provides the structural estimation results and contains our counterfactual policy simulations. We conclude in Section 9.

# 2 Institutional Details

In this section, we describe select institutional features of the U.S. nursing home industry. We focus on Medicaid eligibility and provider reimbursement, and the discharge management of residents, which are most important for the subsequent analysis. For a brief overview of existing home and community based services (HCBS), see Appendix Section A.

[5] A notable exception is Trottmann, Zweifel, and Beck (2012), who study the impact of demand and supply-side cost sharing incentives in Swiss health insurance plans on health care utilization. Also, Dickstein (2015) studies patient and physician incentives in the market for antidepressants. In contrast to our results, Dickstein (2015) finds that more utilization leads to better health outcomes, so the normative implications are not as clearcut as in the case of SNF utilization that we study.

## 2.1 Medicaid Eligibility, Reimbursement, and Cost-Sharing

Nursing home care is largely financed by Medicaid, covering about 65% of all nursing home days. 25% of all days are funded privately with the majority being paid out-of-pocket. Only 16% of privately funded nursing home days (4% of all days) are paid for by private LTC insurance (Hackmann, 2017). The remaining 10% are covered by Medicare which only covers post-acute care of up to 100 days. Since the health profiles of residents whose stay is covered by Medicare are very different from Medicaid beneficiaries and private payers, we do not include them in the analysis.

In order to qualify for Medicaid, seniors cannot have assets above a state-specific resource limit of $2,000 to $4,000 and have to demonstrate LTC need due to medical reasons or functional limitations. States have different level-of-care criteria to determine Medicaid eligibility on medical grounds, but usually a state agency assesses potential beneficiaries for medical conditions that require skilled nursing care and limitations in activities of daily living (ADL).[6]

In addition, Medicaid eligibility is subject to an income test, and beneficiaries with income below the eligibility threshold do not face copayments for SNF care.[7] It is also possible for seniors above the income limit to qualify for Medicaid under so-called medically needy programs. Under these programs, nursing home residents who pass the asset test can deduct medical expenses, including SNF fees, from their income and qualify for Medicaid if their adjusted monthly income falls below a state-specific limit of about $400 to $600. In that case, the resident's entire income except for a personal needs allowance of about $50 is applied to the cost of nursing home care.[8] We abstract away from these co-payments in our baseline analysis which can account for at most 9% of the full nursing home bill.[9] We revisit the significance of co-payments for our main findings in the robustness check analysis in Section 6.3.

In practice, the asset test is the binding constraint for Medicaid eligibility. Using data from the Health and Retirement Study (HRS), we find that among seniors whose assets are below $2,000 and $4,000, respectively, only 1% have income levels that would make them ineligible for Medicaid under a medically needy program. Moreover, Borella, Nardi, and French (2017) find that among seniors in their 90s, up to 60% in the bottom permanent

[6] Of the states in our sample, California requires both ADL limitations and a medical condition that necessitates 24-hour supervision while one of the two requirements is sufficient in New Jersey and Ohio, and in Pennsylvania nursing needs are a necessary condition for Medicaid eligibility (see O'Keeffe, 1999).

[7] See, e.g., https://www.medicaid.gov/medicaid/ltss/institutional/nursing/index.html.

[8] See https://kaiserfamilyfoundation.files.wordpress.com/2014/07/medicallyneedy2003.pdf.

[9] We find an average monthly income of $625 among Medicaid beneficiaries using data from the National Long Term Care Survey. Considering a net allowance of $50 per month and an average private rate of $218 per day, this suggests that Medicaid beneficiaries pay at most $\frac{\$625-\$50}{\$218\times 30\text{days}} = 9\%$ of what private payers pay.

income tercile receive Medicaid benefits. In the middle permanent income tercile, over 20% become eligible for Medicaid as they age. This indicates that Medicaid coverage becomes widespread as seniors age and spend down their assets.

While some nursing home residents qualify for Medicaid at the beginning of their stay, our analysis focusses on residents who are initially *ineligible* for Medicaid because their assets exceed the eligibility threshold. About 10% of these residents spend down their assets during their stay and thereby switch from out-of-pocket pay to Medicaid coverage. Once residents have depleted their assets and become eligible for Medicaid, their price of nursing home care drops sharply. In Pennsylvania, for example, average private rates were $218 per day. At the same time, Medicaid beneficiaries who were discharged from a nursing home during our sample period (2000 to 2005) had to pay for home health care mostly out of pocket. Taken together, these incentives favor prolonged nursing home stays among Medicaid beneficiaries.

Medicaid pays nursing homes a regulated, risk-adjusted, daily reimbursement rate that is usually lower than the out-of-pocket (private) rate. In Pennsylvania, for example, Medicaid reimbursement rates averaged $188 in our sample period. At the same time, federal and state legislation, such as the Omnibus Budget Reconciliation Act of 1987, prohibits nursing homes from offering different quality of care levels by payer source. Grabowski, Gruber, and Angelelli (2008) find that SNFs comply with this regulation, so nursing homes generate lower profits per Medicaid resident than per private payer conditional on LTC needs. While nursing homes may adjust their quality one response to Medicaid reimbursement (Grabowski, 2001, 2004; Hackmann, 2017), we abstract from quality issues in this paper. Despite lower fees, Medicaid beneficiaries are generally profitable for nursing homes because reimbursement rates exceed the marginal cost of care (Hackmann, 2017).

Finally, several nursing homes cannot change their bed capacity at least in the short run because of fixed costs of investments but also because many states require nursing homes to obtain a CON in order to increase the number of beds.[10]

## 2.2 Nursing Home Discharges

Nursing home discharges occur in one of the following ways. If residents' health improves and their LTC needs decline, nursing homes may discharge them to the community. If their LTC needs do not change, the nursing home stay can end because the resident was discharged to a hospital, a different nursing home, or an assisted living facility, or because the resident passed

[10]Of the states in our sample, New Jersey, Ohio, and Pennsylvania had CON laws between 2000 and 2005 while California did not. Some states, including Pennsylvania, only limit the number of SNF beds that can be occupied by Medicaid beneficiaries.

away.[11] In our empirical analysis, we view the latter types of "non-community" discharges as "exogenous" in the sense that they are due to medical reasons, whereas community discharges may be affected by the SNFs and residents' financial incentives.

For seniors who have spent some time in a nursing home where they rely on around-the-clock care, transitioning into the community poses substantial challenges. Specifically, the management of complex medical conditions, support from family members or other informal caregivers, and housing that is adapted to the senior's medical conditions need to be arranged (see, e.g., Meador et al., 2011). Nursing home residents who prefer returning home and their relatives therefore have to exert a considerable effort before a discharge is possible.

From the nursing home's perspective, discharge decisions are made jointly with the resident and her family, and SNFs regularly evaluate their residents' health status to determine if they should remain in the facility or be discharged to the community. Nursing homes also assist residents who express a wish to return home with the above-mentioned tasks. According to discharge managers whom we interviewed, nursing homes have no systematic rules for when to discharge a resident, however. For example, discharge decisions are not tied to a certain case mix index (CMI) value or other objective health outcomes (see Appendix Section C for details on the CMI). While discharge managers asserted that current occupancy rates or waiting lists do not play a role in deciding a resident's discharge, no formal mechanism prevents such behavior.

Although federal regulations such as the Nursing Home Reform Law of 1987 prohibit involuntary discharges from nursing homes, Pipal (2012) argues that residents may not be aware of their rights and SNFs may stipulate the possibility of evictions in their admission agreements. Moreover, recent media coverage has highlighted nursing homes' propensity to discharge residents whose Medicare coverage runs out and who are eligible for Medicaid (Siegel Bernard and Pear, 2018). This shows that nursing homes base their discharge decisions at least partly on financial incentives since Medicare reimburses SNFs at higher rates than Medicaid. Overall, the timing of community discharges is largely at the discharge manager's and the resident's discretion. Therefore, we expect that economic incentives may have profound impacts on the length of stay for relatively healthy residents who can return to the community.

[11] Discharges due to hospitalizations are also affected by bed-hold policies. Since Medicaid reimburses nursing homes for keeping a bed vacant while a resident is hospitalized, SNFs may have a financial incentive to temporarily discharge Medicaid residents to a hospital (Intrator et al., 2007). We abstract from bed-hold policies and provide evidence that hospitalization discharges do not vary systematically with financial incentives arising from occupancy rate variation.

# 3 A Theoretical Model of Nursing Home Discharges

In this section, we sketch a theoretical framework that explains how financial incentives of SNFs and residents affect the timing of community discharges. We provide more details when we discuss the empirical version of the model in Section 7. We consider a single SNF and a single existing resident (the "focal" resident). The SNF maximizes profits and the resident trades off the utility of different care alternatives against the relevant out-of-pocket prices.

**Discharges and Effort:** In order to increase the probability of a discharge in any given week, the SNF and the resident have to exert costly effort, denoted by $e^{SNF} \geq 0$ and $e^{res} \geq 0$, respectively. Motivated by the institutional context and the observed discharged patterns, which we turn to in Section 6, we implicitly rule out "negative" efforts towards delaying the community discharge. The cost of effort includes, for example, the time and resources spent on finding alternative living and care arrangements in the community. We assume that the cost of effort for each agent, $c(e)$, is weakly positive, and strictly increasing and convex in effort. As a result, the SNF and the resident only exert an effort if they prefer a community discharge over staying in the nursing home for an extra period. The SNF and the resident choose their optimal efforts, $e^{SNF,*}(\cdot)$ and $e^{res,*}(\cdot)$, as a weakly increasing function of the financial benefits of discharges denoted by $\text{FinInc}^{SNF}$ and $\text{FinInc}^{res}$, for nursing homes and residents respectively. As a consequence, the discharge probability $Pr[D = 1]$ also weakly increases in $\text{FinInc}^{SNF}$ and $\text{FinInc}^{res}$:

$$D(\tau, oc) = \mathbf{1}\{\alpha \times e^{SNF,*}(\text{FinInc}^{SNF}(\tau, oc)) + \beta \times e^{res,*}(\text{FinInc}^{res}(\tau)) - \epsilon > 0\}.$$

Here, $\alpha \geq 0$ and $\beta \geq 0$ are scalars and capture the effect of financial incentives on nursing home discharges through the nursing home's or the resident's discharge effort. $\tau = P, M$ denotes the focal resident's payer type (private or Medicaid), and $oc$ is the SNF's occupancy rate in beds other than the focal resident's. We assume that the resident's financial discharge incentive and hence her effort are independent of $oc$. $\epsilon \sim F_\epsilon$ captures other factors that determine nursing home discharges, whereby financial incentives only increase discharges in expectation with

$$Pr[D = 1 | e^{SNF,*}, e^{res,*}] = F_\epsilon\Big(\alpha \times e^{SNF,*}(\text{FinInc}^{SNF}(\tau, oc)) + \beta \times e^{res,*}(\text{FinInc}^{res}(\tau))\Big). \quad (1)$$

In the case where $\alpha = 0$, only the resident's financial incentives matter, and if $\beta = 0$ only the SNF's financial incentives determine discharges. If $\alpha, \beta > 0$, the relative parameter magnitudes are important in assessing which agent responds more elastically to financial incentives.

**Financial Incentives and Discharges:** In order to specify how policy interventions affect the length of nursing home stays, we also require a model of the financial discharge incentives. In the case of the nursing home, we consider a dynamic tradeoff. If the focal bed is occupied, the nursing home receives a payer type specific flow profit $\Pi^{\tau}$ with $\Pi^{P} > \Pi^{M} > 0$. If the resident is discharged, the bed can remain empty, in which case the nursing home forgoes the flow payoff. However, with probability $\Phi(oc)$, the bed is filled with a new resident who may be a private payer or a Medicaid beneficiary. Therefore, the nursing home's optimal discharge effort is determined by the tradeoff between the flow payoff and the option value of drawing a more profitable payer type in the future.

Since private payers are more profitable than Medicaid beneficiaries, a nursing home will not exercise a costly discharge effort if the focal bed is filled with a private payer. Importantly, the refill probability $\Phi(oc)$ is weakly increasing in the occupancy rate in the nursing home's *other* beds: $\frac{\partial \Phi(oc)}{\partial oc} \geq 0$. Intuitively, the next arriving resident will seek the focal bed with probability one if all other beds are taken. If multiple beds are vacant, however, the probability of filling the focal bed is $< 1$. Therefore, the financial incentive as well as the optimal discharge effort are weakly increasing in $oc$ if the focal bed is filled with a Medicaid beneficiary. The occupancy rate in defined over *other* beds. We assume that the discharge manager, in charge of the focal resident, does not internalize her indirect effect on the occupancy in other beds. In other words, she takes changes in occupancy as given. This assumption simplifies the estimation of the model. We return to the endogenous equilibrium effects of discharge efforts on occupancy in the counterfactual analysis, see Section 8.2.2.

Turning to the resident's effort decision, we consider a static tradeoff. Staying another week yields the utility of nursing home care minus the co-pay. If the resident leaves the nursing home, she obtains the utility of home-based care minus home care payments. Both Medicaid beneficiaries and private payers pay for home care in full, but only private payers pay for nursing home care since the Medicaid co-pay is zero. Conditional on utilities from the two LTC options, private payers have a larger financial discharge incentive and therefore, they exert more discharge effort. The lower discharge effort among Medicaid beneficiaries leads to longer nursing home stays and constitutes patient moral hazard.

We abstract away from strategic free-riding of residents and their relatives on provider effort. We motivate this simplifying assumption by the presence of asymmetric information over the weekly occupancy rate. Building on the limited number of visits by relatives, we assume that they do not observe the weekly occupancy rate and cannot condition their effort on occupancy accordingly. Mechanically, we assume a constant return to resident effort in the empirical analysis which shuts down resident incentives to free-ride. Importantly, we

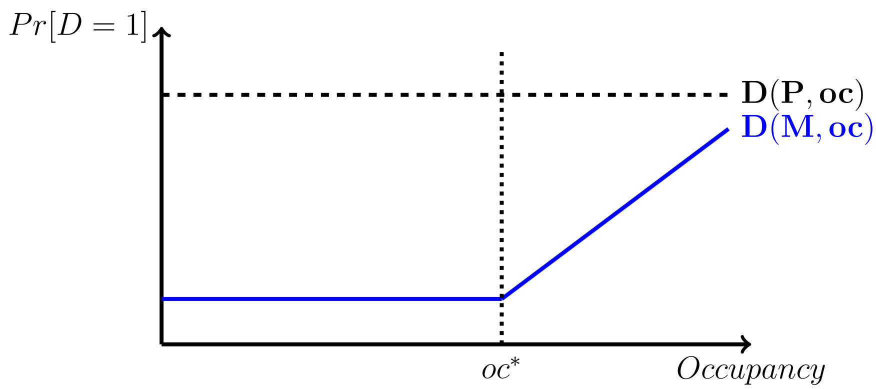



Figure 1: Predicted Discharge Profiles by Payer Type and Across Occupancies

note that free-riding, if present, would work against finding an effect of provider incentives on discharge rates.

**Graphical Discussion:** We summarize these theoretical predictions in Figure 1, which plots the per period discharge probability by payer type on the vertical axis against the nursing home's occupancy rate on the horizontal axis. The occupancy rate only affects the nursing home's financial incentive. As the nursing home will not exercise effort to discharge a private payer ($e^{SNF,*}(P, oc) = 0$ for all $oc$) their discharge rates are constant in occupancy as indicated by the horizontal dashed black line. This is not true for Medicaid beneficiaries. At low occupancy rates, nursing homes are not willing to exercise costly effort as the flow payoff exceeds the option value of drawing a private payer (net of the cost of effort) in the future. Intuitively, the refill probability $\Phi(oc)$ is too small, such that the marginal benefit of effort is strictly smaller than the marginal cost of effort. Hence, the nursing home chooses the corner solution of no effort, $e^{SNF,*}(M, oc) = 0$ for $oc < oc^*$, which explains the horizontal profile in the solid blue line for $oc < oc^*$.

Notice however that the discharge probability is smaller for Medicaid beneficiaries at low occupancy rates. This is because private payers exercise a greater discharge effort as they pay the nursing home rate in full: $e^{res,*}(P) > e^{res,*}(M)$. Hence, the difference in discharge probabilities at low occupancy rates is purely driven by patient incentives, as the nursing home's optimal effort is zero for either payer type at low occupancy rates.

The nursing home's optimal effort decision for Medicaid beneficiaries changes at $oc = oc^*$, where the marginal benefit of effort equals the marginal cost of effort at $e^{SNF} = 0$, providing an interior solution. The marginal benefit of effort continues to increase in the occupancy rate. Hence, the nursing home raises its optimal effort with increasing $oc$, so it equates the marginal benefit and the marginal cost of effort. The latter increases in effort due to the convex nature of the cost of effort. Hence, we have $e^{SNF,*}(M, oc) \geq 0$ and $\frac{\partial e^{SNF,*}(M,oc)}{\partial oc} > 0$ for $oc \geq oc^*$. Therefore, the discharge probability of Medicaid beneficiaries increases in the occupancy rate if $oc \geq oc^*$, as shown in Figure 1. We formally derive this relationship under simplifying assumptions in Appendix Section B.

# 4 Data Description

To investigate the effects of payer type and nursing home occupancy on discharges, we combine resident data from the MDS and Medicaid and Medicare claims data with nursing home characteristics from annual surveys. The MDS contains at least quarterly detailed assessments of SNF residents' health and LTC needs for the universe of residents in Medicaid or Medicare certified nursing homes, about 98% of all nursing homes. The MDS also provides us with exact dates of admission and discharge, as well as the discharge reason. We merge the MDS with Medicaid and Medicare claims data at the nursing home stay level. This allows us to identify the source of payment at each point in time and to specify the timing of transitions from out-of-pocket pay to Medicaid coverage. We merge resident-level data with the On-Line Survey, Certification, and Reporting system (OSCAR), which contains the number of licensed beds, allowing us to calculate weekly occupancy rates. In addition, nursing home surveys from California and Pennsylvania provide us with daily private and Medicaid rates. See Appendix Section C for more details on these data sources.

We use combined data for four states (California, New Jersey, Ohio, and Pennsylvania) for the years 2000 to 2005. Among all individuals who were admitted to a nursing home during this time period, we select those who initially paid for their stay out of pocket, yielding about 1.4 million nursing home stays. We drop Medicare beneficiaries, who require intensive rehabilitative care services and thereby differ from the sample population of interest. We drop residents who are covered by Medicaid from the beginning of their stay to construct a more homogenous sample population with respect to their financial means. Hence, the remaining residents either pay their entire stay out of pocket or become eligible for Medicaid during their stay. We drop about 17,000 resident-stays in which the resident transitions to Medicaid within the first week of the stay because these residents are similar to those who are covered by Medicaid at admission.

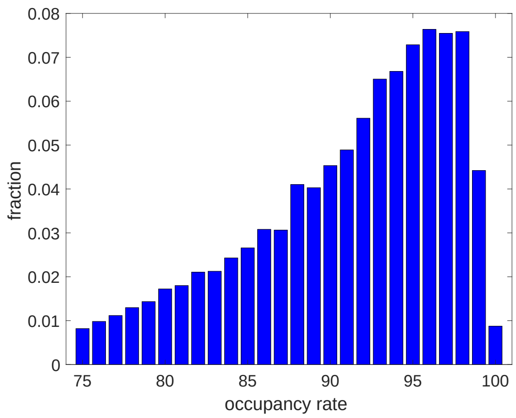
Using the admission and discharge dates from the MDS, we convert the data into a long format where each observation corresponds to a resident-week. We drop all resident-week observations for nursing homes whose occupancy never exceeds 60% or at least once exceeds 130% during the sample period since we are concerned about measurement error in the reported number of licensed beds. This restriction reduces the number of nursing home stays by another 115,000. To further reduce the role of measurement error in occupancy, we only consider nursing home stays for which the occupancy rate varies between 65 and 100 percent during the entire stay. We also restrict the analysis to nursing homes that accept

Medicaid payers. These refinements reduce the number of observations by 350,000. The final sample population consists of about 940,000 nursing home stays and 15.4 million week-stay observations.

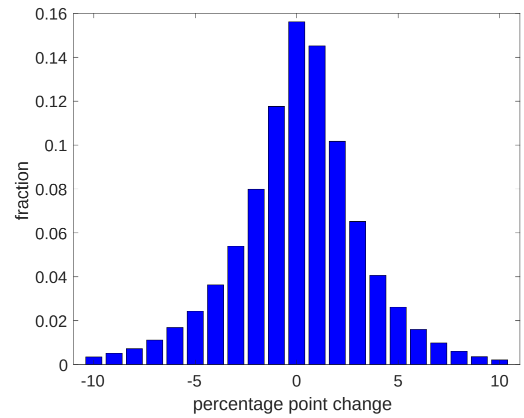
These data provide us with the necessary variation in provider and patient discharge incentives. First, we provide evidence for variation in occupancy rates, which drive nursing homes' incentives to discharge Medicaid residents. Figure 2a summarizes the overall variation in occupancy rates over time (weeks) and between nursing homes. The average occupancy rate equals 91% which translates into 11 empty beds in an average sized nursing home with 120 licensed beds. (See Figure 7 in Appendix Section C for a histogram of the number of beds.) There is considerable occupancy variation between nursing homes and, more importantly for our analysis, within nursing homes in a given year. Within nursing homes, conditional on nursing home and year fixed effects, we find a standard deviation in occupancy of 3.4 percentage points (about 63% of the standard deviation in nursing home fixed effects). Figure 2b displays this variation graphically.

An important driver of the intertemporal variation in occupancy is the volatility in the number of new admissions. Figure 2c shows the frequency of new admissions divided by total number of beds to translate admissions into changes in occupancy rates. It is evident from this tabulation that the relative number of arrivals can vary substantially from week to week leading to unexpected variation in occupancy. Whether this variation affects providers' discharge incentives in a meaningful way depends in part on the persistence of these occupancy shocks. To assess the persistence, Figure 2d displays the impulse response function of occupancy rates to a sudden 3 percentage point increase and decrease in occupancy relative to the sample average. Specifically, we construct an occupancy transition matrix from the data and simulate the occupancy rate profile over time. The response functions indicate that it takes 100 weeks or two years until the occupancy rate reaches its average again. However, it takes only about 25-30 weeks until half of a shock's effects have dissipated, which roughly coincides with the average length of stay of 25.7 weeks in our sample population (indicated by the vertical line in Figure 2d).

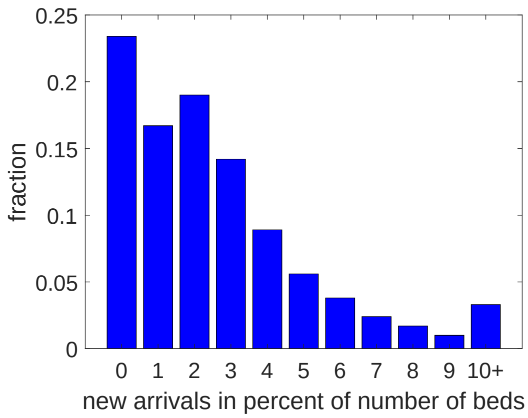
The measured persistence of occupancy shocks indicates that nursing homes are likely to take them into account when planning their discharge efforts. In particular, the variation may impact the nursing home's community discharge efforts, which is the most common discharge reason in our sample, when measured at the nursing home stay level. About 45% of the stays end with a community discharge, whereas only 20% and 15% of stays, respectively, last until a person passes away or is discharged to a hospital without an anticipated return, see Figure 6a in Appendix Section C for details. Another 11% are of stays end with a discharge to a different nursing home and remaining 9% of stays account for discharges to an assisted living



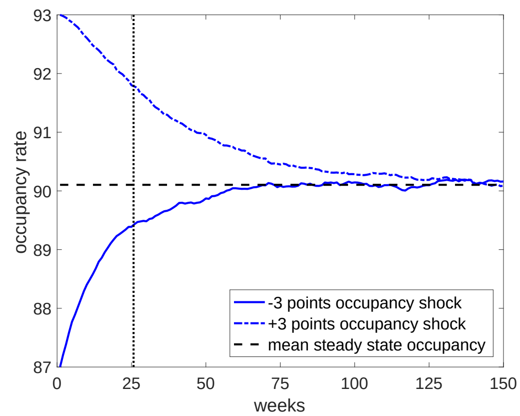

(a) Occupancy Rate Distribution




(b) Occupancy Variation Within SNF and Year




(c) New Arrivals Relative to Number of Beds




(d) Impulse Response to Change in Occupancy

***Notes:*** *Figure 2a presents occupancy rate variation. Figure 2b shows the residual variation conditional on nursing home year fixed effects. Figure 2c summarizes the frequency of weekly arrivals, divided by the number of licensed beds. The unit of observation for Figures 2a, 2b, and 2c is the nursing home week level. Figure 2d presents two impulse response functions, that document the mean reversion of an initial deviation of* $\pm 3$ *percentage points. The vertical line marks the average length of a nursing home stay.*

Figure 2: Variation in Occupancy Rates and New Arrivals by SNF and Week

facility and censoring.[12]

Second, we compare Medicaid beneficiaries with residents who pay out of pocket. Table 1 shows resident-week-level summary statistics for our estimation sample, split by payer type in the given week of the stay. Overall, the payer types are remarkably similar in terms of socio-demographics and their health profiles. However, Medicaid beneficiaries are slightly healthier, based on the CMI, more likely to be black and less likely to reside in a for-profit nursing home. Moreover, Medicaid beneficiaries are substantially more likely

[12] The relative importance of discharge reasons shifts heavily from home discharges to mortality and censoring, when evaluated at the week-of-stay level. The implicit weighting by length of stay reduces the fraction of home discharges to 11%, see Figure 6b in Appendix Section C.

| | Private | | Medicaid | |
|---|---|---|---|---|
| | Mean | SD | Mean | SD |
| Age | 79.05 | (13.83) | 79.19 | (12.42) |
| Female | 0.62 | (0.48) | 0.71 | (0.45 |
| White | 0.85 | (0.36) | 0.83 | (0.38) |
| Black | 0.07 | (0.25) | 0.11 | (0.31) |
| Married | 0.26 | (0.44) | 0.23 | (0.42) |
| Widowed | 0.47 | (0.50) | 0.53 | (0.50) |
| Separated/Divorced | 0.08 | (0.27) | 0.11 | (0.31) |
| Case Mix Index (CMI) | 1.03 | (0.50) | 1.00 | (0.44) |
| Number of ADL | 10.36 | (4.80) | 10.70 | (4.86) |
| Clinical Complexity | 0.41 | (0.49) | 0.35 | (0.48) |
| Depression | 0.34 | (0.47) | 0.43 | (0.49) |
| Weight Loss | 0.11 | (0.31) | 0.09 | (0.29) |
| Impaired Cognition | 0.42 | (0.49) | 0.46 | (0.50) |
| Behavioral Problems | 0.09 | (0.28) | 0.09 | (0.29) |
| Observations | 9,693,761 | | 5,711,288 | |

***Notes:*** *Data are from the MDS 2000 to 2005. The table presents summary statistics by payer source at the week of stay level. The resident's health status is decreasing in each health measure. The CMI is a summary measure of long term care needs, calculated based on methodology 5.01, and normalized to 1. The remaining health measures are direct inputs to the CMI formula and provide more granular information on cognitive and physical disabilities.*

Table 1: Resident-Week-Level Summary Statistics

to be female and widowed. These differences may point to potential differences in access to informal care givers if, for example, women provide informal long term care for their husbands. To investigate potential differences in access to community-based care further, we supplement Table 1 with data from from the National Long Term Care Survey (NLTCS) and the Health and Retirement Study (HRS). The HRS data indicate that access to informal care givers is similar between private payers and Medicaid beneficiaries. With respect to financial outcomes, the NLTCS data point to higher home ownership rates among private payers when compared to residents who transition into Medicaid but the difference is not statistically different from zero, see Appendix Section C. We view these differences between Medicaid beneficiaries and private payers as relatively minor.

# 5 Empirical Strategy

We exploit within-stay (within-spell) variation in patient's and provider's financial incentives and model their effect on the weekly discharge rate (hazard rate). To ease the computational burden of the large number of observations and fixed effects, we estimate a series of linear probability models. We can flexibly control for duration dependence via week-of-stay fixed effects and assume that the residuals are i.i.d. over time.

Specifically, we estimate the following reduced form regression model for equation (1), which expresses financial incentives as flexible functions of occupancy and payer type:

$$Y_{ijst} = \sum_{k=65}^{100} \gamma^k occ_{jt-1}^k + \sum_{k=65}^{100} \delta^k occ_{jt-1}^k Mcaid_{is} + \alpha_{jy} + \alpha_c + \alpha_s + X_i'\beta + \epsilon_{ijst}. \tag{2}$$

Here $Y_{ijst}$ denotes an indicator variable that equals one if resident $i$ in nursing home $j$ was discharged to the community in stay week $s$ and calendar week $t$. $occ_{jt}^k$ is an indicator variable that turns on if the (rounded) occupancy rate equals $k = 65, \ldots, 100$ percent in nursing home $j$ in calendar week $t$. $Mcaid_{is}$ is an indicator variable for resident $i$ being covered by Medicaid in week $s$ of her stay.[13]

The coefficients of interest are $\gamma^k$ and $\gamma^k + \delta^k$, which can be interpreted as the effect of occupancy on weekly discharge probabilities for private payers and Medicaid beneficiaries, respectively. We are particularly interested in the differences between payer types captured by $\delta^k$. These estimates are conditional on facility-year fixed effects $\alpha_{jy}$, calendar month fixed effects $\alpha_c$, week-of-stay fixed effects $\alpha_s$, and health status at admission $X_i$.

Occupancy may be endogenous to discharges for several reasons. First, there is a mechanical reverse relationship between the own discharge process and the occupancy rate. To address this issue, we use a leave-one-out measure for the occupancy rate. Specifically, we explore occupancy rate variation in *other* beds. To implement this strategy, we use the lagged occupancy rate, which only varies in other beds, once we exclude the first week of the stay. To see this, note that an individual resident only affects the facility-level occupancy rate in the weeks when she is admitted and discharged. By dropping the first week of each stay and using the lagged occupancy rate, we remove the variation in the last week of each nursing home stay that is partly due to the resident's own discharge.

Second, to control for potential drivers of discharges that may be correlated with the occupancy rate and the resident's payer type, we add week-of-stay fixed effects $\alpha_s$ in regression (2). These effects control for unobserved differences in health types that are important for the remaining length of stay. For example, the discharge probability in a given week is considerably lower for residents that have already spent six months in the nursing home compared to residents who were admitted within the previous two weeks. In other words, the week of stay fixed effects flexibly control for duration dependence in the discharge profiles. We also control for observable health measures at admission contained in $X_i$ to construct a

[13]Given the Medicaid eligibility rules that depend on asset spend-down (see Section 2), the Medicaid indicator switches from 0 to 1 at most once during a resident's stay.

homogenous comparison population.[14] In addition, we control for calendar month fixed effects, $\alpha_c$, to account for seasonal variation in discharges. Finally, we control for facility-year fixed effects captured by $\alpha_{jy}$ in regression (2). Any unobserved effects that may systematically impact discharge rates, such as nursing homes' management decisions, prices, and quality, are absorbed in this fixed effect. Hence, we only explore variation in occupancy within nursing homes in a given year. We assume that the residuals are i.i.d. over time and therefore uncorrelated with the right-hand side control measures, that are all history-based.

Given our rich set of nursing home-year fixed effects and week of stay fixed effects, we compare residents in the same nursing home and year and the same week of their stay. We therefore attribute remaining differences in discharge rates between payer types and across occupancies to financial incentives.

# 6 Empirical Discharge and Health Patterns

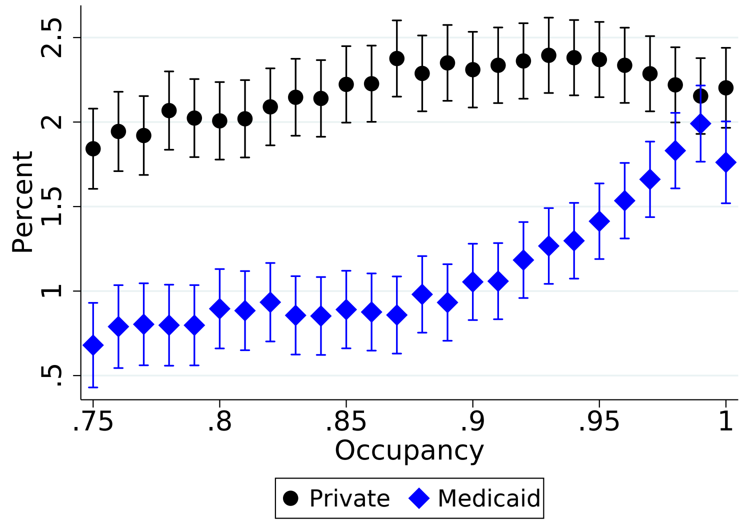
In this section, we provide reduced-form evidence for financial patient and provider incentives driving nursing home discharges, i.e. we test the implications of the theoretical model presented in Section 3.

## 6.1 Home Discharge Patterns by Occupancy and Payer Type

Figure 3 is the empirical analogue to Figure 1 and shows the estimated effects of occupancy and payer type on the home discharge rate along with the 90% confidence intervals. The depicted estimates correspond to the mean-adjusted coefficients $\hat{\gamma}^k$ for private payers and $\hat{\gamma}^k + \hat{\delta}^k$ for Medicaid beneficiaries, respectively. The horizontal axis shows the percentage of occupied beds between 75% and 100%. These estimates are conditional on facility-year, month, and week of stay fixed effects and resident characteristics including health at admission and therefore isolate the effect of occupancies on payer type specific discharge rates from potential selection due to length of stay or resident's health.

The figure provides two important insights. First, private payers have higher discharge rates across the entire range of occupancies. On average, private payers have a 2% chance of being discharged to the community in a given week. In contrast, the discharge rate among Medicaid residents below 90% occupancy equals only 0.8%. This observation is consistent with differences in resident cost-sharing, as discussed in Section 3. Since SNFs do not have a financial incentive to discharge residents of either payer type at low occupancies, the estimated difference in discharge rates suggests that resident incentives affect the length of stay.

[14]These health measures include the individual case mix index, as well as the predicted length of stay. We construct the latter by regressing length of stay on a rich set of disability and health measures and obtaining the predicted outcome for each resident.




***Notes:*** *This figure plots the estimated coefficients* $\hat{\gamma}^k$ *(private) and* $\hat{\gamma}^k + \hat{\delta}^k$ *(Medicaid) from regression (2) for the dependent variable "home discharge" across occupancy rates* $k$*. The coefficients are adjusted to match observed mean discharge rates by payer type. The vertical bars indicate 90% confidence intervals.*

Figure 3: Home Discharge Rates by Payer Type and Occupancy

Second, Medicaid home discharge rates increase above 89% occupancy, reaching almost 2% as occupancy approaches 100%.[15] Interpreted through the lenses of the theoretical model, the kink at around 89% occupancy marks the point at which the nursing home starts to exercise a positive discharge effort. At lower occupancies, nursing homes benefit from extended Medicaid stays, to the extent that Medicaid rates exceed the marginal cost of care. At higher occupancies, this incentive is muted because nursing homes prefer to occupy their scarce beds with more profitable private payers. In contrast, private payers' home discharge rates vary little (between 1.9 and 2.4%) and not systematically with occupancy. Consistent with the theoretical predictions, these findings suggest that provider incentives affect the length of stay as well.

In Appendix Section D.1, we show that discharges to a hospital, another nursing home or due to the resident's death vary with occupancy to a much smaller extent. In fact, as we discuss in greater detail below, changes along these discharge margins predominantly reflect changes in the health composition of residents as a result of endogenous changes in home discharges. We therefore conclude that financial patient and provider incentives operate mostly through the home discharge margin.

[15]With respect to our estimates at 100% occupancy, we note that these are likely contaminated by measurement error, which biases the estimates towards the average discharge probability across occupancy rates. This explains the modest reverse in the discharge pattern.

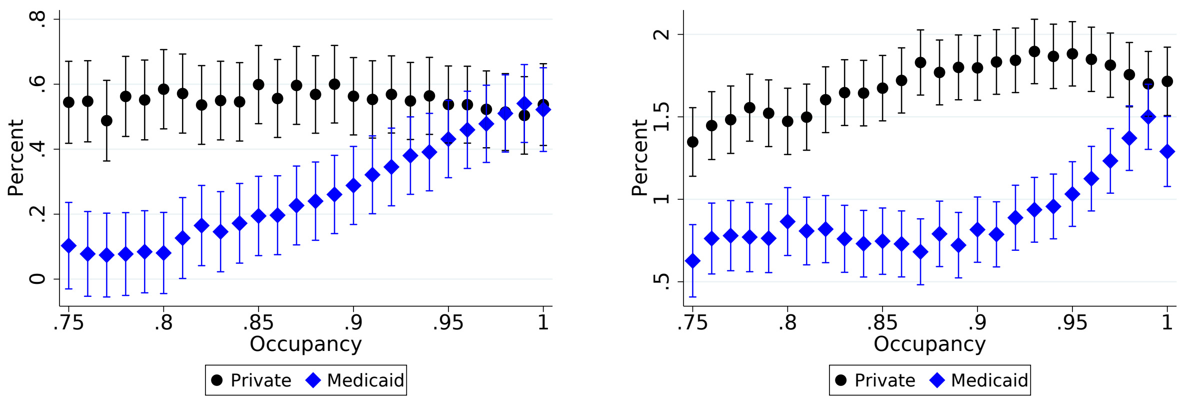
## 6.2 Who Are Residents at the Margin of Being Discharged?

To gain a better understanding of the marginal Medicaid residents who are discharged at high occupancies, we take advantage of more granular information on the discharge code. Specifically, the MDS divides home discharges into cases that require home health follow up care and cases that do not. Figure 4a documents the analogous effects of occupancy and payer type on community discharges that do not require home health care. Here, we see larger changes discharge rates among Medicaid beneficiaries than in Figure 3. Discharge rates increase about tenfold between 75% and 100% occupancy, from a weekly discharge rate of 0.06% to about 0.6%. In contrast, discharge rates among private payers do not vary with occupancy and equal about 0.6% throughout the occupancy range. In relative terms, this increase is much larger than the doubling in Medicaid discharge rates in Figure 3. Figure 4b presents the results for community discharges that require some home health care. Here the evidence is similar to the patterns for community discharges overall.

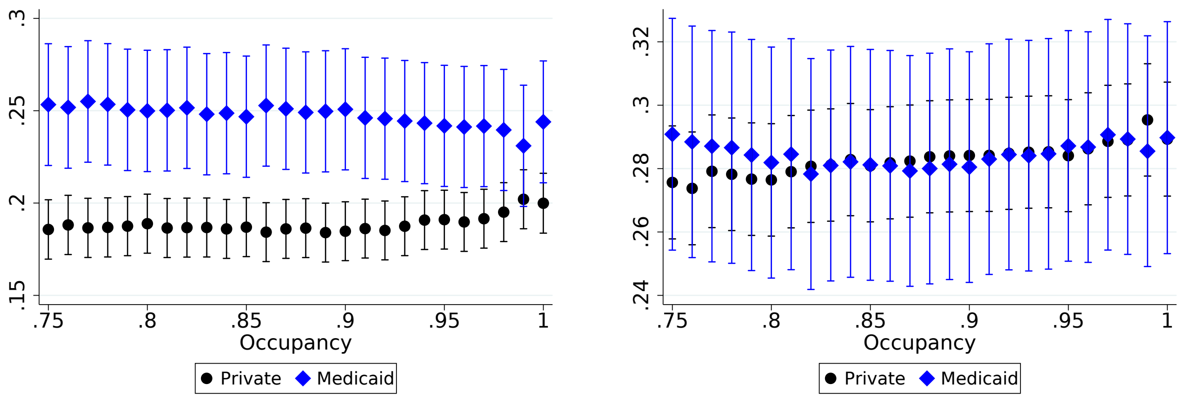
These results indicate that the residents at the margin of being discharged home are relatively healthy and often do not require home health care. Hence, longer nursing home stays among these Medicaid beneficiaries may contribute towards LTC overspending. Interestingly, the discharge rates for Medicaid beneficiaries who are discharged without requiring home health care exhibit a kink point at a lower occupancy rate of about 80%. Nursing homes therefore seem to weigh financial benefits and the LTC needs of their residents when deciding on their discharge effort.

We revisit the health composition of marginal residents in additional analyses, in which we restrict the sample population to residents who do not require SNF care on clinical grounds. Following Mor et al. (2007), we restrict the sample to seniors with very low care needs and high prospects to return to the community. Within this subpopulation, we find higher average discharge rates, larger level differences between Medicaid and private payers at low occupancy rates, and larger increases in discharge rates for Medicaid beneficiaries as the occupancy rate increases from about 87% to 100%, see the left series of graphs in Figure 9 of Appendix Section D.2. These findings corroborate the view that marginal residents appear to be relatively healthy and support the concern that longer nursing home stays may contribute to overspending. We revisit the latter point in Section 6.5.

The fact that marginal residents are relatively healthy also suggests that financial incentives lead (ex post) to an advantageous selection of Medicaid beneficiaries. If financial incentives lead to longer nursing home stays among healthy Medicaid beneficiaries, then conditional on health at admission, Medicaid beneficiaries in nursing homes are advantageously selected when compared to residents who pay out-of-pocket. This prediction is supported




(a) Home Discharges Without Home Health Care (b) Home Discharges With Home Health Care




(c) 1 Year Hospitalization Rate

(d) 1 Year Mortality Rate

***Notes:*** *See notes for Figure 3. The dependent variables are indicator variables that equal one if a resident was discharged to the community without requiring home health care (Figure 4a) or with home health care (Figure 4b) in a given week or if a resident was hospitalized (Figure 4c) or passed away (Figure 4d) within one year from a given week. The vertical bars indicate 90% confidence intervals.*

Figure 4: Home Discharge Rates and Health Outcomes by Payer Type and Occupancy

by lower weekly mortality rates for Medicaid beneficiaries, presented in the bottom right graph of Figure 8, see Appendix Section D.1. Interestingly, the difference in mortality decreases as the occupancy rate increases. Increasing financial provider incentives reduces the length of stay for healthy marginal Medicaid beneficiaries and therefore mutes the ex post advantageous selection. These hypotheses are further supported by the evidence on hospital discharge rates presented in the top right graph of Figure 8.

## 6.3 Patient Incentives and Discharges

Following the theoretical predictions, differences in discharge rates between payer types at low occupancies indicate that patient financial incentives affect the length of nursing home stays. In this section, we briefly summarize a series of robustness exercises that corroborate this interpretation. We are primarily concerned about two sources of bias and organize our discussion around theses threats to identification.

First, differences in health profiles between Medicaid beneficiaries and private payers may explain the different discharge rates. While we control for a rich set of health characteristics at admission, including the predicted length of stay based on observed health measures, unobserved differences in health profiles could explain differences in discharge rates. To alleviate this concern, we note that, if anything, Medicaid beneficiaries appear to be healthier than private residents, which would bias us against finding longer Medicaid stays. Specifically, we find that conditional on week of stay fixed effects, Medicaid residents appear healthier as evidenced by a lower CMI, see Figure 10 in Appendix Section D.3. Furthermore, as discussed in Section 6.2, financial incentives imply that Medicaid beneficiaries are (ex post) advantageously selected, again providing evidence against this concern.

Second, we are concerned that private payers have better access to home health care as they are wealthier than their peers who transitioned into Medicaid. As a first step towards constructing a sample population that is homogenous with respect to financial and medical characteristics, we focus on residents who pay out-of-pocket at the beginning of their stay. Data from the NLTCS suggest that differences in income and wealth become statistically insignificant once we make this refinement, see Appendix Section C for details. Furthermore, we note that differences in discharge rates are also visible among relatively healthy seniors who do not require home-based health care and for whom living in the community is therefore perhaps more affordable, see Section 6.2.

We undertake several additional attempts to further reduce differences in financial means between private payers and Medicaid beneficiaries. First, we use a propensity weighting approach, where we predict Medicaid transitions using the rich demographic information in the MDS, including the zip code of the former residence, the educational attainment, as well as gender, age, and race. This approach balances the covariates of residents who transition into Medicaid and those who keep paying out-of-pocket and delivers results that are similar to our main findings in Figure 3, see Appendix Section D.4. In the other robustness check, we focus on Medicaid applicants. Not all Medicaid applicants are granted Medicaid, providing us with payer type variation among a pool of residents who appear sufficiently poor to consider applying for benefits. Again we find very similar albeit noisier differences in discharge rates between payer types in this smaller sample population, see Appendix Section D.5.

In addition to refining the sample population, we also exploit details on the timing of the Medicaid transition to corroborate the financial patient incentive channel. Specifically, we expect that the reductions in out-of-pocket pay have a disproportionate effect on discharge rates in the short term. To test this prediction, we estimate differences based on the week since Medicaid transition. As expected, we find the largest reductions in the discharge probability in the first week a resident is covered by Medicaid. These differences fade out as residents continue their stay as Medicaid beneficiaries, see Appendix Section D.6.

To provide further evidence on patient moral hazard, we exploit differences in private rates between nursing homes in a difference-in-differences analysis. Our theoretical model predicts larger differences in discharge rates between private payers and Medicaid beneficiaries when the private rate is high. This test is also motivated by the fact that the variation in private rates does not affect the co-payments of Medicaid beneficiaries in medically needy programs. Therefore, we can corroborate our baseline findings against the concern that these small co-payments bias our baseline differences in discharge rates downward. To test this prediction, we divide the sample of nursing homes using the level of the private fee using price data from Pennsylvania and California. We find that the decline in discharge rates after payer type switches is more pronounced in SNFs with higher private fees, i.e. where patient financial incentives are highest. Furthermore, the magnitude of these differences is consistent with the overall difference between private payers and Medicaid beneficiaries in discharge rates observed in Figure 3, when scaling the price difference to the overall private rate, see Appendix Section D.7 for details. Finally, we exploit differences in Medicaid eligibility rules between New Jersey, Ohio, and Pennsylvania in a border analysis. Using state of residence as an instrument for Medicaid eligibility, we find further evidence for reduced discharge rates among Medicaid beneficiaries, see Appendix Section D.8.

Finally, we point out that both concerns, addressed here, suggest that our baseline findings might overstate the role of patient incentives, which is relatively inconsequential for two reasons. First, it works against an important result of this study, which is that providers respond more elastically to financial incentives than residents. Second, in the counterfactual experiments we compensate Medicaid beneficiaries via generous lump sum transfers before exposing them to co-pays. This also mitigates any differences in income and wealth between private and Medicaid residents.

## 6.4 Provider Financial Incentives and Discharges

We explain the observed increase in Medicaid discharge rates at high occupancy rates with an increasing option value of an empty bed. To corroborate this interpretation, we now provide direct evidence on the refill probability, which determines the option value of an empty bed in our framework. Specifically, we combine the observed number of vacant beds and realized admissions to measure the weekly probability that an empty bed is refilled.

To this end, we consider a nursing home with $k \geq 0$ newly arriving seniors per week. We assume that arriving seniors are randomly assigned to $v$ vacant beds. If $k > v$, demand exceeds capacity and the nursing home must turn away $k - v$ of the newly arriving seniors. The probability that a focal bed remains empty in a given week equals:

$$Pr[Not\ Refilled] = \begin{cases} \frac{v-1}{v} \times \frac{v-2}{v-1} \times \cdots \times \frac{v-k}{v-k+1} = \frac{v-k}{v} & \text{if } k < v \\ 0 & \text{otherwise }. \end{cases}$$

Hence, the probability the bed is refilled is simply:

$$\Phi = Pr[Refilled] = 1 - Pr[Not\ Refilled] = 1 - \max\left\{\frac{v-k}{v}, 0\right\}. \tag{3}$$

We note that censoring in admissions, induced by rationing, may bias the number of observed newly arriving seniors $k$ downward. We are, however, less concerned about a downward bias in the refill probability since $\Phi = 1$ for $k \geq v$. In other words, there is no downward bias in $\Phi$ as long as we observe $k = v$ whenever $k \geq v$.

We measure $\Phi$ at the nursing home week level, and construct its conditional mean by weekly occupancy. We find a highly convex relationship between the refill rate and the occupancy rate. The refill probability increases only from 10% to 18% between 75% and 90% occupancy. However, between 90% and 100% occupancy, the refill rate increases drastically from 18% to 60%, see Figure 14 in Appendix Section D.9 for details. We also note that 72% of newly-admitted (non-Medicare) residents pay out-of-pocket at the beginning of their stay. Combined with the high refill probability at high occupancy rates, this provides nursing homes with a strong incentive to discharge Medicaid beneficiaries at high occupancies as observed in Figure 3.

We revisit the effect of provider financial incentives on home discharges in an additional robustness exercise. Using price data from California and Pennsylvania, we divide nursing homes into two groups based on the difference between the daily private rate and the Medicaid reimbursement rate. We find that the increase in Medicaid discharge rates at high occupancies is more pronounced among nursing homes with a larger rate difference, which gain more from discharging Medicaid residents, see Appendix Section D.9 for details.

## 6.5 Health Benefits From Longer Nursing Home Stays

Next we turn to the health effects of longer nursing home stays using the variation in financial incentives as a source of exogenous variation in the remaining length of stay. Specifically, a higher Medicaid discharge rate at a high occupancy rate suggests that Medicaid residents face a shorter expected remaining length of stay when the current occupancy rate is high. Building on this observation, we use regression (2) to quantify differences in health outcomes between payer types and occupancies. For example, if longer nursing home stays lead to better health outcomes, then we expect the best health outcomes for Medicaid beneficiaries at low occupancy rates because of a longer expected remaining length of stay.

We start with an analysis of the one year mortality and hospitalization rate. Using the Medicare claims data, we construct two indicator variables that turn on if the resident deceased or was hospitalized within one year from the current week of stay, respectively. Importantly, we can construct the indicators independently of the realized residual length of stay in the nursing home. Figure 4c shows the mean-adjusted coefficient estimates for the one-year hospitalization rate. Despite a longer remaining average length of stay for Medicaid beneficiaries, we document an overall larger one-year hospitalization rate. If anything, this suggests that longer nursing home stays are harmful for the marginal residents. More importantly, the difference between payer types, $\hat{\delta}^k$, converges in the occupancy rates and we also see a declining hospitalization rate for Medicaid beneficiaries, $\hat{\gamma}^k + \hat{\delta}^k$. The changes in $\hat{\delta}^k$ and $\hat{\gamma}^k + \hat{\delta}^k$ are statistically significant at the 1% level and provide evidence against positive health effects from longer nursing home stays.[16] Figure 4d shows the the analogous mean-adjusted coefficient estimates for the one-year mortality rate. Here we find no evidence that longer nursing home stays lead to lower or higher mortality rates.

Mortality and hospitalizations are two outcomes that indicate severe health problems. In addition, seniors' well-being also depends on their overall health status and the amount of ADL limitations they suffer. We therefore provide additional results on health outcomes and disability measures, evaluated at the time of discharge, in Figure 16 in Appendix Section D.10. If longer SNF stays lead to better health outcomes, we should find that Medicaid residents are generally healthier than private payers but we should see a decline in Medicaid residents' health above 90% occupancy as discharge rates increase. We can reject both predictions. Medicaid beneficiaries have worse health profiles at discharge and we find no evidence for a systematic worsening of Medicaid health profiles at occupancy rates exceeding 90%.

Overall, we conclude that longer stays among Medicaid residents do not lead to improved health outcomes and therefore likely constitute over-utilization of nursing home care.

# 7 Structural Model of Discharges

In the previous section, we documented that both provider and patient incentives affect nursing home discharge rates. In order to give the reduced form parameters a structural

[16] The statistical tests compare the average effects at low occupancies, ranging from 80% to 89%, and high occupancies, ranging from 91% to 100%. Specifically, we test the null hypotheses $H_0 : \frac{1}{10}\sum_{k=80}^{89}\delta^k - \frac{1}{10}\sum_{k=91}^{100}\delta^k = 0$ and $H_0 : \frac{1}{10}\sum_{k=80}^{89}(\delta^k + \gamma^k) - \frac{1}{10}\sum_{k=91}^{100}(\delta^k + \gamma^k = 0$.

interpretation, to compare the significance of resident and provider incentives, and to evaluate counterfactual policy experiments, we now turn to a structural model of nursing home discharges.

## 7.1 The Empirical Model

**Discharge Probabilities:** Our empirical discharge model builds on the theoretical model outlined in equation (1). Specifically, we assume that exogenous discharge factors ($\epsilon$ in equation (1)), which are not affected by financial incentives, are uniformly distributed, allowing us to express the probability of a discharge as follows:

$$\Pr[D = 1 | e^{SNF}, e^{res}] = D^{exog} + \alpha \times e^{SNF}(\text{FinInc}^{SNF}(\tau, oc)) + \beta \times e^{res}(\text{FinInc}^{res}(\tau)) \quad . \quad (4)$$

Here, $D$ denotes any discharge, which includes endogenous community discharges (our focus) but also discharges to a hospital, a different nursing home, or death, captured in an exogenous discharge rate $D^{exog}$. We extend our model of community discharges to overall discharges in order to capture the profit motives of nursing homes, which do not depend on the discharge reason, more accurately. While consumer payoffs depend on the discharge destination, we show below that the marginal benefit of resident effort remains unchanged in this extended framework. We define a period to be one week.

**Resident's Effort Choice:** In order to derive the resident's optimal effort choice, and to predict how discharge efforts and probabilities change in counterfactual policy experiments, we need to specify the resident's financial benefit from discharge. To this end, we assume the following indirect conditional utility:

$$W(\tau, D, \eta) = \begin{cases} u - \kappa p^{\tau} + \eta^{SNF} & \text{if } D = 0 \\ \eta^{home} & \text{if } D = 1 \end{cases}, \quad (5)$$

where $u$ is the resident's gross utility from a week of nursing home care, and we normalize utility from home health care to zero. To simplify notation, we set the utility from discharge equal to the utility from home discharge, $\eta^{home}$. Utility differences between discharge reasons will not affect optimal effort choices as shown below.

We aim to match the discharge profiles for private payers and residents who transition to Medicaid and assume that they have the same relative utility for nursing home care, $u$. As outlined in Section 6.3, better access to community based care among private payers would imply $u^P < u^M$ and suggest that our baseline estimates might overstate the role of financial resident incentives. $\kappa$ is a price coefficient, which maps the per-period price paid by the resident, $p^\tau$, into utility. $\eta^{SNF}$ and $\eta^{home}$ are type I extreme value taste shocks that are observed by the resident before choosing the effort level and unobserved by the SNF.

Residents choose the optimal discharge effort given by:

$$e^{res,*} = \arg\max_{e^{res}\geq 0} \Big\{ \Pr[D=1|\cdot, e^{res}] \times W(\tau, D=1, \eta) + (1 - \Pr[D=1|\cdot, e^{res}]) \\ \times W(\tau, D=0, \eta) - \kappa \times c(e^{res}) \Big\}, \quad (6)$$

where $c(e)$ denotes the cost of effort, measured in dollars. To translate the cost of effort back into utility, we multiply the cost by the price coefficient $\kappa$. The discharge probability is conditional on $D^{exog}$ and the resident's beliefs about $e^{SNF}$, captured by "$\cdot$" in $\Pr[D = 1|\cdot, e^{res}]$. However, these factors do not affect the resident's optimal effort because of the uniform distribution of $\epsilon$, see equation (1), shutting down potential free-riding incentives, as shown by the first order condition:

$$e^{res,*}(\tau, \eta) = \begin{cases} mc_e^{-1}\Big(\frac{\beta}{\kappa} \times \big(W(\tau, D=1, \eta) - W(\tau, D=0, \eta)\big)\Big) & \text{if } W(\tau, D=1, \eta) \\ & \quad > W(\tau, D=0, \eta) \\ 0 & \text{otherwise} \end{cases}, \quad (7)$$

where $mc_e^{-1}(\cdot)$ is the inverse marginal cost of effort function.

**Provider's Effort Choice:** Turning to the LTC provider, we assume that the SNF does not observe the resident's taste shocks nor the resident's discharge effort. Instead, the nursing home observes the payer type of the focal resident and forms expectations over her optimal effort level. Therefore, we assume that nursing homes maximize efforts under the following belief:

$$\Pr[D = 1|e^{SNF}, \tau] = D^{exog} + \alpha \times e^{SNF} + \beta \times E_\eta[e^{res,*}|\tau] \ . \quad (8)$$

To derive the optimal provider effort, $e^{SNF,*}$, we impose the following timing of events. During the period (week), providers choose their discharge effort level $e^{SNF}$ and realize the weekly flow payoff:

$$\pi(\tau) = \begin{cases} 0 & \text{if bed is empty: } \tau = 0 \\ r^\tau - mc & \text{otherwise} \end{cases},$$

where $r^\tau$ is the private fee or Medicaid reimbursement rate and $mc$ is the weekly marginal cost of providing LTC. Discharges, arrivals, and payer type transitions are random events which realize at the *end* of the period. Arrivals and payer type transitions are exogenous and characterized by the weekly refill probability $\Phi$, see equation (3) and the transition probability $\psi$, respectively. Discharges, in contrast, depend on endogenous discharge efforts as outlined in equation (8) and together with arrivals determine the occupancy rate in other beds $oc$.

To simplify the analysis, we assume that discharge managers do not coordinate their discharge efforts between residents (beds). In other words, discharge managers do not internalize the effect of their "focal" discharge decision on the occupancy rate and discharges in other beds, which are both endogenous equilibrium objects. Instead, we assume that, in equilibrium, the discharge manager takes the time series process of the occupancy rate in other beds as given and chooses the discharge effort in the focal bed optimally. To reduce the state space, we model occupancy rate transitions as a Markov process, which is characterized by a week-to-week transition matrix, $\Theta$. This transition matrix denotes the conditional probability mass function over next week's occupancy rate in other beds, $oc'$, conditional on today's occupancy rate on other beds, $oc$: $\Theta(oc, oc') = \Pr[oc'|oc]$.[17]

Combining the timing assumptions with the assumptions on arrivals, discharges, and payer type transitions, we can express the SNF's optimal discharge efforts through the following Bellman equation:

$$V(\tau, oc) = \max_{e^{SNF} \geq 0} \left\{ \pi(\tau) - c(e^{SNF}) + \delta E\Big[ V(\tau', oc') \Big| \tau, oc, e^{SNF} \Big] \right\} \quad , \tag{9}$$

where $\delta$ is a discount factor, $c(e^{SNF})$ is the cost of effort measured in dollars, and

$$\begin{aligned}
E\Big[V\Big|0, oc, e^{SNF}\Big] &= \sum_{oc'} \Theta(oc, oc') \times \Big[(1 - \Phi(oc')) \times V(0, oc') \\
&\quad + \Phi(oc') \times \Big(\rho V(P, oc') + (1-\rho) V(M, oc')\Big)\Big] \qquad (10) \\
E\Big[V\Big|M, oc, e^{SNF}\Big] &= \sum_{oc'} \Theta(oc, oc') \times \Big[\Big(1 - \Pr[D = 1|e^{SNF}, M]\Big) \times V(M, oc') \\
&\quad + \Pr[D = 1|e^{SNF}, M] \times \Big((1 - \Phi(oc')) \times V(0, oc') \\
&\quad + \Phi(oc') \times \Big(\rho V(P, oc') + (1-\rho) V(M, oc')\Big)\Big)\Big] \qquad (11) \\
E\Big[V\Big|P, oc, e^{SNF}\Big] &= \sum_{oc'} \Theta(oc, oc') \times \Big[\Big(1 - \Pr[D = 1|e^{SNF}, P]\Big) \times \Big((1-\psi) V(P, oc') \\
&\quad + \psi V(M, oc')\Big) + \Pr[D = 1|e^{SNF}, P] \times \Big((1 - \Phi(oc')) \times V(0, oc') \\
&\quad + \Phi(oc') \times \Big(\rho V(P, oc') + (1-\rho) V(M, oc')\Big)\Big)\Big]. \qquad (12)
\end{aligned}$$

[17] Conceptually, $oc$ refers to the occupancy in other beds. In practice, we approximate $oc$ by the overall occupancy rate in the structural estimation. We also assume that discharge managers form beliefs over next week's occupancy rate based on the current occupancy rate only. That simplification implies that the discharge manager does not condition on the payer type distribution nor the payer type in the focal bed.

In words, the value function combines the flow profit net of the cost of effort and a continuation value, see equation (9). The continuation value of an empty bed, as indicated in equation (10), is given by the probability of drawing a new resident, captured by the refill probability vector $\Phi(oc')$, multiplied by the payer type probability at admission. For example, the new resident is a private payer with probability $\rho$ delivering a payoff vector of $V(P, oc')$. Furthermore, expectations are taken over next week's occupancy rate as indicated by the transition matrix $\Theta(oc, oc')$. The continuation value of a bed filled with a Medicaid beneficiary, see equation (11), adds to this the possibility that the focal person may be discharged, which depends on the efforts of the nursing home and the resident. Finally, the continuation value of a bed filled with a private payer, see equation (12), adds to this a payer type transition to Medicaid, which happens with probability $\psi$.

## 7.2 Identification and Estimation

**Parameters estimated outside of the model:** We estimate several objects of the empirical analysis outside of the model, which are summarized in panel A of Table 2. We estimate the weekly refill probabilities according to equation (3). The results are summarized in Figure 14. The week-to-week occupancy transition matrix is an endogenous equilibrium object. For the purpose of estimation, we use the empirical transition matrix. An excerpt is provided in Table 4, see Appendix Section C. In the counterfactual exercises, we endogenize the transition matrix to allow for changes in discharge efforts affecting occupancy transitions, which in turn feed back into optimal effort choices.

We also estimate the probability of a payer type transition from private to Medicaid from observed week-to-week changes, which happens in 1.1% of cases. Excluding Medicare beneficiaries, about 72.3% of new residents initially pay out-of-pocket. Therefore, we set the probability that a new resident is a private payer to $\rho = 0.723$. To construct $D^{exog}$, we sum the average discharge rate to other destinations, e.g. hospitals, nursing homes, or mortality, see Figure 8. We simply take the unweighted average over occupancy and payer type specific discharge rates and find an average of 2.3%. This calculation abstracts away from observed differences between payer types, which we attribute to ex post advantageous selection among Medicaid beneficiaries, see again Section 6.2 for a discussion.[18] Finally, the private rate and the Medicaid reimbursement rate correspond to the average rates in Pennsylvania in the sample period.

[18]We abstract away from this as we do not model differences in health in the structural analysis. Importantly, these differences are considerably smaller than the differences in home discharges, which are captured in the model.

**Calibrated parameters:** We set the weekly discount factor to $\delta = 0.95^{1/52}$, as indicated in panel B of Table 2. We require a scale normalization on either the cost of effort or the return on effort since, naturally, they cannot be separately identified from our data. We decide to normalize the cost of effort function to $c(e) = e^2$ and thereby load differences in cost functions between payer types onto differences in the returns of effort.[19] Finally, we also normalize the utility from nursing home care $u$ as we can only identify utility up to scale. This is because utility affects discharges through effort in our specification, which is again scaled by the factor $\beta$.

**Parameters estimated within the model:** Finally, we turn to key structural parameters for our analysis that we estimate using the model. These include the occupancy threshold $oc^*$, the marginal cost of nursing home care per resident and day $mc$, the price coefficient $\kappa$ and the effort parameters $\alpha$ and $\beta$. Identification comes from the estimated discharge profiles in Figure 3. We next discuss the estimation strategy, which provides an identification argument by construction.

Our estimation routine first recovers the occupancy threshold $oc^*$. To this end, we construct a flexible algorithm that searches for a kink point in the Medicaid discharge profile, as suggested by the theory, see Figure 1. Specifically, we estimate linear regression models of the following form for any possible kink point occupancy rate $oc^k$:

$$D_{oc}^{Home,M} = \tau_1^k + \tau_2^k \times oc + \tau_3^k \times \mathbf{1}\{oc \geq oc^k\} + \tau_4^k \times oc \times \mathbf{1}\{oc \geq oc^k\} + \epsilon_{oc},$$

where $D_{oc}^{Home,M}$ denotes the estimated home discharge rate for Medicaid beneficiaries denoted by the blue diamonds in Figure 3. $\mathbf{1}\{oc \geq oc^k\}$ is an indicator that turns on if the occupancy weakly exceeds the hypothetical kink point $oc^k$. We are primarily interested in the slope change around $oc^k$, captured by parameter $\tau_4^k$, and choose the kink point with the most pronounced slope change in a statistical sense. Specifically, we estimate $\tau_4^k$ for any possible kink point and then choose the kink point that delivers the highest value from a comparison of $t$-statistics: $k^* = \arg\max_k \left\{|\hat{\tau}_4^k|/\widehat{SE}_{\hat{\tau}_4^k}\right\}$, where $\widehat{SE}_{\hat{\tau}_4^k}$ is the estimated standard error.

Next, we leverage the estimated kink point to recover the marginal cost of care per resident and day. At $oc^*$, the marginal benefit of effort to discharge a Medicaid beneficiary equals the marginal cost of effort when evaluated at $e^{SNF} = 0 = mc_e(0)$. Hence, the marginal benefit must be zero as well, which trades off the Medicaid flow profit $\pi(M)$ against the option

[19] To see this, consider the resident's optimal effort choice in equation (7). We observe the private rate $p^P$ as well as the discharge rates. Suppose we generalize the cost $c(e) = \gamma e^2$. Then we have $\frac{\partial Pr[D=1|\cdot]}{\partial p^P} = \beta \times \frac{\partial e^{res,*}}{\partial p^P} = \frac{\beta^{2*}}{2\gamma}$. Hence, scaling up the cost parameter $\gamma$ would simply scale up $\beta$. We note however, that the curvature in the Medicaid discharge profile for $oc > oc^*$, which we haven not exploited yet, could potentially be used to identify the curvature of the cost function.

value of drawing a new resident tomorrow. The option value increases in the refill probability. Intuitively, we can pin down the marginal cost that equates $\pi(M)$ with the option value when evaluated at $oc^*$. In practice, we solve the Bellman equation via contraction mapping for a guess of marginal costs. We then compare the occupancy threshold with the empirical counterpart and adjust the marginal cost guess accordingly. We repeat this procedure until predicted and observed occupancy thresholds coincide.

Finally, we recover the remaining parameters from the observed home discharge rates for Medicaid beneficiaries and private payers at low and high occupancy rates. To identify the resident coefficients, $\beta$ and $\kappa$, we focus on discharge rates at low occupancy rates, $oc < oc^*$, where providers do not exert effort. Specifically, we can pin down the two parameters by matching the observed and the predicted discharge rates by payer type as shown in equations (13) and (14) below. Here, $D_k^{Home,M}$ and $D_k^{Home,P}$ denote the measured community discharge rate for Medicaid beneficiaries and private payers at occupancy rate $k$ from Figure 3, respectively. Finally, conditional on $\beta$ and $\kappa$, the increase in discharge rates for Medicaid beneficiaries at higher occupancy rates, $oc \geq oc^*$, is only driven by provider incentives, which identifies $\alpha$. Specifically, we recover these parameters by matching the following three moments:

$$\frac{1}{N^D}\sum_{k=75\%}^{oc^*} D_k^{Home,M} = \frac{1}{N^D}\sum_{k=75\%}^{oc^*} \beta \times e^{res,*}(M) \tag{13}$$

$$\frac{1}{N^D}\sum_{k=75\%}^{oc^*} D_k^{Home,P} = \frac{1}{N^D}\sum_{k=75\%}^{oc^*} \beta \times e^{res,*}(P) \tag{14}$$

$$\frac{1}{N^{DS}}\sum_{k=oc^*}^{99\%} D_k^{Home,M} = \frac{1}{N^{DS}}\sum_{k=oc^*}^{99\%} \Big[\alpha \times e^{SNF,*}(M,k) + \beta \times e^{res,*}(M)\Big] \ . \tag{15}$$

While not explicit in equations (13) to (15), $\kappa$ enters the moment conditions through $e^{res,*}(\tau)$, see equation (7). To construct the means, we divide by the number of observations $N^D = 100 \times (oc^* - 75\%)$ and $N^{DS} = 100 \times (99\% - oc^*)$. We omit the estimated discharge rate at 100% occupancy in equation (15), to mitigate the effect of measurement error.

The moment equations (13) to (15) just identify the remaining parameters $\alpha, \beta$ and $\kappa$, whereby we do not require a weighting matrix.

**Inference:** We conduct inference via bootstrap. One computational limitation of this procedure is that estimating the regression model (2) is very time and memory consuming due to the large number of fixed effects and about 15 million observations. This makes it challenging to estimate the model many times as required in a standard bootstrap procedure. Instead, we leverage the observation that the OLS estimator for the vector $\tau = [\gamma^{75}, ..., \gamma^{100}, \delta^{75}, ..., \delta^{100}]$

is jointly normally distributed, $\hat{\tau} \sim N\Big(\tau, \Sigma\Big)$, because of the central limit theorem. Therefore, we only estimate the variance-covariance matrix for entire vector, $\Sigma$, once and then draw discharge coefficients. For each bootstrap iteration $b = 1, ..., B$, we draw $\hat{\tau}^b \sim N\Big(\hat{\tau}, \hat{\Sigma}\Big)$ and then re-estimate the parameters $oc^*, mc, \alpha, \beta$ and $\kappa$. Finally, we obtain 95% confidence intervals by ordering bootstrapped parameters, which are re-centered around the respective point estimates, and reporting the 2.5th and the 97.5th percentile.

# 8 Model Results and Policy Counterfactuals

In this section, we first discuss the model fit and model parameter estimates, and then use these results to conduct policy counterfactual simulations.

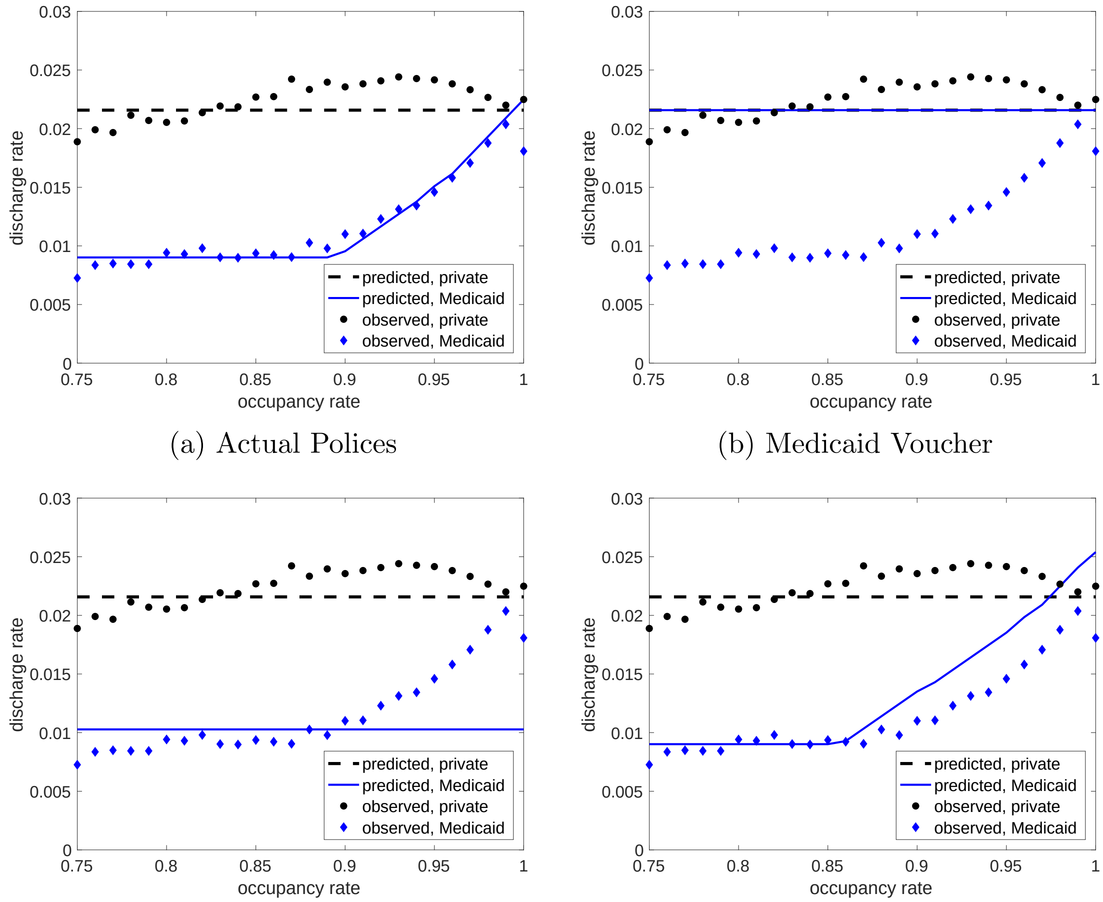
## 8.1 Estimates of Model Parameters and Model Fit

The model provides a very good fit to the observed community discharge rates, as shown in Figure 5a. The circles and diamonds denote the estimated discharge rates from Figure 3, and the lines denote the analogous model fit.

We estimate a kink point for the Medicaid discharge profile at $oc^* = 89\%$, as shown in Panel C of Table 2. The kink point allows us to pin down the marginal costs per resident and day, which we estimate at \$107. This estimate is smaller than marginal cost estimates from the literature. For example, Hackmann (2017) estimates an average marginal cost of \$159 using data from Pennsylvania over the entire resident population. Since we focus on a healthier marginal population of nursing home residents that are discharged back to the community, our smaller marginal cost estimate is highly plausible.

To provide intuition for the estimated location of the kink point and the magnitude of marginal costs, we consider a simple back of the envelope calculation. At the kink point, the weekly refill probability, $\Phi(oc^*)$, equals about 18%, see Figure 14, and the chance of drawing a private payer is $\Phi(oc^*) \times \rho = 18\% \times 72.3\% = 13\%$. If a private payer replaces a Medicaid beneficiary, then the weekly flow payoff increases by $7 \times (\$218 - \$188) = \$210$. On the other hand, the bed may remain empty for an extra week, which occurs with probability $1 - \Phi(oc^*) = 82\%$. In this case, the nursing home would forgo the Medicaid flow payoff of $7 \times (r^M - mc) = 7 \times (\$188 - \$107) = \$567$. To make the SNF indifferent, a new private payer would need to stay for about $\frac{82\% \times \$567}{13\% \times \$210} = 17$ weeks, which is roughly in the ballpark of the simulated length of stay of 22 weeks, see the first column of Table 3.

Both estimated discharge effort parameters are positive, $\hat{\alpha} > 0$ and $\hat{\beta} > 0$, implying that provider and resident discharge efforts increase the discharge probability. A naive comparison of these "marginal" coefficients would suggest that residents respond more elastically to financial incentives as $\hat{\beta} = 0.38 > \hat{\alpha} = 0.053$. However, these parameter estimates ignore




(a) Actual Polices (b) Medicaid Voucher

(c) Top-Up Payment (d) Front-Loaded Medicaid Reimbursement

Figure 5: Simulated Discharge Rates Under Different Policies

the fact that decisions on discharge efforts are often infra-marginal, in which case the effort decision does not respond to marginal incentives. We correct for this in a simple thought experiment.

We define FinInc as the net benefit from discharge, denominated in dollars, relative to staying an extra week. For residents, we have

$$\text{FinInc}^{res}(\tau) = \frac{1}{\kappa} \times \left(W(\tau, D = 1, \eta) - W(\tau, D = 0, \eta)\right) .$$

Equating marginal benefits and marginal costs, we have $e^{res,*} = \max\left\{\frac{\beta}{2} \times \text{FinInc}^{res}(\tau), 0\right\}$, see equation (7). Taking expectations over $\eta$, we find the following average marginal effect of incentives on discharges:

$$\frac{\partial D}{\partial \text{FinInc}^{res}(\tau)} = \frac{\beta^2}{2} \times Pr[W(\tau, D = 1, \eta) - W(\tau, D = 0, \eta) > 0].$$

| | | Mean | SE | 95% CI |
|---|---|---|---|---|
| *A. Estimated Outside of Model* | | | | |
| Refill Probability | $\Phi(oc)$ | See Figure 14. | | |
| Occupancy Transition Matrix | $\Theta(oc, oc')$ | See Table 4. | | |
| Pr[Payer Type Transition to Medicaid] | $\psi$ | 1.1% | | |
| Pr[Private Payer at Admission] | $\rho$ | 72.3% | | |
| Exogenous Discharge Rate | $D^{exog}$ | 2.3% | | |
| Daily Private Rate | $p^P$ | \$218 | | |
| Daily Medicaid Rate | $p^M$ | \$188 | | |
| *B. Calibrated* | | | | |
| Discount Factor | $\delta$ | $0.95^{\frac{1}{52}}$ | | |
| Cost of Effort Function | $c(e)$ | $e^2$ | | |
| Utility of Nursing Home Care | $u$ | 5 | | |
| *C. Estimated Inside of Model* | | | | |
| Occupancy Threshold for Medicaid | $oc^*$ | 89% | 0.20% | [89%,90%] |
| SNF Effort Parameter | $\alpha$ | 0.053 | 0.002 | [0.049,0.056] |
| Resident Effort Parameter | $\beta$ | 0.38 | 0.04 | 0.27,0.43] |
| Resident Price Coefficient | $\kappa$ | 0.064 | 0.03 | [0.053,0.167] |
| Marginal Cost of Care per Person and Day | $mc$ | \$107 | \$2 | [\$103,\$112] |

***Notes:*** *The table is organized into three panels. The first panel summarizes the objects that are estimated outside of the model. The second panel presents parameters that are calibrated. Finally, the third panel presents the parameters that are estimated inside the model. Inference is conducted via a bootstrap procedure.*

Table 2: Structural Parameter Estimates

The second factor, $Pr[\cdot]$, captures the observation that not all residents adjust their discharge efforts in response to marginal changes in incentives. For Medicaid beneficiaries, we have $Pr[\cdot] = \frac{1}{\exp(u)} = \frac{1}{\exp(5)} = 0.67\%$, so the effort decision is infra-marginal in more than 99% of cases. This implies $\frac{\partial D}{\partial \text{FinInc}^{res}(M)} = \frac{\beta^2}{2} \times 0.67\% = 0.048\%$.

For providers, we have the following effect of incentives on discharges

$$\frac{\partial D}{\partial \text{FinInc}^{SNF}} = \frac{\alpha^2 \delta}{2} \times Pr[\text{FinInc}^{SNF}(\tau) > 0].$$

Effort decisions are infra-marginal when occupancy is below the kink point $oc^*$. In practice, this happens with probability $Pr[oc \leq oc^*] = 36\%$. Hence, we find that $\frac{\partial D}{\partial \text{FinInc}^{SNF}} = \frac{\alpha^2 \delta}{2} \times (1 - 0.36) = 0.09\%$. Hence, Medicaid discharge rates respond almost twice as elastically to provider incentives than to resident incentives. We revisit this observation in the analysis of several policy counterfactuals.

| | Actual | Voucher | Top-up | 1% Front |
|---|---|---|---|---|
| Private LOS | 22.63 | 22.63 | 22.63 | 22.63 |
| Medicaid LOS | 28.06 | 22.61 | 30.83 | 26.71 |
| Average Occupancy | 91.1% | 88.1% | 92.4% | 90.3% |
| Δ Medicaid LOS | | −5.45 | 2.8 | −1.4 |
| Δ spending per stay in $ | | −198 | 9,055 | −500 |
| Δ annual spending in bn$ | | −0.03 | 1.6 | −0.09 |

***Notes:*** *The table summarizes the length of stay (LOS) in weeks, average occupancy rates, Medicaid savings per stay, and national Medicaid savings for the three counterfactual policy experiments.*

Table 3: Simulated Length of Stay and Cost Savings Under Policy Counterfactuals

## 8.2 Cost Savings and Policy Counterfactuals

### 8.2.1 Potential Cost Savings

Before turning to the policy counterfactuals, we consider the scope for cost savings in this context. To this end, we start by comparing Medicaid spending on nursing home care to expenditures that would arise when discharged nursing home residents lived in the community instead. These expenditures consist of home health care and spending on other medical care but also the cost of living expenditures including housing, food, etc. We note that Medicaid will cover only a fraction of these community expenditures. Therefore, our cost savings can be interpreted as overall LTC savings and as a lower bound on Medicaid savings. Using data from the Medical Expenditure Panel Survey and the Consumer Expenditure Survey, we find mean annual expenditures of $1,018 for formal home health care and of $12,903 for all other expenditures in the age group 80 and over.

We also add the opportunity cost of informal care provided by family members. Using data from the HRS, we find that the average individual receives 39 hours of informal care per month while the average conditional on receiving any informal care equals 92 hours. These values are roughly similar to previous findings in the literature. To calculate the opportunity cost of informal care, we use median weekly earnings from the Bureau of Labor Statistics for men and women for the years 2000 to 2005. Using the fact that 71% of informal caregivers of HRS respondents are women and assuming a 40-hour work week, we calculate an average hourly wage of $13.72. This implies annual opportunity costs of informal care giving of $\$14 \times 39 \times 12 = \$6{,}552$. Taken together, the expenditures and opportunity costs of home health care equal $\$1{,}018 + \$12{,}903 + \$6{,}552 = \$20{,}473$ per year or $56 per person and day. We note that our estimates on formal and informal community based care may overstate the expenditures for our sample popualation since many marginal nursing home residents do not receive home health care (see the results in Figure 4.)

Nevertheless, this estimate is considerably smaller than the average daily Medicaid reimbursement rate of \$188 and only about half as large as the estimated marginal cost of nursing home care of \$107. Medicaid spending could therefore be lowered by at least $(\$188 - \$56) \times 7$ days $= \$924$ per resident if the beneficiary was discharged one week earlier. The first column of Table 3 indicates that Medicaid stays exceed the length of stay of private payers by 5.4 weeks on average. Since we find no evidence that longer stays lead to health improvements, this difference points to Medicaid overspending of $5.4 \times \$924 = \$4{,}990$ per Medicaid stay or about 13.5% of Medicaid spending per nursing home stay.[20] Multiplying this fraction with national Medicaid nursing home spending of \$55bn in 2015 suggests annual savings of up to \$7.4bn.

To estimate the national savings more conservatively, we calculate the overall savings in our sample population and scale the estimates by a factor of 5, since the four states in our sample account for about 20% of overall Medicaid nursing home spending.[21] Combining a 13.5% reduction in Medicaid spending with a base of 1m Medicaid weeks per year in our sample, we calculate annual savings of \$0.2bn and \$0.9bn in our sample states and at the national level, respectively.

#### 8.2.2 Policy Counterfactuals

We consider three policy counterfactuals that change the discharge incentives for SNFs and nursing home residents. When simulating their effects on the length of stay and Medicaid spending, we take endogenous changes in occupancy rates into account, which in turn affect provider discharge efforts. To this end, we divide the nursing home into two wings. The additional wing allows us to incorporate admissions and discharges among residents that were excluded from the estimation sample but also affect overall occupancy. We treat these admissions and discharges as exogenous. For the study population (nursing home wing) of interest, we take observed weekly admissions as exogenous, and use our structural model to predict discharge rates under alternative policy regimes. Combining admission and discharge profiles between wings allows us incorporate the effect of policy changes on occupancy rates. In the simulation analysis, we then add an outer loop to the optimization problem, which searches for a fixed point in the discharge profiles, see Appendix Section E for details.

**Voucher Program:** First, we consider the introduction of a voucher program, which equates the marginal prices for nursing home care between Medicaid beneficiaries and pri-

[20] Overall Medicaid spending equals 28 weeks $\times$ 7 days $\times$ \$188 = \$36,848.

[21] See http://files.kff.org/attachment/Infographic-Medicaids-Role-in-Nursing-Home-Care.

vate payers, $p^M = p^P$. We compensate Medicaid beneficiaries for their expected outlays through a lump-sum transfer, which equals the average length of stay of private payers, 22.6 weeks, times the weekly private rate. This amounts to a lump-sum of \$34,533 per Medicaid stay. The program affects resident and provider incentives in opposing directions. On one hand, Medicaid beneficiaries now have a larger incentive to end their stays early. On the other hand, SNFs are now indifferent between private and Medicaid residents because they generate identical profits. Therefore, nursing homes will reduce their discharge effort for Medicaid beneficiaries to zero.

Figure 5b shows that Medicaid beneficiaries and private payers are now discharged at the same rate (the discharge rates lie on top of each other). As indicated in the second column in Table 3, the length of stay for Medicaid beneficiaries is reduced to 22.6 weeks, which also reduces the average occupancy rate to 88.1%. Medicaid saves 28 weeks of nursing home payments worth \$36,848 but provides transfers worth \$34,533 under this policy. We add costs for the additional 5.4 weeks spent in the community, worth 5.4 weeks $\times$ 7 days $\times$ \$56 = \$2,117 per Medicaid stay. Hence, overall savings equal \$36,848 − \$34,533 − \$2,117 = \$198 per Medicaid stay, corresponding to a 0.5% decline in Medicaid spending. Using the conservative approach we find that annual Medicaid savings of at least 0.5% $\times$ 1m weeks $\times$ 7 days $\times$ \$188 $\times$ 5 = \$0.03bn.

**Top-Up Program:** Next, we consider a top-up policy, where Medicaid residents pay the difference between the private fee and the Medicaid reimbursement rate, $p^P - p^M$, out-of-pocket. This equates the revenues among payer types and affects provider and patient incentives in opposing directions, as discussed in the voucher counterfactual.

Figure 5c shows that home discharge rates among Medicaid residents at low occupancies are slightly higher than under current policies (compared to Figure 5a) due to residents' increased discharge incentives. In contrast to current policies, we do not observe an increase in Medicaid discharges at high occupancies under top-up payments since both payer types are equally profitable for nursing homes. The incentive effects for providers outweigh the incentive effects for patients, leading to an overall increase in the length of stay by 2.8 weeks and a 1.3 percentage point rise in the average occupancy rate, see the third column in Table 3. The effect on spending is twofold. First, Medicaid has to cover 2.8 extra weeks of nursing home care leading to additional expenses of (\$188 − \$56) $\times$ 7 days $\times$ 2.8 weeks = \$2,587. On top of that, Medicaid may have to reimburse the patient expenditures stemming from the difference between the private and the Medicaid rate. This adds another (\$218 − \$188) $\times$ 7 days $\times$ 30.83 weeks = \$6,468. Taken together, this policy may increase spending by \$9,055 per Medicaid enrollee, or \$1.6bn overall.

**Episode-Based Reimbursements:** Finally, we simulate the effects of a small shift from per-diem reimbursement to an episode-based reimbursement approach. In this counterfactual, we reduce the daily Medicaid reimbursement rate by 1% and compensate the provider for the forgone variable profits with an up-front payment. This compensation maintains the profitability of Medicaid beneficiaries and counteracts a provider's incentive to respond along other margins such as changing the quality of care, see Hackmann (2017). Specifically, the change in variable profits combines two components, described below:

$$\begin{aligned}\Delta_1 &= (\$188 - \$107) \times 7 \times \Delta\text{weeks} \\ \Delta_2 &= 1\% \times \text{weeks}^{\text{new}} \times \$188\end{aligned}$$

The first component, $\Delta_1$, denotes the Medicaid rate markup over marginal costs times the change in the length of stay. The second component captures the lower Medicaid rates for the counterfactual length of stay. Finally, the overall compensation, $\Delta^{1\%}$, is simply the sum of the two components. We then replace $V(M, oc')$ by $V(M, oc') + \Delta^{1\%}$ in equations (10) and (12). Intuitively, the provider receives the up-front compensation whenever a new Medicaid beneficiary arrives or a private payer transitions into Medicaid.

The simulated discharge rates in Figure 5d suggest a large increase in provider discharge efforts. The new kink point is at 85% occupancy instead of 89% under the status quo. We find a reduction in length of stay of Medicaid beneficiaries of about 1.4 weeks and a corresponding decline in average occupancy to 90%, as indicated in the last column of Table 3. Taking the compensation payments into account, changes in Medicaid spending per nursing home stay are given by the difference between the marginal cost of nursing home care and the cost of alternative community care scaled by the 3.8 week change in the length of stay. This suggest savings of about $(\$107 - \$56) \times 7$ days $\times 1.4$ weeks $= \$500$ under this policy, and total annual spending is reduced by at least \$0.09 billion.[22] This policy, which increases financial discharge incentives among providers, therefore achieves the largest cost saving among the counterfactuals considered here.

Furthermore, we find that transitioning only 3.2% of per-diem payments to an up front episode-based reimbursement is as effective as the voucher program in reducing the average length of a Medicaid stay to the average length of a private stay. However, the implied cost savings are substantially larger. In this case we find cost savings worth \$1,950 per stay or about \$0.34 billion in total annual savings. Therefore, we conclude that transitioning to an

[22] More precisely, Medicaid savings equal \$188 − \$56 per reduced day of nursing home care plus the 1% reimbursement savings on remaining weeks of nursing home care, $\Delta_2$. Subtracting the compensation payments to nursing homes $\Delta^{1\%} = \Delta_1 + \Delta_2$ yields net savings of \$107 − \$56 per reduced day of nursing home care.

episode-based reimbursement model is more effective than increasing resident cost-sharing in shortening the length of Medicaid stays.

# 9 Conclusion

How do patient and provider incentives affect long term care utilization and health care spending? We answer this question in the context of the U.S. nursing home industry. Our findings indicate that resident cost-sharing incentives and profit incentives of nursing homes affect the length of nursing home stays among relatively healthy residents who can return to the community.

We find no evidence for improved health outcomes from longer stays. If anything our findings point in the opposite direction, suggesting that a substantial part of the Medicaid expenditures that burden state budgets may be wasteful. Our evidence on provider incentives suggests that excess capacity of nursing homes may contribute towards excessive nursing home utilization, providing an argument for Certificate of Need laws, which restrict entry and capacity investments of nursing homes. This observation is consistent with an older claim made by Roemer (1961) on provider financial incentives in hospitals: "in an insured population, a hospital bed built is a filled bed."

Many states have recognized that it is more cost-effective to reimburse home health care for Medicaid-eligible seniors instead of paying for long nursing home stays. Programs that are authorized under section 1915(c) waivers provide funding for community-based LTC. Most of these programs were started after our sample period that ends in 2005, but our results can nevertheless speak to the likely success of these new policies. While these programs largely affect patient incentives, our findings indicate that incentivizing providers may be more cost-effective in reducing nursing home utilization. Specifically, our findings suggest that transitioning from per-diem to episode-based provider reimbursement is more effective than increasing resident cost-sharing in shortening Medicaid stays. To investigate the effects of recent policy changes, we intend to expand our analysis using more recent data in future research and to account for different implementations of LTC vouchers.

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# For Online Publication

## A Home Health Care Programs

Home and community based services (HCBS) range from intensive LTC such as round-the-clock services and private duty nursing to temporary services including home health aides and help with housekeeping tasks, but may also encompass home-delivered meals, support of informal caregivers, and home accessibility adaptations. States offer various programs covering HCBS in order to reduce their Medicaid LTC spending and to offer eligible seniors more flexibility in meeting their LTC needs. Under waivers based on Section 1915(c) of the Social Security Act, states have developed programs that cover services such as adult day care, assisted living, home modifications, home delivered meals, and respite for informal caregivers.[23] 1915(c) waivers explicitly require states to demonstrate that the covered HCBS are cheaper to provide than institutional LTC.

In addition, the Money Follows the Person (MFP) demonstration program makes grants available to states to pursue programs that pay for HCBS of previous nursing home residents.[24] MFP is therefore closely related to the discharge incentives of existing nursing home residents. By subsidizing HCBS, it makes prolonging a nursing home stay relatively less attractive from the resident's perspective.

Overall, MFP and programs authorized under 1915(c) waivers increase seniors' freedom of choice of LTC arrangements and therefore approximate a policy that provides LTC vouchers. Under a typical voucher program, Medicaid beneficiaries would receive a lump-sum payment that they can spend on their preferred LTC type while incurring the price that a private payer may pay. In practice, seniors covered by Medicaid do not face private fees, but otherwise the existing HCBS programs resemble to a policy that provides LTC vouchers. We simulate nursing home utilization under a voucher program in Section 8.2.

## B Details for Theoretical Discussion

In this section, we describe the key predictions of our model in greater detail. See Section 3 for the model set-up. Specifically, we assume that nursing homes and residents can exert positive levels of effort, $e^{SNF}, e^{res} \geq 0$, to increase the probability of a discharge, $D = 1$:

$$Pr[D = 1|e^{SNF}, e^{res}] = D^{exog} + \alpha e^{SNF} + \beta e^{res}, \tag{16}$$

[23]See https://www.medicaid.gov/medicaid/hcbs/authorities/1915-c/index.html for details and https://www.cms.gov/Outreach-and-Education/American-Indian-Alaska-Native/AIAN/LTSS-TA-Center/info/1915-c-waivers-by-state.html for a list of waivers by state.

[24]See https://www.medicaid.gov/medicaid/ltss/money-follows-the-person/index.html for details. We cannot study the effects of MFP because it started in 2009, i.e. after our sample period.

where $D^{exog}$ denotes an exogenous discharge rate in the absence of any positive effort. We assume that efforts weakly increase the discharge probability: $\alpha, \beta \geq 0$.

## B.1 Resident Effort

Residents trade off the utility from staying an additional week in the nursing home versus returning to the community. We assume the following indirect conditional utility:

$$W(\tau) = \begin{cases} u - \kappa p^{\tau} + \eta^{SNF} & \text{if stay} \\ \eta^{home} & \text{otherwise} \end{cases},$$

where $u$ is the resident's gross utility from nursing home care and we normalize utility from home health care to zero. $\kappa$ is a price coefficient and $p^{\tau}$ is the per-period price paid by the resident (i.e. $p^M = 0$). $\eta^{SNF}$ and $\eta^{home}$ are type I extreme value taste shocks. Residents prefer to be discharged and hence exert strictly positive effort if $u - \kappa p^{\tau} \leq \eta^{home} - \eta^{SNF}$.

## B.2 Provider Effort

Here, we show that the Medicaid discharge rate increases in the occupancy rate above some occupancy threshold $oc^*$—as shown in Figure 1 in the main text—under simplifying assumptions that yield a closed-form solution. Specifically, we assume that the occupancy rate is fixed, newly admitted residents are private payers, there are no payer type transitions, and the exogenous discharge rate as well as the resident's discharge effort are equal to zero. Hence, a resident is only discharged if the nursing home provides strictly positive effort. The focal bed can either be empty, $\tau = 0$, or filled with a private payer or Medicaid beneficiary: $\tau = P, M$. We assume that providers exert discharge effort during the period, but that discharges continue to be stochastic and are realized at the time as new arrivals at the end of the period. We can then define the following Bellman equation:

$$V(\tau, oc) = \begin{cases} \frac{\Pi(P)}{1-\delta} & \text{if } \tau = P \\ \max_{e \geq 0} \{\Pi(M) - c(e) + D(e)V(0, oc) + (1 - D(e))\delta V(M)\} & \text{if } \tau = M \\ \delta[\phi(oc)V(P, oc) + (1 - \phi(oc))V(0, oc)] & \text{if } \tau = 0 \end{cases},$$

where $\Pi(\tau)$ is the payer type-specific per-period profit, $c(e)$ denotes the provider's cost of effort, $D(e)$ is the discharge probability as a function of the SNF's effort, $\phi(oc)$ is the probability of refilling a vacant bed, and $\delta$ is the discount factor. Note that the nursing never has an incentive to discharge a private payer in this model, which leads to the functional form of $V(P, oc)$.

For Medicaid-covered residents, the nursing home has no incentive to exert strictly positive effort below occupancy level $oc < oc^*$ because the refill probability is too low and the option value of vacating a bed does not compensate for forgone Medicaid profits. Hence, $V(M, oc) = \frac{\Pi(M)}{1-\delta}$ for $oc < oc^*$. For $oc \geq oc^*$, we have the first order condition:

$$c'(e) = D'(e)[V(0, oc) - \delta V(M, oc)].$$

Assuming $c(e) = e^2$ and with $D'(e) = \beta$, see equation (16), we have

$$e^* = \frac{\beta}{2}[V(0, oc) - \delta V(M, oc)]$$

and

$$V(M, oc) = \frac{\Pi(M) - c(e^*)}{1 - \delta(1 - D(e^*))} + \frac{D(e^*)}{1 - \delta(1 - D(e^*))} V(0, oc)$$

Defining:

$$F = e^* - \frac{\beta}{2}[V(0, oc) - \delta V(M, oc)] = 0,$$

we have $dF/de^* = 1$ as $V(M, oc)/de^* = 0$ because of the first order condition. We also have

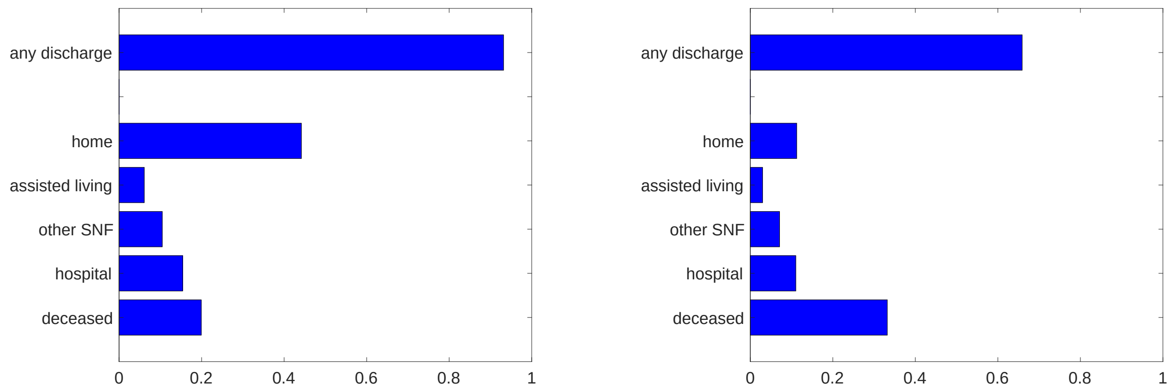
$$\frac{dF}{doc} = -\frac{\beta}{2}\left[1 - \frac{\delta D(e^*)}{1 - \delta(1 - D(e^*))}\right]\frac{dV(0, oc)}{doc}.$$

Since we have $dV(0, oc)/doc > 0$ and $\left[1 - \frac{\delta\mu}{1-\delta(1-\mu)}\right] > 0$, we get $dF/doc < 0$. This implies $de^*/doc > 0$ based on the implicit function theorem. Hence, provider efforts and consequently Medicaid discharge rates increase in the occupancy rate for $oc \geq oc^*$.

# C Data Appendix

In this Section, we provide additional information about our data sources and describe the sample in more detail.

The quarterly or more frequent nursing home resident assessment in the MDS provide information on a large number of health measures that cover a variety of cognitive, physical functional, behavioral, communication, and disease-related conditions. We reduce the wealth of health measures to a few key statistics that are commonly used in Medicaid and Medicare reimbursement methodologies. Most importantly, these include the resident's CMI, which is normalized to one and summarizes the expected resource utilization relative to the average resident. We also consider five other health measures that enter into the calculation of the CMI: (i) physical disabilities that are measured by the amount of required help with activities of daily living (ADL) such as toileting or assistance with eating, bed mobility, and transferring, (ii) whether the resident requires clinically complex treatments such as




(a) Discharge Probabilties on the Stay Level (b) Discharge Probabilities Weighted by Length of Stay

Figure 6: Discharge Probabilities by Destination

***Notes:*** *The figure summarizes the propensity of different discharge reasons. Any discharge is an indicator that turns on if the stay ends with a discharge. The remaining stays are censored. "Home", "Assisted Living", "Other SNF", "Hospital", and "Deceased" are indictor variables that turn on if the stay ends with a discharge to the community, to an assisted living facility, a different SNF, a hospitals or if the resident deceased, respectively. The figures shows the fraction of stays that end accordingly. In the left figure, the unit of observation is the nursing home stay. In the right figure, the unit of observation is week of the nursing home stay.*

chemotherapy, oxygen therapy, or dialysis, (iii) depression, (iv) impaired cognition, and (v) behavioral problems.

The MDS also indicates the admission and discharge dates for each resident, which allows us to construct the exact length of each nursing home stay. We observe a discharge code, which provides information on the reason of discharge and the institution the resident was discharged to. Figure 6a displays stay-level discharge probabilities overall and by destination. In our sample, 93% of nursing home residents have been discharged by the end of the sample period. Almost half of them return to the community, 20% of residents pass away in the nursing home, and the remaining residents enter an assisted living facility, another nursing home, or a hospital upon being discharged. When weighting the discharge probability by length of stay, i.e. using week-level discharges, Figure 6b shows a different pattern. Specifically, we find that only 65% of residents have been discharged by the end of the sample period. Since residents' health status deteriorates over time, fewer SNF stays end in a home discharge (11%) and more stays end through the resident's death (33%) when weighting by the length of stay, thereby explaining the differences between Figures 6a and 6b.

To investigate differences in discharge patterns between residents with different payer sources, we combine the MDS with administrative Medicare and Medicaid claims data from the MedPAR and MAX databases, respectively. We link the information at the nursing home-stay level and are thus able to quantify the days of each stay covered by Medicare and Medicaid. We assume that the remaining days are paid out-of-pocket, given that only a small fraction of residents have private LTC insurance. Importantly, we observe the exact transition date between private pay and Medicaid eligibility, which is key for our empirical analysis.

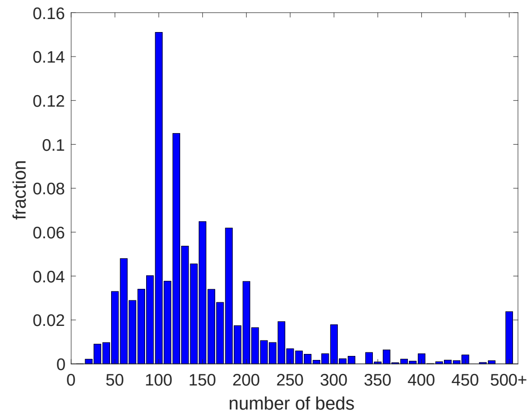
To quantify the occupancy rate of each nursing home at any point in time, we combine admission and discharge date information from the MDS with information on the number of licensed beds from the On-Line Survey, Certification, and Reporting system (OSCAR). This survey provides information from state surveys on all federally-certified Medicaid and Medicare nursing homes in the U.S. (see, e.g., Grabowski, 2001). Figure 7 presents a histogram for the number of licensed beds. While about 30% of all SNFs have between 100 and 120 beds, there is substantial variation in facility size. Table 4 summarizes within nursing home week-to-week variation in the occupancy rate. The cells document the relative frequency compared to any transition between 90% and 100% occupancy. The cells on the main diagonal sum to 58% suggesting that occupancy rate changes from week-to-week of one or more percentage points occur 42% of the time.s

We also use data from distinct nursing home surveys from California and Pennsylvania, which collect information on private and Medicaid rates not included in OSCAR (see Hackmann, 2017).[25] For California SNFs, we infer daily private and Medicaid rates by dividing annual revenue by the number of resident-days for each payer type. The average daily private rate amounts to $218 and $158 in Pennsylvania and California, respectively, and the corresponding average Medicaid rates are $188 and $127. These differences indicate that the profitability varies by payer types, a fact that we leverage in some of our empirical analyses.

To investigate potential differences in access to community-based care further, we supplement Table 1 with data from two surveys. First, we select respondents from the HRS who have spent at least one night in a nursing home in the two years prior to a given survey wave. Table 5 shows selected summary statistics by payer type.[26] We first consider the role of respondents' children as potential informal caregivers. The fractions of respondents who receive help with ADL or IADL are roughly similar across payer types. Private payers expect more help in the future, but Medicaid beneficiaries are more likely to receive financial support to pay for medical bills and receive more help at the intensive margin. Overall, there

[25]The Pennsylvania survey data were provided by the Bureau of Health Statistics and Research of the Pennsylvania Department of Health. California data come from the Office of Statewide Health Planning and Development (see http://www.oshpd.ca.gov/HID/Products/LTC/AnnFinanclData/PivotProfls/default.asp).

[26]HRS respondents are designated as Medicaid beneficiaries if they report being covered by a Medicaid plan. We define them as private payers if their nursing home stay was not covered any type of government plan (Medicaid or Medicare) or if they hold private long term care insurance.




***Notes:*** *The figure presents a histogram of the overall number of licensed beds. The unit of observation is the week of the nursing home stay.*

Figure 7: Number of Licensed Beds

Table 4: Week-to-Week Transitions in Occupancy 90%-100%

| | 90% | 91% | 92% | 93% | 94% | 95% | 96% | 97% | 98% | 99% | 100% |
|---|---|---|---|---|---|---|---|---|---|---|---|
| 90% | 3.51% | 0.93% | 0.63% | 0.36% | 0.16% | 0.08% | 0.04% | 0.02% | 0.01% | 0.00% | 0.00% |
| 91% | 0.77% | 3.75% | 1.05% | 0.73% | 0.37% | 0.18% | 0.07% | 0.03% | 0.01% | 0.00% | 0.00% |
| 92% | 0.49% | 0.83% | 4.55% | 1.24% | 0.71% | 0.40% | 0.17% | 0.08% | 0.03% | 0.01% | 0.00% |
| 93% | 0.31% | 0.57% | 1.02% | 5.48% | 1.32% | 0.86% | 0.39% | 0.19% | 0.06% | 0.02% | 0.00% |
| 94% | 0.17% | 0.34% | 0.59% | 1.08% | 5.66% | 1.46% | 0.84% | 0.36% | 0.14% | 0.03% | 0.01% |
| 95% | 0.11% | 0.19% | 0.35% | 0.70% | 1.22% | 6.35% | 1.55% | 0.89% | 0.34% | 0.08% | 0.02% |
| 96% | 0.05% | 0.09% | 0.18% | 0.36% | 0.70% | 1.33% | 7.18% | 1.54% | 0.79% | 0.21% | 0.03% |
| 97% | 0.03% | 0.04% | 0.09% | 0.21% | 0.37% | 0.74% | 1.29% | 7.47% | 1.61% | 0.47% | 0.07% |
| 98% | 0.02% | 0.02% | 0.05% | 0.10% | 0.17% | 0.34% | 0.72% | 1.35% | 8.45% | 1.13% | 0.14% |
| 99% | 0.00% | 0.01% | 0.01% | 0.03% | 0.04% | 0.10% | 0.22% | 0.46% | 1.03% | 5.14% | 0.28% |
| 100% | 0.00% | 0.00% | 0.00% | 0.01% | 0.01% | 0.02% | 0.04% | 0.06% | 0.14% | 0.26% | 0.87% |

***Notes:*** *The table summarizes within nursing home week-to-week variation in the occupancy rate. Current and next week's occupancy rate are displayed in rows and columns, respectively. Each cell documents the relative frequency, compared to any transition between 90 and 100% occupacny. For expositional reasons, occupancy variation is only shown between 90% and 100%.*

Table 5: Summary Statistics: Informal Caregivers, Income, and Asset (HRS)

| | Medicaid | | Private | |
|---|---|---|---|---|
| | Mean | SD | Mean | SD |
| Kids help with ADL | 0.203 | (0.402) | 0.131 | (0.338) |
| Kids help with IADL | 0.157 | (0.364) | 0.144 | (0.351) |
| Kids will help in future | 0.0433 | (0.204) | 0.141 | (0.348) |
| Kids help with health care cost | 0.0777 | (0.268) | 0.0279 | (0.165) |
| Hours per month kids helping | 30.11 | (99.79) | 21.67 | (80.95) |
| Home ownership | 0.274 | (0.446) | 0.597 | (0.491) |
| | Mean | SD | Mean | SD |
| | [p25, p50, p75] | | [p25, p50, p75] | |
| Social security income | 7018.1 | (4345.5) | 5881.3 | (3837.2) |
| | [4704.8, 6891.8, 9508.0] | | [2861.1, 6528.9, 9199.1] | |
| Total household income | 11933.6 | (11235.4) | 19214.1 | (10292.3) |
| | [6718.3, 9235.1, 13703.3] | | [11271.1, 17718.8, 26515.2] | |
| Net financial assets | 5495.2 | (51800.3) | 27845.8 | (45702.0) |
| | [0, 0, 854.2] | | [0,4750.8, 50015.1] | |
| Total wealth | 32989.8 | (119989.1) | 130945.6 | (142362.5) |
| | [0, 394.9, 15846.6] | | [3316.8, 81361.9, 237699.4] | |
| Observations | 1,693 | | 526 | |

***Note:*** *All income and asset amounts in 2000 dollars.*

***Source:*** *Health and Retirement Study 1998 to 2010.*

Table 6: Summary Statistics: Monthly Income and Assets (NLTCS)

| | (1) Mcd/Mcd | | (2) Prv/Mcd | | (3) Prv/Prv | | (4) |
|---|---|---|---|---|---|---|---|
| | Mean | SD | Mean | SD | Mean | SD | $p$-value |
| Total income | 622.6 | (582.6) | 792.2 | (585.7) | 1117.3 | (1516.3) | 0.0518 |
| Social Security benefits | 620.8 | (367.3) | 770.7 | (319.6) | 872.0 | (545.1) | 0.1190 |
| Other retirement income | 60.33 | (203.1) | 153.0 | (361.0) | 337.9 | (875.8) | 0.0636 |
| SSI | 10.94 | (125.7) | 0 | (0) | 10.88 | (101.0) | 0.4396 |
| Home ownership | 0.0819 | (0.280) | 0.0899 | (0.295) | 0.160 | (0.376) | 0.2933 |
| Net home value | 5264.2 | (25189.0) | 3995.2 | (18530.6) | 23287.9 | (115402.0) | 0.2368 |
| Observations | 203 | | 48 | | 266 | | |

***Notes:*** *p-value from the F-test of the null hypothesis* $H_0 : \beta_P = \beta_M$ *in the regression* $Y_i = \beta_0 + \beta_P Prv_i + \beta_M Mcd_i + \sum_t \delta_t \mathbf{1}\{LOS_i = t\} + \epsilon_i$, *where* $Prv_i$ *and* $Mcd_i$ *are indicators for payer type at the interview, given that payer type at admission is private (excluded category is Medicaid at admission). All amounts in 2004 dollars.*
***Source:*** *National Long Term Care Survey 1999 and 2004.*

are no systematic differences in informal care.

Second, we consider nursing home residents' income and assets. Using data from the HRS again, the bottom part of Table 5 indicates that private payers have higher income and asset levels than Medicaid beneficiaries and are more likely to own a home. These differences are less substantial, however, when taking into account the cost of nursing home care. The income difference at the median would pay for a nursing home stay of about one month. In practice, most private payers would pay down their assets. At median asset levels, they could afford about one year of SNF care before qualifying for Medicaid. The comparison between private and Medicaid residents in Table 5 does not reflect the income and asset distributions in our MDS sample because the HRS sample includes SNF residents who qualify for Medicaid at the beginning of their stay, thereby biasing the income and asset measures downward. In our MDS sample, Medicaid and private resident are therefore more homogenous with respect to income and assets than in the HRS.

Using data from the NLTCS of 1999 and 2004, we can construct a sample that is more comparable to the MDS. In contrast to the HRS, the NLTCS samples individuals who are currently residing in a nursing home. Moreover, the NLTCS contains information on the payer type at admission and at the time of the interview, so it allows us to observe payer type switches. Table 6 shows nursing home residents' average income and assets by payer type including residents who switch from private to Medicaid during their nursing home stays. Since we drop residents who are covered by Medicaid at admission in our main analysis, we are particularly interested in any differences between column (2) and (3) in Table 6. The length of stay adjusted $p$-values in column (4) show that none of the income and asset measures are statistically significantly different at the 95% level between payer types. However, we note that the averages point to higher home ownership rates among private payers. We revisit the effect of remaining differences between payer types on our main findings in Section 6.3.

# D Robustness Checks and Additional Empirical Results

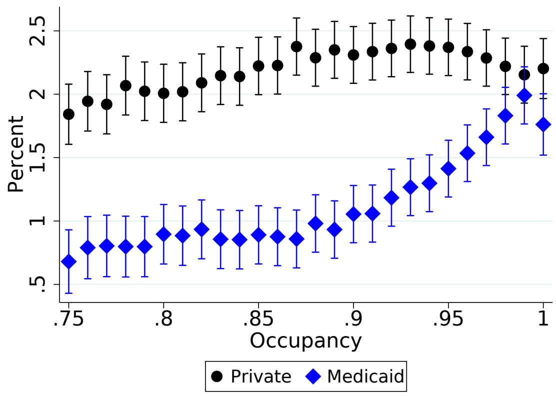
## D.1 Discharge Patterns to Other Destinations

We have argued above that financial discharge incentives of providers and patients are most likely to operate in the case of residents who ultimately return to the community. The results shown in Figure 3 confirm this hypothesis. Here, we estimate regression (2) with indicators for different types of discharges as the dependent variable to verify that discharges that are less likely to be affected by financial incentives do not exhibit the same pattern. Figure 8a repeats the results from Figure 3.

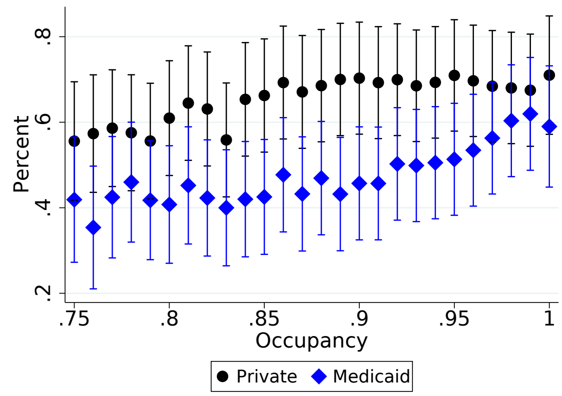
Compared to these home discharge patterns, we observe smaller differences between private and Medicaid residents in discharges to a hospital, another nursing home, and due to the resident's death in Figures 8b, 8c, and 8d, respectively. While private payers are discharged to all destinations more frequently, we find that the difference to Medicaid discharge rates amounts to less than half a percentage point. In contrast, home discharge rates differ by over one percentage point at low occupancy rates. We also note that the patterns for discharges to a hospital, another nursing home, and due to the resident's death exhibit a kink for Medicaid residents around an occupancy rate of 90%. These kinks are much less pronounced than for home discharge rates, however, and are likely due to advantageous selection. Since mostly healthy Medicaid beneficiaries are discharged back to the community at high discharge rates, the Medicaid residents who remain in the nursing home are less healthy and therefore more likely to be hospitalized and to die, see also Section 6.2 and Section D.2 below. The overall discharge patterns in Figure 8e are similar to home discharge rates, but from Figures 8a to 8d it is clear that the sharp increase in Medicaid discharges above 89% occupancy is primarily driven by discharges to the community.

## D.2 Discharge Rates Among Low-Care Residents

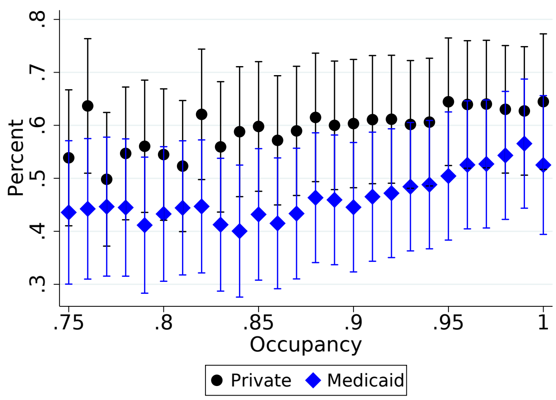
To further verify that Medicaid residents with high LTC needs are not the ones who are discharged at high occupancies, we consider a homogenous subsample of nursing home residents



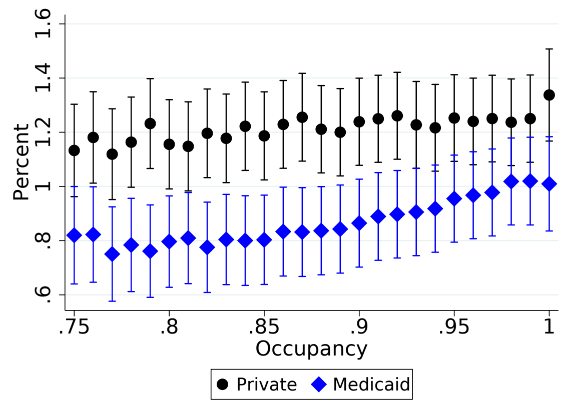

(a) Discharged Home



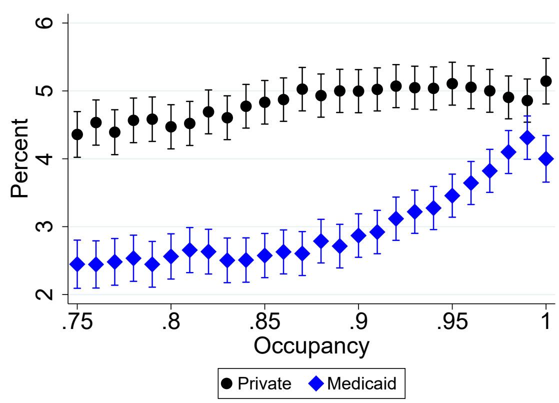

(b) Hospitalized




(c) Discharged to Another Nursing Home




(d) Deceased




(e) Discharged to Any Destination

***Notes:*** *See notes for Figure 3. The dependent variables are indicator variables that equal one if a resident was discharged to the community (Figure 8a), to a hospital (Figure 8b), to a different nursing home (Figure 8c), deceased (Figure 8d), or discharged overall (any of the above, Figure 8e) in a given week. The vertical bars indicate 90% confidence intervals.*

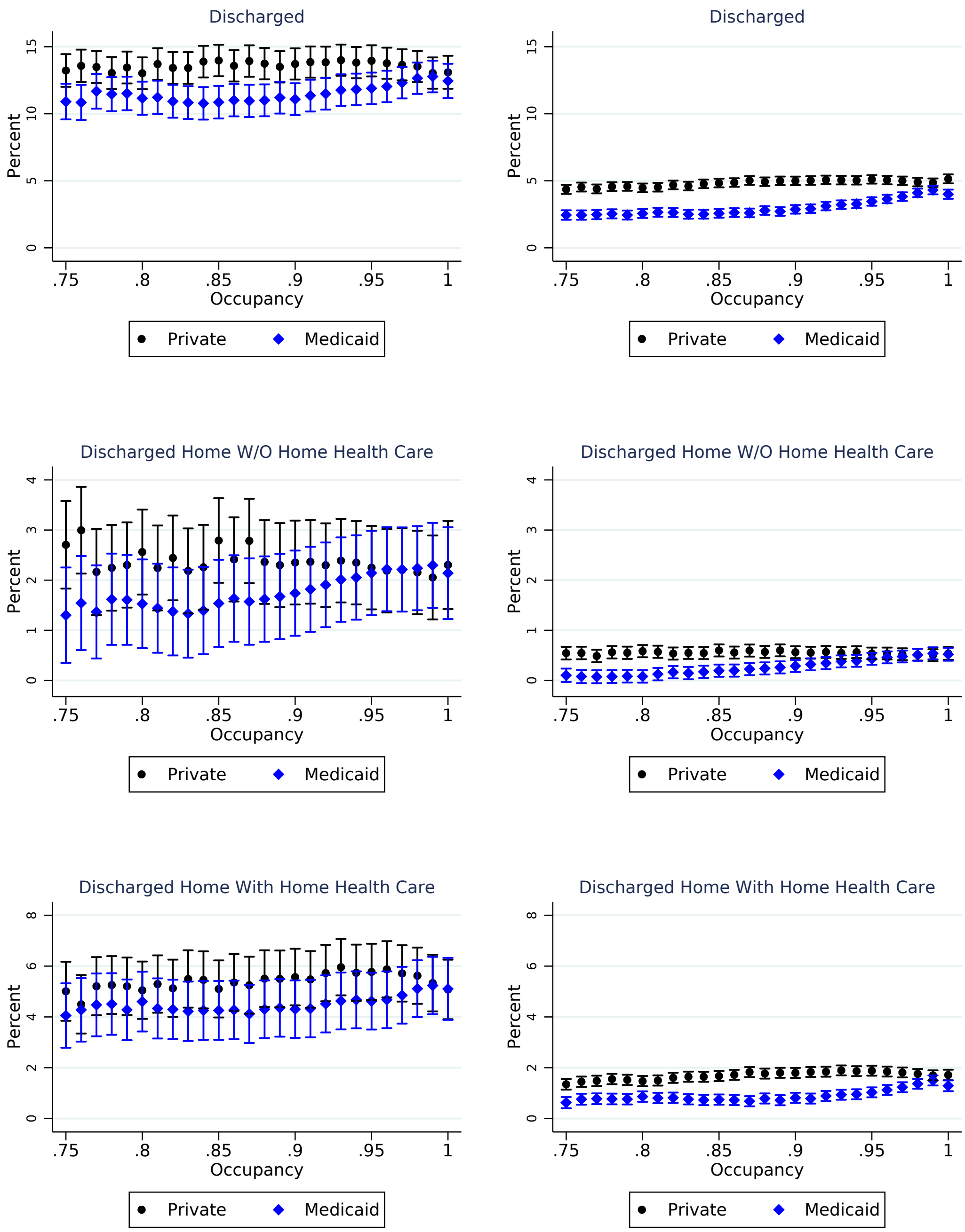
Figure 8: Discharge Rates to Different Destinations by Occupancy and Payer Type

with very little LTC needs. Thereby, we can further isolate the effect financial incentives on nursing home utilization. To identify nursing home residents who have low LTC needs, we follow the methodology developed by Mor et al. (2007) who define residents belonging to a broad category of low care need as those who do not require physical assistance in four ADLs (bed mobility, transferring, using the toilet, and eating) and are not classified as "special rehabilitation" or "clinically complex" residents. Residents who fall into this category have very low care needs and should not receive care in a SNF from a medical point of view but rather use home health care. If we nevertheless observe these patients in a nursing home, they should have very short lengths of stay and therefore high discharge rates. Hence, low discharge rates among low-care Medicaid residents point to overuse of LTC.

Figure 9 shows three key discharge outcomes for the low-care sample on the left and for comparison purposes for the whole sample on the right: discharges overall and community discharges with and without home health care. First, overall discharge rates among low-care residents are almost three times as high as for the whole sample reflecting the fact that the low-care residents have shorter stays, as indicated in the first row. Second, the community discharge rates for private payers in the low-care sample are substantially higher than for Medicaid beneficiaries at low occupancy rates, see the middle and the bottom left figures. This is particularly the case for residents who do not need LTC in the community, see the middle graph. This result provides further evidence that longer stays of Medicaid beneficiaries contribute to LTC overspending since differences in unobserved health status correlated with payer type are unlikely to explain the observed difference in discharge rates given a highly selected sample with homogenous health profiles. Compared to the discharge differences at low occupancy rates, we find an increase in Medicaid discharge rates at higher occupancies, suggesting again that both consumer and provider profit incentives contribute to longer stays among residents who have very low care needs and should not receive care in a SNF from a medical point of view.

## D.3 Effect of Occupancy Rates and Payer Type on Health Outcomes

Next, we provide evidence on selection on health types by analyzing how current health profiles vary with occupancy. We can thereby test if systematic differences in health between payer types and across occupancies may lead to the observed discharge patterns in Figure 3. Specifically, if Medicaid residents are healthier at higher occupancy rates they may be discharged disproportionately not due to financial incentives of the SNF, but rather because they have lower LTC needs. To test for selection based on health, we estimate regression (2) with the respective health outcome as the dependent variable. The main difference to the




***Notes:*** *See notes for Figure 3. This figure compares overall discharges (first row), home discharges without home health care (middle row), home discharges with home health care (bottom row) between resident groups. In the left figures, we focus on residents with low care needs and the right graphs we present the evidence for the baseline sample. The vertical bars indicate 90% confidence intervals.*

Figure 9: Discharge Rates Among Low Care Need Residents and All Residents

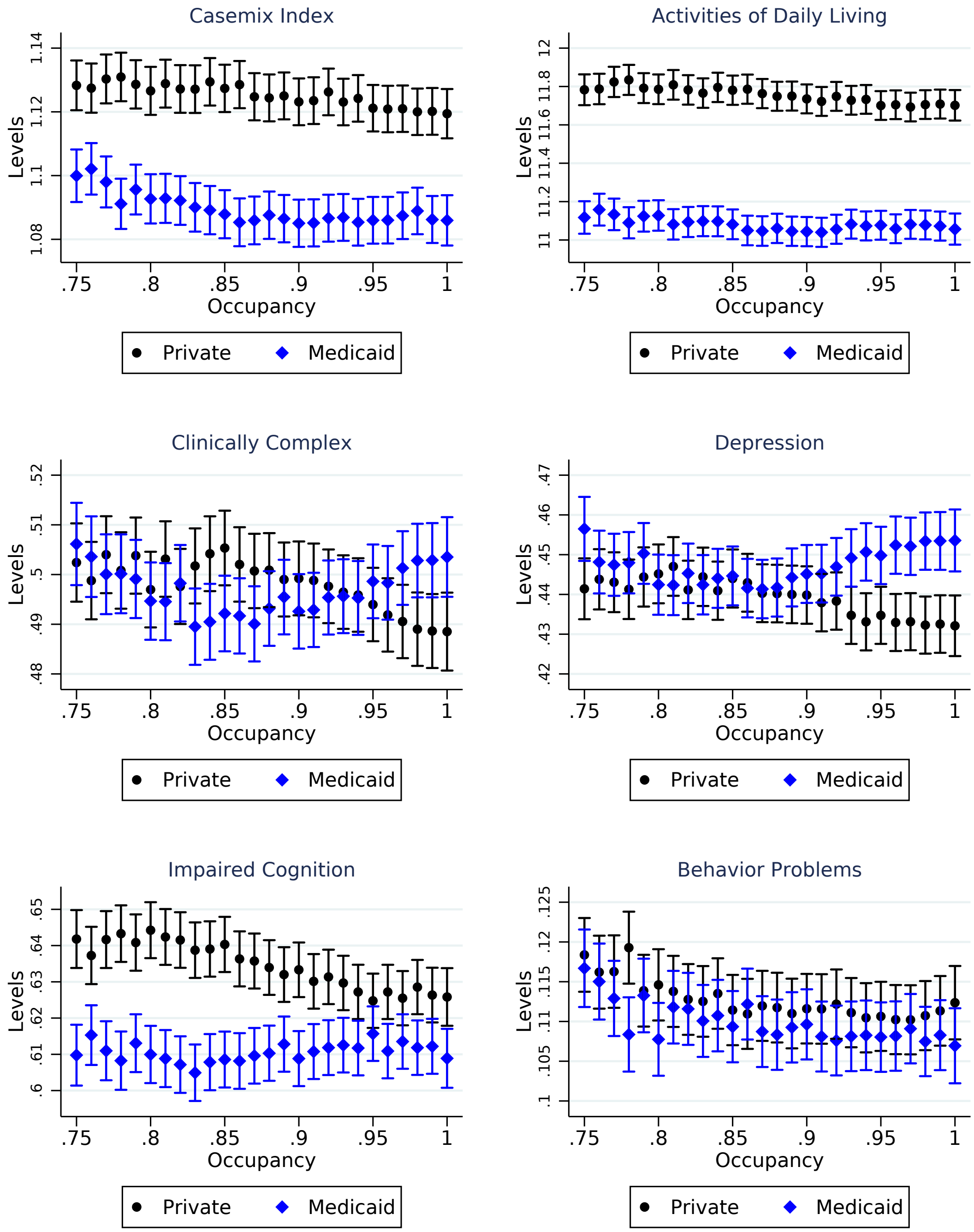
results from the analysis shown in Figure 3 is that we do not control for the health profile at admission in this analysis. We consider six health outcomes from residents' most recent assessment: the CMI, ADL limitations, and indicators for clinical complexity, depression, impaired cognition, and behavioral problems. For all health outcomes, a higher value implies that the resident is less healthy.

The top-left panel in Figure 10 shows that private payers are sicker than Medicaid beneficiaries across the entire occupancy range according to the CMI, a summary measure, conditional on nursing home-year and week-of-stay fixed effects. The CMI differs by less than one tenth of a standard deviation between payer types and across all occupancy rates (see Table 1 for summary statistics of the CMI and other health measures). The graph also indicates that the CMI is slightly decreasing in occupancy for both private and Medicaid residents, i.e. residents are on average slightly healthier at higher occupancy rates. Overall, the CMI trends look very similar between payer types. The downward trend flattens out for Medicaid payers at about 90% occupancy, which indicates that disproportionately healthier Medicaid residents remain in the nursing home, when compared to private payers, at high occupancies.

In addition to the CMI as a summary measure, we test for selection based on health type for five specific health outcomes. The remaining panels in Figure 10 do not show many systematic changes in health outcomes as occupancy rates increase that would indicate selection based on health. For clinical complexity, depression, and impaired cognition, the health difference between private and Medicaid residents declines at higher occupancy rates, and Medicaid beneficiaries are less healthy at the highest occupancy rates when it comes to clinical complexity and depression. The health status differences do not exceed 5% of a standard deviation, however. Overall, we find some evidence for selective discharges based on health types but, the evidence is not as striking as the results on selection on payer type presented in Section 6.1, which suggests that mostly healthy Medicaid beneficiaries are discharged to the community.

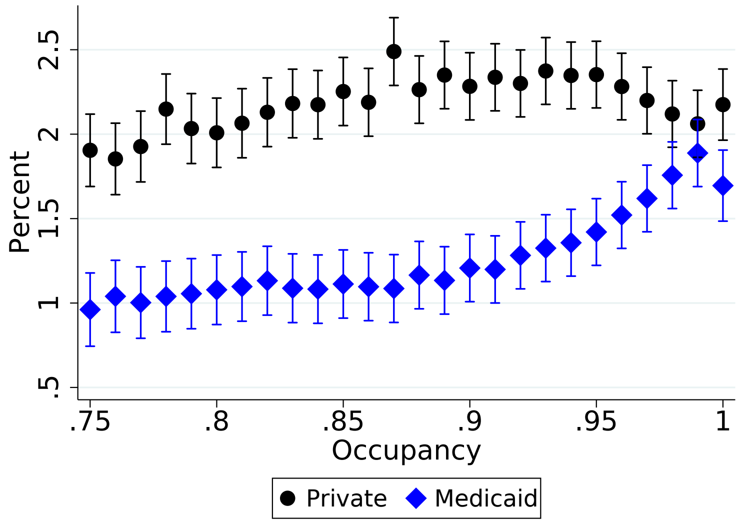
## D.4 Patient Incentives: Propensity Score Analysis

In this section, we revisit the baseline estimates in a population of private payers and Medicaid beneficiaries that is even more homogenous than in our baseline results. To this end, we a adopt a propensity score weighting approach. We first estimate the probability of being Medicaid eligible in the given week of the nursing home stay using a flexible logit model that conditions on the highest education level, race, and zip code of former residence fixed effects. All right hand side characteristics remain unchanged during the nursing home stay. Therefore, the predicted probability $\hat{p}_i$ does not vary across weeks of the stay. The logit model




***Notes:*** *See notes for Figure 3. This figure compares the most recent health profile at the nursing home week level between payer types and occupancy rates. The resident's health status is decreasing in each health measure. The CMI is a summary measure of long term care needs, calculated based on methodology 5.01, and normalized to 1. The remaining health measures are direct inputs to the CMI formula and provide more granular information on cognitive and physical disabilities. The vertical bars indicate 90% confidence intervals.*

Figure 10: Most Recent Health Profiles by Payer Type and Occupancy




***Notes:*** *See notes for Figure 3. The dependent variables are indicator variables that equal one if a resident was discharged to the community. The underlying regression model, see equation (2), is weighted by a propensity score of Medicaid eligibility. The vertical bars indicate 90% confidence intervals.*

Figure 11: Home Discharges: Propensity Score Weighted

yields delivers an $R^2$ of 7.2%. For Medicaid beneficiaries, we find the following percentiles for the predicted probabilities. At the 10th percentile, we find a Medicaid probability of 29%. This probability increases to 45% and 61% at the 50th and the 90th percentile, respectively. For private payers, the Medicaid probabilities increase from 20% to 39% and 54% at the 10th, the 50th, and the 90th percentile, respectively. The overlap between the two predicted discharge distributions is therefore very good, suggesting that an inverse propensity score weighting approach yields a relatively small bias compared to other matching algorithms (Busso, DiNardo, and McCrary, 2014).

We then weight Medicaid observations by 1 and assign a weight of $\frac{\hat{p}_i}{1-\hat{p}_i}$ to private payer observations. Intuitively, we assign private payers a higher weight if they are more likely to transition into Medicaid conditional on observables. We then estimate equation (2) using these regression weights. Our findings are summarized in Figure 11, which are very similar to our baseline findings presented in Figure 3. Since Medicaid-covered and private residents are balanced based on observables in this result, we can further isolate the role of residents' financial incentives in determining discharge rates. Specifically, we find a similar gap between discharge rates at low occupancies as in Figure 3, suggesting that this difference is indeed driven by patient moral hazard.

## D.5 Patient Incentives: Medicaid Applications

In this section, we revisit the effect of resident financial incentives on the home discharge probability among Medicaid applicants. In the MDS assessment data, we are able to infer whether a resident has applied for Medicaid coverage and is waiting on a final decision. We restrict the sample to this subsample of residents and focus on the weeks following the assessment date when a pending Medicaid application was first indicated. This sample refinement mitigates remaining differences between private payers and Medicaid. By comparing residents who have applied for Medicaid and whose application has or has not yet been approved, we effectively condition on the residents' level of assets since only residents with resources close to the Medicaid eligibility threshold would consider applying for benefits. Hence, the application outcome only affects home discharge rates through changes in financial incentives.

Estimating regression (18) on the subsample of Medicaid applicants, we find a reduction in the home discharge rates among Medicaid beneficiaries of 0.5 percentage points, $\hat{\delta} = -0.005$ (S.E. 0.01 percentage points). This estimate is slightly smaller in magnitude than the average difference in home discharge rates presented in Figure 3. Nevertheless, the point estimate confirms that financial resident incentives affect the length of stay.

## D.6 Patient Incentives: Discharge Timing

To further establish that patient financial incentives affect discharge rates, we exploit the timing of payer type transitions within nursing home stays in greater detail. That is, we estimate the week-specific effect of a payer type switch from private to Medicaid on the resident's discharge probability in subsequent weeks, conditional on the current length of stay, facility-year and calendar month fixed effects, and individual characteristics. We estimate the following regression

$$Y_{ijsmt} = \sum_{m=1}^{M} \phi^m Mcaid_{is}^m + \alpha_{jy} + \alpha_s + \alpha_c + X_i'\beta + \epsilon_{ijsmt}, \tag{17}$$

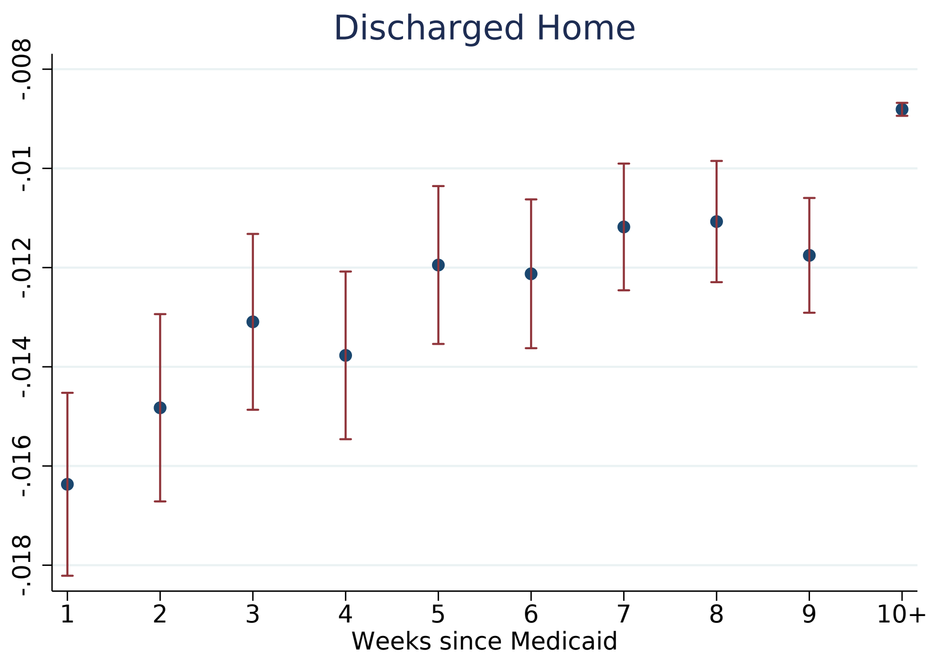
where $Y_{ijmst}$ is a discharge indicator for resident $i$ in facility $j$, week of stay $s$, week of being covered by Medicaid $m$, and calendar week $t$. $Mcaid_{is}^m$ is an indicator variable that turns on if the resident is in her $s$th week of her nursing home stay and has been covered by Medicaid for $m$ weeks. For example, if person $i$ is in the 10th week of her stay, paid the first 9 weeks out of pocket and just transitioned into Medicaid, then $Mcaid_{i10}^1 = 1$. Furthermore, $Mcaid_{is}^m = 0$ for a person who pays out-of-pocket in week $s$ of her stay. We restrict this analysis to residents who are in their 10th to 20th week of their nursing home stay to obtain a homogenous sample. Based on our theoretical model, we predict that a payer type switch lowers the discharge propensity. Moreover, if resident incentives affect home discharges, this decline should be most pronounced immediately after a resident qualifies for Medicaid. To test this prediction, we include indicators for the first 9 weeks that a resident's stay is covered by Medicaid, $\phi^m, m = 1, \ldots, 9$, as well as a medium-run effect, $\phi^{10+}$.

Figure 12 displays the corresponding discharge effects based on the week since Medicaid transition together with 90% confidence intervals. The last estimate (10+) reflects the coefficient on the overall Medicaid indicator variable, the medium term effect. All point estimates are negative ranging between $-0.8$ and $-1.6$ percentage points. Hence, Medicaid beneficiaries have smaller discharge probabilities than residents who continue to pay out of pocket. This finding is consistent with patient moral hazard as predicted by the theoretical model. Second, the point estimates are decreasing in absolute value. Specifically, we see that the largest reduction in the weekly discharge rate falls in the first Medicaid week, and this discharge rate is about twice as large as the middle run estimate. Hence, residents who become eligible for Medicaid immediately react to changing discharge incentives. Since it is unlikely that resident's LTC needs change during the same week they become eligible for Medicaid, this finding further strengthens our conclusion that the observed increases in the home discharge rate are due to patient moral hazard.

Unfortunately, we are not able to conduct a "true" event study because we only observe Medicaid eligibility conditional on no prior discharge having occurred. Therefore, the overall reduction in the discharge rate may be confounded by unobserved differences between Medicaid beneficiaries and private payers. Based on observable characteristics, private payers are generally very similar to Medicaid beneficiaries as discussed in Section 4. If anything, private payers have observably worse health profiles at admission and during their stay once we condition on the week of stay and nursing-home year fixed effects, suggesting longer stays, as we show in Section D.3 below.

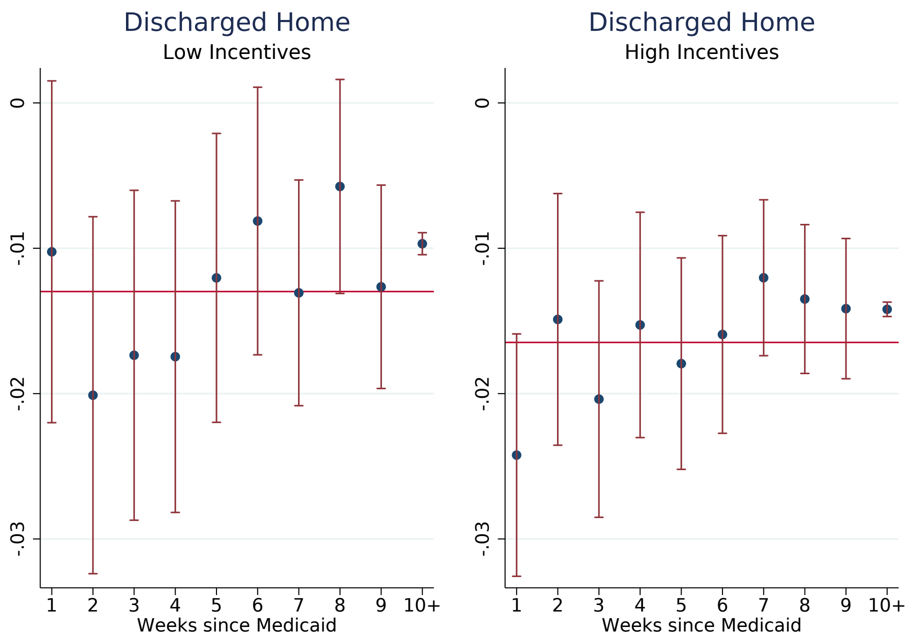
## D.7 Patient Incentives: Difference-in-Differences

Since we are not able to conduct a "true" event study, we cannot rule out that Medicaid and private residents differ in unobservables, as discussed above. To address this concern, we estimate the same regression (17) separately for nursing homes with higher than average private rates and lower than average private rates in the states of Pennsylvania and California, where we have access to price data. We expect a larger reduction in discharge probabilities in nursing homes with higher private rates because of a larger price drop upon start of a resident's Medicaid coverage. High price nursing homes therefore provide stronger incentives for consumer moral hazard. Figure 13 presents the estimates for lower priced and higher




***Notes:*** *The dependent variables are indicator variables that equal one if a resident was discharged to the community. The figure presents the average difference in discharge rates (averaged over occupancy rates) between private payers and Medicaid beneficiaries by the week since Medicaid transition. A negative difference implies a higher discharge rate for private payers. The vertical bars indicate 90% confidence intervals.*

Figure 12: Discharge Effect for Medicaid Beneficiaries by Week Since Payer Type Switch




***Notes:*** *The dependent variables are indicator variables that equal one if a resident was discharged to the community. The figure presents the average difference in discharge rates (averaged over occupancy rates) between private payers and Medicaid beneficiaries by the week since Medicaid transition. A negative difference implies a higher discharge rate for private payers. The left graph investigates residents with small difference in discharge incentives between private payers and Medicaid beneficiaries (small private rates). The right graph investigates residents with large difference in discharge incentives between private payers and Medicaid beneficiaries (high private rates). The horizontal lines display the average difference in discharge rates across weeks since Medicaid transition. The vertical bars indicate 90% confidence intervals.*

Figure 13: Discharge Effect for Medicaid Beneficiaries by Week Since Transition and Incentives

priced nursing homes in the left and right panel, respectively.

As expected, we see larger reductions in weekly discharge rates in nursing homes with higher private rates. To put this difference into perspective, the average price difference between these groups equals about $79 per day which corresponds to 43% of the average daily price of $183 over all nursing homes in Pennsylvania and California. The difference in the reduction in weekly discharge rates equals 0.0035 percentage points for the first 9 weeks under Medicaid and 0.0045 percentage points in the middle run (when comparing the point estimates on the very right only). Scaled from 43% to 100%, this suggests that consumer incentives can account for an overall reduction in weekly discharge rates of $\frac{0.0035}{0.43} \times 100 = 0.8$ percentage point points over the first 9 weeks. This accounts for about 62% of the average 1.29 percentage point reduction over the first 9 weeks observed in Figure 12.

## D.8 Patient Incentives: Border Analysis

In this section, we exploit differences in Medicaid eligibility criteria between states in a border analysis. In Pennsylvania, Medicaid beneficiaries can retain higher asset levels ($2,400 to $8,000) when compared to Ohio (a most $1,500).[27] Furthermore, Pennsylvania permits the spouse living in the community to retain a larger set of the couple's assets when considering the eligibility for Medicaid. In particular, 401k and other retirement plans of the community spouse are not counted as resources for purposes of the institutionalized spouse qualifying for Medicaid in Pennsylvania, but they are in Ohio and New Jersey. Finally, Ohio has a more aggressive estate recovery program which looks to the Medicaid recipient's assets to reimburse the state for the money spent on the individual's care provided in the nursing home.

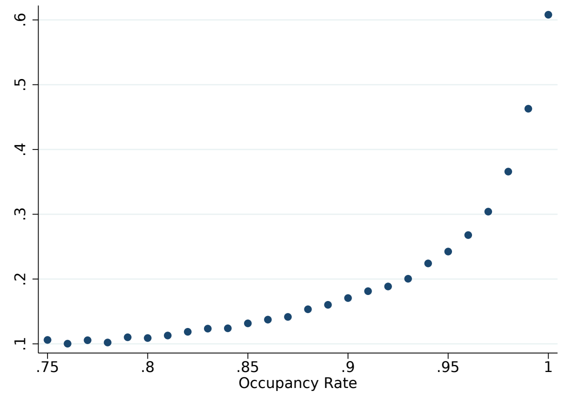
These rules indicate that Medicaid eligibility standards are more generous in Pennsylvania than in the neighboring states of Ohio and New Jersey. All else equal, we therefore expect a larger fraction of nursing home residents in Pennsylvania to be on Medicaid. To isolate this source of variation, we conduct a state border analysis and restrict the sample to seniors who resided in counties on the Ohio-Pennsylvania border or the Pennsylvania-New Jersey border prior to their nursing home stay. We then revise the baseline regression as follows:

$$Y_{ijst} = \delta \times Mcaid_{is} + \sum_{k=75}^{100} \gamma^k occ_{jt-1}^k + border_i + \alpha_c + \alpha_s + X_i'\beta + \epsilon_{ijst}. \tag{18}$$

For simplicity, we have omitted the full sequence of interaction terms between the occupancy rate fixed effects and Medicaid eligibility. Instead, we include only the Medicaid indicator, whereby $\delta$ denotes the average difference in weekly home discharge rate between payer types. We also drop nursing-home year fixed effects as we exploit cross-sectional variation between counties on either side of the border. Instead, we only control for a state-border fixed effect, $border_i$. To isolate the variation in Medicaid eligibility between states, we instrument for $Mcaid_{is}$ using a Pennsylvania and a New Jersey state indicator variable (Ohio is the omitted category).

Our first stage parameters of interest equal 0.043 (S.E. 0.0013) for the Pennsylvania indicator and −0.023 (S.E. 0.003) for the New Jersey indicator. These estimates suggests that residents residing in Pennsylvania are more likely to transition into Medicaid coverage when compared to residents from New Jersey or Ohio. Specifically, the probability of transitioning into Medicaid in Pennsylvania's bordering counties, bordering Ohio, exceeds the probability

[27] See http://hartleelderlaw.com/download/i/mark_dl/u/1038829/15106695/Ohio and PA Medicaid Differences.pdf and http://www.state.nj.us/humanservices/dmahs/clients/medicaid/medicaid_program_eligibility.pdf.




***Notes:*** *This figure plots the average weekly refill probability of an empty bed against the nursing home's occupancy rate, see equation* (3) *in Section 6.4 in the main text for details.*

Figure 14: Weekly Refill Probability by Occupancy Rate

in Ohio's bordering counties by 4.3 percentage points. The negative coefficient on the New Jersey indicator suggests that this difference is more pronounced at the Pennsylvania-New Jersey border. The IV estimate for $\delta$ in regression (18) equals $-0.009$ (S.E. 0.006). This estimate suggests that the discharge rate falls by 0.9 percentage points as seniors transition into Medicaid. This difference is similar to the average difference in discharge profiles presented in Figure 3. However, due to the additional variation introduced by the IV, the point estimate is not significant at the 10% level with a $t$-statistic of 1.5. Nevertheless, this result supports the conclusion that differences in discharge rates are partly driven by patient incentives.

## D.9 Provider Incentives

Based on the discussion in Section 6.4 in the main text, Figure 14 plots the average weekly refill probability of an empty bed on the vertical axis against the weekly occupancy rate on the horizontal axis. The figure documents a highly convex relationship highlighting an increasing option value of an empty bed at occupancies exceeding 90%. This relationship further supports the conclusion that the increasing Medicaid discharge rates at high occupancy rates are driven by provider incentives.

As an additional test of financial incentives, we compare the discharge rates across payer types between facilities that differ in how much they benefit from replacing a Medicaid beneficiary with a private payer, which is a function of the difference between the daily rate that private residents pay and the average Medicaid reimbursement rate. Specifically, we calculate the average percentage difference between the two prices on the facility-year level using information on private and Medicaid rates from Pennsylvania and California (see Section 4 for details). Then we designate the quarter of nursing homes with the highest relative differences in daily rates as the high-incentive group, and the bottom quarter constitutes the low-incentive group. We expect that the relationship between occupancy and discharge rates of Medicaid residents is more pronounced in the high incentive group. The differences between private and Medicaid rates in the high-incentive group are larger than 28%, and they are below 6% in the low-incentive group. Since nursing homes set their private rates freely, there is a possible concern that they may be correlated with unobserved resident characteristics, leading to an endogenous sample split. However, it is unlikely that SNFs frequently change their prices due to changes in occupancy or unobservables, and we control for facility-year effects.

Using this sample split, Figure 15 displays estimated discharges rates by payer type and occupancy for high-incentive nursing homes on the left and low-incentive facilities on the right. Compared to the results for the whole sample in Figure 3, we find a steeper relationship between occupancy and discharge rates of Medicaid residents for the high-incentive group (left plots in Figure 15) and a flatter relationship in the low-incentive group (right plots). The baseline discharge rates below occupancy levels of 80% are similar in the two subsamples. Hence, we find clear evidence that SNFs who have more to gain from discharging Medicaid residents are more likely to do so at high occupancy rates.

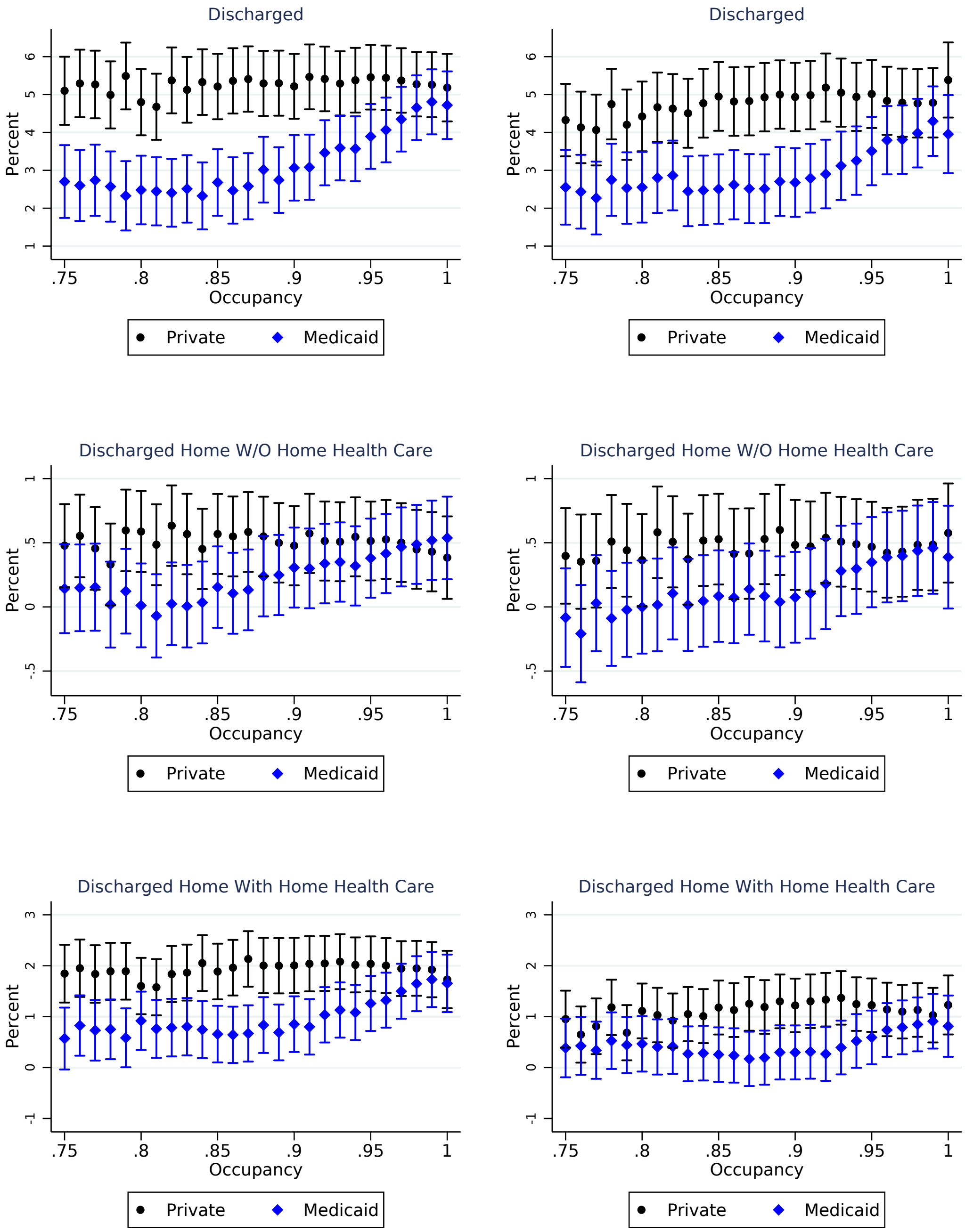
## D.10 Additional Health Benefits From Longer Nursing Home Stays

In addition to mortality and hospitalizations, we consider six health outcomes from residents' last health assessments before discharge: the CMI, ADL limitations, and indicators for clinical complexity, depression, impaired cognition, and behavioral problems.[28] For all health outcomes, a higher value implies that the resident is less healthy. If longer nursing home stays of Medicaid residents improve their health, it should be reflected in these measures.[29] In particular, we would expect to see a level difference between private and Medicaid residents with the latter being healthier. Moreover, we would expect that the health outcomes of Medicaid residents start to get relatively worth above 90% occupancy since their discharge rates increase above that threshold.

Overall, Figure 16 does not provide any evidence for health benefits of longer nursing home stays. First, Medicaid residents have worse health outcomes for all but one measure (behavioral problems) across all occupancy rates than private residents. Note that this

[28]In Figure 10 we use the same outcomes to analyze health-related selection. However, in those regression, we use current health outcomes whereas we use health at or close to discharge in the present analysis.

[29]Health improvements among SNF residents are relative. We do not expect to see absolute improvements in health status, but we rather interpret the absence of worsening health status as evidence for "health benefits" of longer nursing home stays.




***Notes:*** *See notes for Figure 3. This figure compares overall discharges (first row), home discharges without home health care (middle row), home discharges with home health care (bottom row) between resident groups. In the left figures, we focus on SNFs with high incentives (large differences between private and Medicaid rates). In the right graphs we present the evidence for SNFs with low incentives (small differences between private and Medicaid rates). The vertical bars indicate 90% confidence intervals.*

Figure 15: Discharge Rates by Nursing Home Incentives (California and Pennsylvania) – High Incentives SNFs on the Left and Low Incentive SNFs on the Right

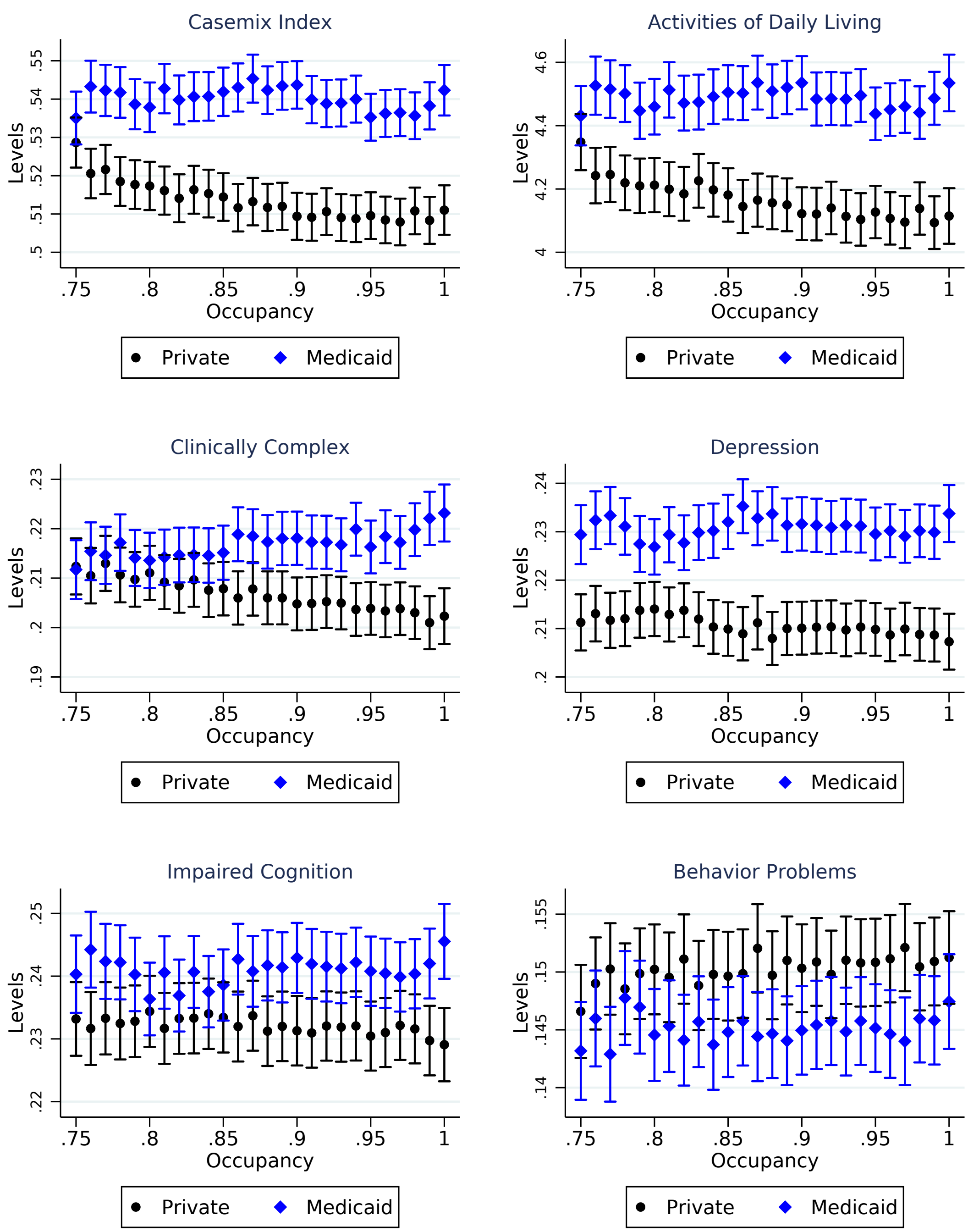
finding is in contrast to the results in Figure 10. That is, Medicaid residents have worse health outcomes at discharge although they are positively selected on the same measures during their stay. Second, we do not find any changes in the relationship between occupancy and health outcomes of Medicaid beneficiaries around 90% occupancy that would correspond to the kink in the discharge rate observed in Figure 3. Hence, we can reject that longer SNF stays leads to better health outcomes based on the predictions above. While there are some systematic differences between payer types—most notably the divergence in clinical complexity rates at higher occupancies—there does not seem to be a connection between the observed change in discharge rates and these health outcomes. Overall, we find no evidence suggesting that the longer nursing home stays at lower occupancies lead to any health benefits among Medicaid residents. This provides further evidence that extended stays among Medicaid beneficiaries contribute to LTC overspending.

# E Endogenous Occupancy Rates

In the counterfactual analysis, we take endogenous changes in occupancy rates into account, which in turn affect provider discharge efforts. To this end, we divide the nursing home into two wings. The additional (external) wing allows us to incorporate admissions and discharges among residents that were excluded from the estimation sample but also affect overall occupancy. These include the 22% of nursing home stays, who were initially covered by Medicare, and an additional 22% of stays that were initially covered by Medicaid. We treat these admissions and discharges as exogenous to the counterfactual policy changes. For the study population (nursing home wing) of interest, we take observed weekly admissions as exogenous, and use our structural model to predict discharge rates under alternative policy regimes.

We calibrate admissions and discharges in the external wing to match observed changes in occupancy rates conditional on observed admissions and the estimated discharge strategies in the focal wing of interest. Specifically, we consider a nursing home of $b$ beds and simulate occupancy changes in the focal wing of interest. To this end, we draw a sequence of shocks, $\epsilon^s = \left\{\epsilon^s_{occ}, \epsilon^s_{arr}, \epsilon^s_{\phi}, \epsilon^s_{\rho}, \epsilon^s_{dis}\right\}$ for each simulation iteration $s \in 1, ..., S$. The first shock $\epsilon^s_{occ}$ determines the change in occupancy rate for the entire nursing home. In combination with the occupancy transition matrix $\Theta(oc, oc')$, this shock specifies the occupancy for the next simulation draw (or next week) $oc^{s+1}$ conditional on today's occupancy rate, $oc^s$.

The remaining shocks govern admissions, payer type changes, and discharges in the focal wing of interest. $\epsilon^s_{arr}$, in conjunction with the arrival process outlined in Figure 2c, determines the number of new arrivals. $\epsilon^s_{\phi}$ and $\epsilon^s_{\rho}$ specify, in combination with $\phi$ and $\rho$ in Table 2, the payer type composition of new and previously admitted residents. Finally, $\epsilon^s_{dis}$, in




***Notes:*** *See notes for Figure 3. This figure compares the health profile at the time of discharge between payer types and occupancy rates. The resident's health status is decreasing in each health measure. The CMI is a summary measure of long term care needs, calculated based on methodology 5.01, and normalized to 1. The remaining health measures are direct inputs to the CMI formula and provide more granular information on cognitive and physical disabilities. The vertical bars indicate 90% confidence intervals.*

Figure 16: Health Profiles at Time of Discharge by Payer Type and Occupancy

combination discharge probabilities by occupancy rate and payer type, see Figure 5, specify the number of discharged residents.

Finally, we calibrate net changes in the number of residents in the external wing to match the overall change in the occupancy rate as a result of shock $\epsilon^s_{occ}$. For instance, suppose we started out with 90 occupied beds at time $s$ in the entire nursing home and that $\epsilon^s_{occ}$ implied a net increase to 92 occupied beds by $s+1$. Furthermore, suppose that the remaining shocks implied that the number of occupied bed in the focal wing of interest decreased from 38 to 37. Then we would assume a net increase of $\Delta^s_{ext} = 3$ seniors in the external wing to reconcile to overall increase from 90 to 92. This procedure generates a sequence of resident changes in the external wing $\{\Delta^s_{ext}\}$ for $s \in 1, ..., S$.

In the counterfactual analysis, we hold the sequence of shocks to the focal wing and resident changes in the external wing, $\epsilon^s = \left\{\epsilon^s_{arr}, \epsilon^s_{\phi}, \epsilon^s_{\rho}, \epsilon^s_{dis}, \Delta^s_{ext}\right\}$ for $s \in 1, ..., S$, fixed. Importantly, we can now ignore the sequence of occupancy shocks, $\epsilon^s_{occ}$. Of course, absent any policy changes, we are able to replicate the overall occupancy rate changes by inverting the strategy discussed in the previous paragraph that identified the sequence $\Delta^s_{ext}$. In the counterfactual analysis, we document changes in the discharge policies, see Figure 5, which we use to simulated a new sequence of overall occupancy rates. We summarize the mean occupancy rate over the simulation draws in the third row of Table 3.