

# The Expansionary and Contractionary Supply-Side Effects of Health Insurance\*

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## Abstract

We examine how health insurance expansions affect the entry and location decisions of health care providers. We first show theoretically that expansions of health insurance with relatively high, market-based prices (e.g., private insurance) should lead the supply side to expand, whereas expansions of insurance with low, administered prices (e.g., Medicaid coverage) should instead lead the supply side to contract. We test these predictions in the setting of retail clinics, where we exploit county-level changes in insurance coverage following the Affordable Care Act and 1,721 retail clinic entries and exits. Using two-way fixed effects and instrumental variable designs, we find that growth in private insurance leads to large growth in clinic entry, whereas clinic penetration is dampened by increases in Medicaid coverage. Further consistent with our theoretical model, we find that the contrasting supply-side effects of private insurance and Medicaid expansions are concentrated in states with low provider reimbursements under Medicaid. While it has long been accepted that reductions in the prices paid by consumers following insurance expansions should lead the supply side to expand to meet increased demand (Arrow, 1963), our results demonstrate that whether health insurance expansions cause the supply side to expand or contract further depends on how the prices received by providers are affected.

JEL: H44, I11, I13, I14, L11

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# I Introduction

Insurance expansions increase demand for health care by decreasing the prices paid by consumers (Manning et al., 1987). This increase in demand is in turn anticipated to cause supply-side responses that increase the market-level supply of health care resources (Arrow, 1963). Recent empirical work has confirmed these positive general equilibrium effects, showing that firm entry, technology adoption, and labor supply in the health care sector increase in response to sizable insurance expansions (Finkelstein, 2007; Kondo and Shigeoka, 2013; Hackmann et al., 2021). However, all health insurance is not created equal, and expansions of insurance with generous patient cost-sharing but low reimbursement rates for providers might generate the anticipated demand-side, but not supply-side, responses. As an increasing number of countries adopt two-tiered health care systems with both private and public insurance coverage (Einav and Finkelstein, 2023), an understanding of how relative insurance generosity affects supply-side responses to health insurance provision is key.

In this paper, we examine how the supply side responds to expansions of different types of health insurance coverage that differ in their provider reimbursement generosity. We exploit variation in health insurance coverage following the largest health insurance expansion in the United States in decades and focus on the entry and location decisions of on-demand health care clinics. On-demand clinics such as retail clinics and urgent care centers have been key contributors to growth in health care markets in recent decades and serve almost one third of the adult population (NCHS, 2022).<sup>1</sup> Combining data from 2010 to 2016 on the share of the population with health insurance of different types from the one-year American Community Survey (ACS) with information on the entry and exit decisions of the universe of retail clinics from Merchant Medicine, we find that growth in private insurance coverage leads to significant increases in clinic concentration, whereas growth in Medicaid coverage

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<sup>1</sup>On-demand health care clinics compete with traditional health care providers by offering convenience, and, in the case of retail clinics, lower and more transparent pricing. Retail clinics are located in retail outlets, are staffed by nurse practitioners or physician assistants, and treat a limited range of low-acuity conditions and provide preventative care. Urgent care centers treat more severe conditions, are typically staffed by physicians in addition to nurse practitioners and physician assistants, and often have imaging equipment available on site. According to data from the National Health Interview Survey, nearly 30 percent of adults reported going to a retail clinic or an urgent care center in 2019 (NCHS, 2022). This figure is significantly higher than the 22 percent of adults that reported going to a hospital emergency room.

significantly diminishes clinic penetration. We document that urgent centers exhibit similar location patterns using data on the universe of urgent care centers in 2021 from the National Urgent Care Realty (NUCR) database. Our findings indicate that supply-side responses to sizable insurance expansions can depend critically on the type of insurance being expanded and can even cause the supply side to contract.

We begin by introducing a theoretical model outlining how changes in insurance provision should influence clinic entry and exit. The model is based on the framework of [Sloan et al. \(1978\)](#) and considers a firm that faces demand from patients in a market with administered prices (in our setting, Medicaid patients) and patients in a market with market-based pricing (patients with private or no insurance). As in [Garthwaite \(2012\)](#), we model Medicaid expansions as shifting a portion of the population from market-based pricing to the administered price. Private expansions, on the other hand, increase the willingness to pay among patients in the non-Medicaid market.

The model generates three sets of predictions about the effects of changes in insurance coverage on firm entry patterns. First, if the firm was not accepting Medicaid patients at baseline, increases in private coverage serve to increase profits, thereby inducing entry. This occurs because demand among patients in the non-Medicaid market becomes more inelastic, allowing the firm to charge higher prices. Second, growth in Medicaid coverage reduces profits and induces exit, again among firms that were not accepting Medicaid patients at baseline. This occurs because shifting patients from the non-Medicaid to the Medicaid market reduces the size of the population being served by the clinic. Finally, expansions of either private insurance or Medicaid coverage generally have no effects on the profits of clinics that were serving both markets at baseline. The model demonstrates that whether clinics accept Medicaid is closely linked to Medicaid payment rates, with clinics being more likely to accept Medicaid when the administered payment under Medicaid is higher. The opposing supply-side responses to growth in private insurance and Medicaid coverage should therefore be more pronounced when Medicaid rates are lower.

To examine the relationship between health insurance provision and market structure empirically, we begin by using a two-way fixed effects specification to examine how retail clinic concentration covaries with the share of the population with any health insurance

coverage at the county-year level. This analysis reveals a surprising result demonstrating that positive supply-side responses need not accompany sizable insurance expansions: despite significant variation in insurance growth and clinic penetration over our sample period, there was no association between within-county changes in health insurance coverage and clinic growth. While counties in the highest decile of insurance growth experienced average increases in health insurance coverage of nearly 12 percentage points compared to only 2 percentage points among counties in the lowest decile, counties in both groups saw an increase of approximately 0.15 retail clinics per 100,000 people. To examine whether clinic growth depends on the type of insurance being expanded, we then estimate two-way fixed effects specifications that exploit conditional variation in the share of the population with private insurance and Medicaid coverage within counties over time. This analysis reveals pronounced heterogeneity by insurance type, with growth in private insurance coverage associated with large growth in the concentration of clinics and growth in Medicaid coverage associated with reduced clinic penetration.

While informative, our two-way fixed effects specifications might be confounded by changes in local socio-demographics. Notably, income eligibility requirements for Medicaid and the provision of the majority of private insurance through employers ensures that health insurance in the United States is closely tied to income. As such, changes in the share of the population with private insurance or Medicaid coverage over our sample period will capture changes in insurance provision driven by policy as well as changing socio-demographics within locations. To isolate variation in health insurance driven by policy, we instrument for changes in insurance levels and types using four features of the Affordable Care Act (ACA).

First, following previous work, we exploit the fact that growth in Medicaid coverage was larger in counties that were in states that expanded Medicaid and that had a larger share of their population below 138 percent of the federal poverty level (FPL). Moreover, we introduce two novel instruments that affect private insurance coverage.<sup>2</sup> The first exploits baseline variation in the share of the population between 138 and 400 percent of the FPL

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<sup>2</sup>Conceptually, Medicaid expansions shift Medicaid coverage and ACA provisions surrounding direct purchase and employer-sponsored coverage shift private insurance coverage. However, our instruments will shift both types of insurance coverage simultaneously due to crowd-out, and thus we include all of our instruments when predicting changes in the local provision of each insurance type.

(i.e., the population eligible for subsidies on the exchanges) to shift direct purchase insurance. The second exploits baseline variation in the share of the population employed (i.e., the population targeted by the mandate for large employers to provide health insurance) to shift employer-sponsored coverage.<sup>3</sup> These instruments are powerful and help deal not only with endogeneity concerns but also measurement error in the available county-year level data on health insurance coverage that are both self-reported and collected from a 1 percent sample of households.

Results from our instrumental variables analysis confirm the patterns observed in the two-way fixed effects specifications. The results are large and show that growth in private insurance coverage of 5 percentage points—the average increase experienced by counties in our sample over our time period—leads to an increase of 0.16 retail clinics per 100,000 people, or about 25 percent relative to the mean. In contrast, growth in Medicaid coverage of 4 percentage points—the average increase experienced by counties in our sample over our time period—leads to a reduction of 0.20 retail clinics per 100,000 people, or over 30 percent relative to the mean. Additional analyses using first-difference specifications show that these impacts are driven by effects on both entries and exits, with private insurance growth leading to increased clinic entry and Medicaid growth leading to increased clinic exit.

We show that these opposing supply-side responses to insurance expansions are likely driven by low reimbursement rates under Medicaid.<sup>4</sup> As outlined above, theory predicts that the supply-side effects of private insurance and Medicaid expansions should be concentrated in areas in which it is not profitable to serve the Medicaid market (i.e., areas in which Medicaid reimbursement rates are low). Using data on state-level Medicaid reimbursement rates for office visits from [Alexander and Schnell \(2019\)](#), we show that the positive effects of growth in private insurance coverage on retail clinic penetration are most

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<sup>3</sup>We show that counties with different income profiles and employment rates were on similar trends in retail clinic growth before the ACA. We further show that our results are very similar if we use the baseline share of the population employed by firms with at least 50 employees when constructing the instrument.

<sup>4</sup>Alternatively, gaining Medicaid coverage might lead patients to shift toward more traditional sources of care, thereby reducing demand for retail clinics and promoting exit. We find little support for this alternative mechanism: the negative supply-side effects of Medicaid coverage are most pronounced in areas with the least health care resources at baseline (i.e., areas in which there is little scope for substitution), and our results are unaffected when we control for growth in federally funded, non-profit health care clinics that target underserved populations.

pronounced among counties in the bottom tercile of Medicaid reimbursements at baseline. In fact, there is no significant relationship between growth in private insurance coverage and clinic penetration among counties in states with the highest Medicaid payments. We observe similar heterogeneity in the relationship between county-level growth in Medicaid coverage and clinic concentration, with the negative supply-side effects of Medicaid growth concentrated in locations with low Medicaid reimbursement rates at baseline.

The supply-side responses that we document are likely inefficient from the perspective of the social planner. When we split the sample into areas that were and were not designated as primary care shortage areas by the Health Resources and Services Administration (HRSA) at baseline, we find that growth in clinics following private insurance expansions is predominately concentrated in areas with sufficient baseline resources. In contrast, reductions in clinic penetration following growth in Medicaid coverage are observed across both shortage and non-shortage areas. This suggests that insurance-induced changes in the concentration of clinics are unlikely to address existing access barriers and may exacerbate unnecessary service provision in well-resourced areas.

On-demand health care clinics are the ideal setting in which to examine the supply-side effects of health insurance. While seminal work by Finkelstein (2007) focused on the entry of hospitals in the late 1960s, entry of hospitals in recent decades has been rare, with fewer than 200 openings between 2010 and 2016.<sup>5</sup> In contrast, nearly 200 retail clinics opened in each year of our sample period, allowing for a careful statistical examination of how entry is driven by changes in local health insurance rates and composition. Moreover, retail clinics offer a limited set of services at relatively low prices (Thygeson et al., 2008; Mehrotra et al., 2009). This limits the scope for variation in insurance status and payment generosity to affect their location and entry decisions, suggesting that analogous effects on the entry of other types of health care providers, such as traditional doctors' offices, might be even more pronounced.<sup>6</sup>

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<sup>5</sup>The number of hospital entries is computed using annual surveys from the American Hospital Association (AHA). We identify entering hospitals from 2010 to 2016 as short-term, non-federal hospitals that were not present in AHA surveys in either 2006 or 2008 but responded to an AHA survey at some point between 2010 and 2016. To avoid misclassifying changes in hospital ownership as hospital entry, we exclude new hospital identifiers that are located in the same geographic location as a previous hospital identifier.

<sup>6</sup>We use our theoretical model to develop insights into the supply-side effects of health insurance in other care settings in Section III.B.

Our work contributes to three literatures. First, we build on the literature examining the supply-side effects of health insurance. Finkelstein (2007) found that the introduction of Medicare in 1965 led to hospital entry and increased adoption of new technologies. Outside of the United States, Kondo and Shigeoka (2013) demonstrated that the 1961 introduction of universal health insurance coverage in Japan led to increases in the number of hospital beds but had no conclusive effects on the number of medical institutions or medical labor supply. In Germany, a recent paper by Hackmann et al. (2021) documents that the introduction of universal, long-term care (LTC) insurance in 1995 led to sizable increases in the number of LTC firms and workers. We contribute to this work by examining the impacts of the largest insurance expansion in the United States in decades and show that firm-level entry responses—an anticipated mechanism through which supply keeps pace with growing demand—depend on the type of coverage being expanded. This result has implications beyond the United States; as an increasing number of countries expand two-tiered health insurance systems where providers can opt out of accepting less generous public insurance coverage, provider responses to relative insurer generosity will become an increasingly important driver of access to care.<sup>7</sup>

Our finding that the type of coverage being expanded is important for shaping supply-side responses relates to work documenting the importance of insurance generosity on the behavior of providers. Recent work shows that the reluctance of providers to accept Medicaid is driven by the program's low reimbursement rates for providers relative to other payers (Alexander and Schnell, 2019) and billing hassles that plague the Medicaid system (Dunn et al., 2021). Focusing on technology adoption, Freedman et al. (2015) showed that expansions of Medicaid eligibility for pregnant women in the 1980s and 1990s did not affect hospitals' adoption of neonatal intensive care units. In the pharmaceutical space, Garthwaite et al. (2021) document that research and development activities—which are typically linked to

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<sup>7</sup>Many countries have supplemental private health insurance programs, including Australia, Canada, Germany, Israel, Singapore, and the United Kingdom (Einav and Finkelstein, 2023). In many two-tiered health care systems, providers can choose to only offer their services in the private market. For instance, in the market for dental services in the United Kingdom, many dentists opt out of providing services through the National Health Service in favor of setting their own (higher) rates in the private market (HSCSC, 2023). One important implication of our theoretical model is that an expansion of private health insurance can result in a decrease in access to care in the public sector even when it induces additional entry into the provider market overall.

market size—did not respond to recent Medicaid expansions.<sup>8</sup> Both Freedman et al. (2015) and Garthwaite et al. (2021) attribute their null findings to Medicaid’s low reimbursement rates relative to other providers. We add to this work by showing that firm-level entry responses are likewise shaped by the generosity of insurance for providers, uncovering the contractionary supply-side effects of Medicaid expansions in a setting in which the expansionary effects of growth in private insurance coverage can be simultaneously confirmed.

Finally, our work contributes to recent discussions surrounding the impacts of new health care delivery mechanisms on access to and use of care. Much of the recent literature has focused on the effects of retail clinics and urgent care centers on emergency room (ER) use and aggregate health care costs. While Alexander et al. (2019) and Allen et al. (2021) show that retail clinics and urgent care centers reduce unnecessary ER use, respectively, Ashwood et al. (2016), Currie et al. (2021), and Wang et al. (2021) show that these clinics can nevertheless lead to increased costs by increasing total health care utilization. In contrast to prior work, we focus on the location decisions of such clinics and examine how their expansion patterns have been shaped by changing insurance landscapes. This focus relates to recent work by Magnolfi et al. (2022), who estimate an equilibrium model of market structure for urgent care centers and hospitals and find that hospital presence deters the entry of urgent care centers. While these location patterns could suggest that on-demand health care will help equalize access across the United States, our findings indicate that retail clinics and urgent care centers—by avoiding areas with growth in Medicaid beneficiaries and concentrating in areas with high rates of private insurance coverage and existing health care resources—are unlikely to meaningfully address access barriers faced by disadvantaged populations as many health policy experts have hoped (see, for example, Bechraich and Frohlich, 2016).<sup>9</sup>

This paper proceeds as follows. In Section II, we outline the data sets that we use.

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<sup>8</sup>Focusing on physician labor supply, Garthwaite (2012) found that physicians decreased their time spent with patients following the 1997 implementation of the State Children’s Health Insurance Program (SCHIP)—a program that expanded health insurance for low-income children who do not qualify for traditional Medicaid.

<sup>9</sup>A 2010 report by the RAND Corporation notes that “some champions have argued that retail clinics may improve access to care for populations in underserved areas” (Weinick et al., 2010). However, the report emphasizes that “the viability of retail clinics in underserved areas is uncertain and remains largely unexplored as a model for improving access to care in such areas.”

Section III introduces a theoretical framework that delivers predictions about the impacts of changes in insurance provision on clinic entry and exit. Section IV presents our empirical strategies and discusses identification. Our main results are presented in Section V, and extensions are presented in Section VI. Section VII provides a discussion and concludes.

## II Data

We use data from two main sources. The locations and operating dates of all retail clinics in the United States from 2010 to 2016 come from Merchant Medicine. Data on the county-level shares of the population with different types of health insurance coverage and local socio-demographics in each year over the same period come from the ACS. We describe each of these primary data sources in more detail below. Additional data sources—including those used to construct the instruments and to examine heterogeneity by Medicaid reimbursement rates—are introduced in subsequent sections.

### II.A On-demand health care clinics

Information on retail clinics comes from Merchant Medicine, a management consulting firm serving the on-demand health care market.<sup>10</sup> These data are comprehensive and contain the geo-coded locations and operating dates of all retail clinics ever operating in the United States. Using this data set, we create a panel of the total number of operating clinics and the number of entries and exits at the county-quarter level. As shown in Figure 1(a), the number of retail clinics grew steadily over our sample period, with the number of clinics nationally increasing by 66 percent from 1,224 at the beginning of 2010 to 2,036 by the end of 2016. Moreover, there was substantial churn in the market, with an average of over 50 entries and nearly 22 exits in each quarter from 2010 to 2016 (see Figure A1).

We supplement these data with information on the locations of all urgent care centers operating in the United States in 2021 from the NUCR database. The NUCR data are less comprehensive than the Merchant Medicine data and only include information on the year of

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<sup>10</sup>This proprietary data is available for purchase to qualified researchers. For more information, contact UCP Merchant Medicine here: <https://www.ucpmm.com/contact-us>.

entry and geo-coded locations for clinics that remained open in 2021. Given this limitation, we focus much of our analysis on the entry and exit behavior of retail clinics, but we use the NUCR data to examine whether the location patterns of urgent care centers and retail clinics exhibit a similar relationship with health insurance provision in the cross-section. As shown in Figure A2(a), there were nearly 13,000 urgent care centers operating in 2021, only 25 percent of which were open in 2010.

## II.B Insurance shares and other county-level characteristics

We combine these data on on-demand health care clinic locations with information on county-level characteristics from the ACS. Our main independent variables of interest are the shares of the population with any health insurance coverage and health insurance coverage of different types at the county-year level. We focus on private insurance (employer-sponsored and direct purchase) and Medicaid coverage, although we control for Medicare coverage and all other types of health insurance in our analyses.<sup>11</sup>

To capture annual changes in these variables, we use data from the one-year ACS in our primary analyses. Because only counties with at least 65,000 residents are included in the single-year files, we restrict these analyses to the 555 counties that were in every one-year ACS from 2010 to 2016; these counties account for over 75 percent of the total U.S. population and nearly 87 percent of all operating retail clinics in 2016 (see Table A1 and Figure 1(b)). When considering location patterns of retail clinics and urgent care centers in the cross-section, we instead use information on health insurance coverage for all U.S. counties from the five-year pooled ACS.<sup>12</sup>

As shown in Figure 2(a), the share of the population with health insurance coverage across the United States increased from less than 85 percent in 2010 to nearly 92 percent in 2016. This increase was driven almost entirely by the implementation of the ACA in 2014. Moreover, as shown in Figure 2(b), the increase in health insurance coverage nationally was

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<sup>11</sup>When constructing insurance shares, we hold population fixed as of 2010 to ensure that our analyses capture changes in the number of patients with health insurance of a given type available to clinics rather than changes in population. All of our analyses include a time-varying control for population to account for population growth.

<sup>12</sup>More precisely, we compare location patterns of retail clinics in 2016 (urgent care centers in 2021) to local characteristics from the 2012–2016 (2016–2020) ACS.

driven by sizable increases in both private insurance and Medicaid coverage. At the county level, there is variation in the levels and types of coverage expanding between 2013 and 2015, with some areas seeing large growth in Medicaid coverage only, private insurance coverage only, or both (Figure 2(c)).

To control for other differences across counties that could influence clinic penetration, we further consider a range of socio-demographics from the ACS. As shown in Table 1, retail clinics in 2016 were more likely to be located in dense, urban hubs. This can further be seen in Figure 1(b), which shows that the geographic distribution of retail clinics across the United States in 2016 largely mirrored the distribution of the population. Reflecting the demographics of urban areas, retail clinics in 2016 were more likely to be located in areas with a more diverse, educated population and with higher median incomes and rates of employment. This is true both when considering only counties in the one-year ACS (Table 1) and all U.S. counties (Table A1). Moreover, while growth in urgent care centers had spread beyond the largest metropolitan areas by 2021, Table A2 shows that such centers were likewise concentrated in counties that were relatively economically prosperous and racially diverse, with nearly every county with a retail clinic in 2016 seeing an urgent care center operating in 2021 (see Figure A2(b)).

### III Theoretical framework

In this section, we introduce a theoretical model of entry and exit for on-demand health care clinics (hereafter referred to as “clinics”). In our setting, clinics decide whether to enter the market and, if so, which price to charge. We assume below that clinics charge the same price to all patients; we consider an extension in which firms charge different prices to patients who are and are not covered by Medicaid in Appendix C. We aim to examine how expansions of private insurance and Medicaid coverage affect firm profits and in turn, firm-level entry and exit decisions. The model delivers a number of theoretical insights that both rationalize our main findings and motivate additional empirical exercises.

We follow the general setup of Sloan et al. (1978), who introduce a mixed-economy model formalizing providers’ decisions surrounding optimal participation in government insurance

programs.<sup>13</sup> Our setting differs from this previous work in two key ways. First, in contrast to settings in which providers can ration appointment availability based on insurance coverage and type, on-demand health care clinics serve patients on a first come, first served basis. Clinics in our setting are therefore faced with a binary decision of whether to accept Medicaid patients—rather than a decision of how many Medicaid patients to accept—and face the potential of seeing any number of patients covered by public insurance once they opt to serve the program’s beneficiaries. Moreover, while previous work has focused predominately on the decision of whether to accept a given type of insurance coverage conditional on entry, our primary goal is to examine how changes in the share of the population covered by different types of health insurance affect firm profitability and equilibrium market structure.

### III.A Baseline model

Clinics face demand from consumers in two markets: (1) a market with administered prices in which they serve the  $s_M$  share of the population covered by Medicaid, and (2) a market with market-based prices in which they serve the  $1 - s_M$  share of the population that is privately insured or uninsured (“non-Medicaid patients”). Let  $p_M$  denote the price that clinics receive when treating patients covered by Medicaid; this price reflects the administered price net of any hassle costs associated with program billing.

Total demand facing the clinic is shown in Figure 3(a). Because Medicaid patients do not incur any out-of-pocket costs and cannot opt for self-pay, the demand curve is perfectly elastic at  $p_M$  with length  $s_M$ .<sup>14</sup> At all other prices, the firm faces downward-sloping demand from the non-Medicaid population. As shown in Figure 3(a), the resulting kinks in the total demand curve lead to discontinuities in the associated marginal revenue curve. In particular, while marginal revenue is downward sloping and lies below the demand curve when demand is downward sloping, the marginal revenue and demand curves overlap on the perfectly elastic portion of the demand curve.<sup>15</sup> This reflects the fact that clinics do not need to lower the

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<sup>13</sup>More recently, Garthwaite (2012) adapted the model introduced by Sloan et al. (1978) to outline predictions of the introduction of SCHIP on physicians’ program participation and labor supply.

<sup>14</sup>Providers are generally not allowed to accept payment from known Medicaid patients, thereby preventing the program’s beneficiaries from opting for self-pay (ASHA, 2023).

<sup>15</sup>In Figure 3(a), marginal revenue is negative when  $q > D(p_M)$ ; this will typically be the case unless very few patients are covered by Medicaid (i.e., the elastic portion of the total demand curve is very short) or the

price to attract an additional Medicaid patient when  $p = p_M$ , thereby keeping revenue from both marginal and inframarginal patients constant.

Figure 3(a) shows the profit-maximizing prices and quantities set by clinics faced with different marginal cost curves. When there is a single intersection between marginal revenue and a given marginal cost curve, this intersection determines the quantity of patients that the clinic serves (i.e.,  $q^*$  is such that  $MR(q^*) = MC(q^*)$ ). To achieve this optimal quantity, the firm sets  $p^* = D^{-1}(q^*)$ . Whether the firm accepts Medicaid depends on how  $p^*$  compares to  $p_M$ : if  $p^* > p_M$ , the clinic does not serve Medicaid patients, whereas the clinic accepts patients covered by Medicaid if  $p^* \leq p_M$ . As  $p_M$  increases—that is, as the program becomes more generous for providers—it becomes more likely that the firm will accept Medicaid. Note that when the clinic serves both market segments, the share of the clinic's  $q^*$  patients that are covered by Medicaid is indeterminant and depends on patient arrival patterns.

Figure 3(a) shows two examples of marginal cost curves with single marginal revenue intersections. When marginal costs are given by  $MC_1$ , the clinic sets  $p = p_1^*$ . Since  $p_1^* > p_M$ , the clinic does not serve patients covered by Medicaid and instead serves  $q_1^*$  patients coming from the non-Medicaid market. In contrast, when marginal costs are given by  $MC_2$ , the single intersection between marginal revenue and marginal costs occurs on the perfectly elastic portion of the demand curve. The firm sets  $p = p_2^* = p_M$ , thereby accepting patients covered by Medicaid, and serves  $q_2^*$  patients coming from both the Medicaid and non-Medicaid markets.

Given the discontinuities in the marginal revenue curve, there need not be a single intersection between marginal revenue and marginal costs. In particular, there can be two intersections between the discontinuous marginal revenue curve and a given marginal cost curve near the first jump in marginal revenue and no intersections near the second jump. Figure A4 shows how prices and quantities are determined in each of these cases. When there are two intersections between marginal revenue and marginal costs (see Figure A4(a)), the firm must further consider average total costs to compare profits at each potential set of prices and quantities. When the marginal cost curve instead lies entirely between the different portions of the marginal revenue curve (as in Figure A4(b)), the firm sets  $p^* = p_M$  and

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Medicaid price is relatively high (i.e., few non-Medicaid patients have willingness to pay greater than  $p_M$ ).

sees all patients willing to pay at least  $p_M$  (i.e., the firm does not need to restrict capacity).

### III.B Insurance expansions and firm entry decisions

We aim to examine how health insurance expansions affect firm profits and in turn, firm entry and exit decisions. We begin by demonstrating how expansions of Medicaid coverage and private insurance affect the demand facing clinics. As shown in Figure 3(b), Medicaid expansions increase the share of the population covered by Medicaid, thereby lengthening the perfectly elastic component of the total demand curve. In contrast, expansions of private insurance coverage increase the willingness to pay among non-Medicaid patients by reducing out-of-pocket costs for some segments of this population. As shown in Figure 3(b), this serves to rotate the demand curve among the non-Medicaid population upward.

How do such insurance expansions affect firm entry? Firms will enter (exit) the market when the average total cost is below (above) the profit-maximizing price. We first consider the effects of a private insurance expansion. As shown in Figure 4(a), an upward rotation of the demand curve among non-Medicaid patients leads clinics that did not accept Medicaid at baseline to increase quantity and prices (i.e.,  $q_2^* > q_1^*$  and  $p_2^* > p_1^*$ ). Profits increase, inducing additional entry into the market. In contrast, it is possible for expansions of private insurance to have no effects on firm profits when the firm accepted Medicaid at baseline.<sup>16</sup> As shown in Figure 4(b), optimal quantity and price do not change (i.e.,  $q_2^* = q_1^* = p_M$  and  $p_2^* = p_1^* = p_M$ ) when the marginal cost curve intersects the marginal revenue curve on a portion that is unaffected by the private expansion. Profits stay the same, and the expansion has no effects on firm entry.

Now consider the effects of a Medicaid expansion. As shown in Figure 4(c), an inward shift in the demand curve among non-Medicaid patients leads clinics that did not accept Medicaid at baseline to decrease quantity and prices (i.e.,  $q_2^* < q_1^*$  and  $p_2^* < p_1^*$ ). Profits

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<sup>16</sup>There are two cases in which a private insurance expansion will cause profits to increase among firms that accepted Medicaid at baseline. First, suppose that marginal costs and marginal revenue intersect along the perfectly elastic portion of the demand curve but close to the first kink in total demand at baseline. If the upward rotation in demand among the non-Medicaid market is sufficient to cause the firm to stop accepting Medicaid (i.e.,  $p_2^* > p_1^* = p_M$ ), then profits will increase. Moreover, suppose that the marginal cost curve lies between the portions of the marginal revenue curve at baseline. As shown in Figure A5(a), while optimal pricing is unaffected (i.e.,  $p_2^* = p_1^* = p_M$ ), optimal quantity—and firm profits—will increase.

decrease, and firms exit the market. However, as was the case with private expansions, it is possible for Medicaid expansions to have no effects on firm profits when the firm accepted Medicaid at baseline.<sup>17</sup> As shown in Figure 4(d), profits stay the same and the expansion does not lead to exit when the marginal revenue curve is unaffected by the expansion near its intersection with marginal costs.

Our baseline model therefore delivers two key sets of predictions regarding the supply-side effects of insurance expansions. First, expansions of private insurance should generally serve to increase firm profitability, thereby leading to additional entry. In contrast, expansions of Medicaid should tend to reduce firm profits and lead to clinic exit. These opposing supply-side effects of insurance expansions based on the type of coverage being expanded are the focus of Sections IV and V.

Second, the model predicts that the supply-side effects of insurance expansions should vary depending on whether clinics were accepting Medicaid at baseline (or would have accepted Medicaid had they entered). In particular, the positive (negative) supply-side effects of private insurance (Medicaid) coverage should be concentrated in areas that were not accepting Medicaid at baseline. Because the level of Medicaid reimbursement rates dictates the vertical positioning of the perfectly elastic component of demand, with clinics being more likely to accept Medicaid in areas with higher reimbursement rates under the program, it follows that the supply-side responses to growth in private insurance and Medicaid coverage should be more pronounced in areas with low Medicaid reimbursement rates. This prediction motivates our analysis of the heterogeneous supply-side effects of insurance expansions by baseline Medicaid reimbursement rates in Section VI.

We have thus far assumed that clinics charge the same price to all payers. We relax this assumption in Appendix C and allow firms to charge different prices to Medicaid and non-Medicaid patients. This two-price extension differs from the one-price case outlined above in that firms care about the composition of patients by payer type rather than only the total number of patients served. Nevertheless, the two-price setting delivers predictions for the

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<sup>17</sup>There is one case in which a Medicaid expansion will cause profits to increase among firms that accepted Medicaid at baseline. Suppose that the marginal cost curve lies between the different portions of the marginal revenue curve at baseline (i.e., there is no intersection). As shown in Figure A5(b), while optimal pricing will be unaffected by an expansion of Medicaid coverage in this case (i.e.,  $p_2^* = p_1^* = p_M$ ), optimal quantity increases (i.e.,  $q_2^* > q_1^*$ ). Profits likewise increase, leading to firm entry.

supply-side effects of insurance expansions that mirror those derived above. In particular, we show in Appendix C.2 that expansions of private insurance should lead firm profits to increase whereas expansions of Medicaid should generally lead firm profits to decline. Moreover, as in the one-price case, we show that areas with low Medicaid reimbursement rates should experience the most pronounced supply-side responses to expansions of both private insurance and Medicaid coverage.<sup>18</sup>

We can further use the model to develop insights into the supply-side effects of health insurance in other care settings. Recall that while on-demand health care clinics can choose whether to accept Medicaid, patients are treated on a first come, first served basis, and thus on-demand clinics cannot control payer mix conditional on accepting multiple insurance types. In contrast, traditional doctors' offices can both choose whether to accept Medicaid and ration appointment availability to directly control the number of patients that they see with each type of insurance that they accept.<sup>19</sup> All of the theoretical results hold when providers can “pick” their patients, and thus we anticipate that doctors' offices and on-demand health care clinics will respond similarly to insurance expansions.<sup>20</sup> However, since emergency departments are required to stabilize patients regardless of their ability to pay for care, the predictions likely differ for hospitals. In particular, since Medicaid expansions reduce the amount of uncompensated care that hospitals provide (Dranove et al., 2016; Garthwaite et al., 2018), such expansions could instead lead to expansionary responses in hospital settings.

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<sup>18</sup>A key difference from the one-price case is that firms that accepted Medicaid at baseline will also be affected by insurance expansions when we allow for two prices. However, the supply-side effects of insurance expansions still depend on whether clinics found it profitable to accept Medicaid at baseline.

<sup>19</sup>Because on-demand health care clinics and traditional doctors' offices have the ability to not accept Medicaid if seeing Medicaid patients would lead the clinic to incur losses, there are no incentives for such clinics to engage in cost-shifting behavior.

<sup>20</sup>Recall that the magnitudes of the supply-side responses to insurance expansions of different types are determined in part by the relative difference between the market-based and the administered price. As this price differential is likely larger in office-based and urgent care settings than in retail clinics (due to low, posted prices in retail clinics; see Thygeson et al., 2008 and Mehrotra et al., 2009), the responses to insurance expansions might be even more pronounced in these other settings.

## IV Empirical strategies

We now turn to an empirical investigation of the supply-side effects of health insurance expansions. This section introduces our primary empirical methods; results and extensions are presented in Sections V and VI, respectively.

### IV.A Two-way fixed effects

We begin with simple two-way fixed effects designs. Let  $Clinics_{ct}$  denote the number of open retail clinics per 100,000 people in county  $c$  in year  $t$ . To examine how county-level changes in the share of the population with health insurance covary with county-level changes in the number of retail clinics, we estimate the following specification:

$$Clinics_{ct} = \beta \cdot Insured_{ct} + \delta \cdot X_{ct} + \gamma_c + \gamma_t + \epsilon_{ct}, \quad (1)$$

where  $Insured_{ct}$  is the share of the population in county  $c$  in year  $t$  with health insurance coverage of any type;  $X_{ct}$  are time-varying, county-level controls listed in Table 1; and  $\gamma_c$  and  $\gamma_t$  are county and year fixed effects, respectively. Observations are weighted by county population in 2010, and standard errors are clustered by county.

Estimation of equation (1) shows how the concentration of retail clinics within a county covaries with the share of the population covered by any type of health insurance over time. To examine whether retail clinics respond differently to changes in the share of the population covered by different types of health insurance, we estimate the following specification:

$$Clinics_{ct} = \beta_1 \cdot Private_{ct} + \beta_2 \cdot Medicaid_{ct} + \delta \cdot X_{ct} + \gamma_c + \gamma_t + \epsilon_{ct}, \quad (2)$$

where  $Private_{ct}$  and  $Medicaid_{ct}$  are the shares of the population with private insurance or Medicaid coverage in county  $c$  in year  $t$ , respectively, and all other variables are defined as in equation (1).<sup>21</sup> Observations are again weighted by county population in 2010, and standard

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<sup>21</sup>When considering effects by insurance type,  $X_{ct}$  further includes controls for the share of the population with Medicare and the share of the population with other types of health insurance. With these additional controls,  $\beta_1$  and  $\beta_2$  reflect the impacts of increases in private insurance and Medicaid coverage, respectively, relative to the uninsured. Since changes in the share of the population covered by health insurance other

errors are clustered by county.

While informative, results from estimation of equations (1) and (2) may not provide evidence of the causal effects of changes in health insurance on the concentration of clinics. Notably, health insurance in the United States is closely tied to income: while income eligibility requirements ensure a close connection between the share of the population with Medicaid coverage and the share living in poverty, the correlation between income and employment also leads the share of the population with private insurance—which is largely provided through employers—to be closely linked to the share of the population with incomes well above the FPL. While equations (1) and (2) control for county-level median income, the share of the population living in poverty, and the employment rate, it could nevertheless be the case that any supposed preference for the privately insured (and distaste for those covered by Medicaid) is simply capturing retail clinics’ preference for the wealthy (and distaste for the poor) rather than heterogeneity in the supply-side effects of health insurance coverage by insurance type.

## IV.B Instrumental variables

To isolate the impacts of different types of insurance coverage separately from other county-level characteristics that are correlated with health insurance composition and might directly influence the entry and exit decisions of retail clinics, we instrument for changes in insurance coverage driven by changes in policy. In particular, we leverage four features of the ACA that drove differential changes in the share of the population covered by private insurance and the share of the population covered by Medicaid in each county.

The first two features of the ACA that we exploit directly affected the share of the population covered by Medicaid and have been commonly used in previous work. Most notably, 20 states expanded their Medicaid programs to extend eligibility to low-income, childless adults in 2014, and five states made similar changes between 2010 and 2013 (see Figure A6). As shown in Figure A7(a), state-level variation in the decision of whether to expand Medicaid was a key driver of changes in Medicaid enrollment from 2013 to 2015.

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than private or Medicaid were minimal over our sample period, results are nearly identical if we exclude these additional controls.

Second, nearly all Medicaid expansions expanded coverage to individuals with incomes up to 138 percent of the FPL. We therefore additionally exploit county-level variation in exposure to Medicaid expansions driven by variation in the share of the population in 2013 that was uninsured, between ages 18 and 64, and making less than 138 percent of the FPL as reported in the Small Area Health Insurance Estimates (SAHIE). As shown in Figure A7(b), counties that were in states that expanded Medicaid by 2014 and had an above-median share of the population at baseline who stood to gain insurance coverage under a Medicaid expansion saw among the largest increases in the shares of their populations with Medicaid coverage from 2013 to 2015.

We exploit two additional features of the ACA that directly affected the share of the population with private insurance coverage. To the best of our knowledge, these instruments are novel and might be useful to researchers in other contexts. First, the ACA mandated that employers with 50 or more full-time employees provide health insurance coverage. Since over 70 percent of jobs are in companies with at least 50 employees (QWI, 2013), this provision led to increases in private insurance that were closely tied to a county's employment rate at baseline (see Figure A7(c)). As such, in our primary specifications we use variation in private insurance growth that was driven by variation in the share of the population that was employed in 2013. Moreover, using data from the Quarterly Census of Employment and Wages (QCEW), we show that our results are robust to using proxies for the baseline shares of the population that were employed by firms that were most affected by the employer mandate.

Finally, starting in 2014, the ACA directed the federal government to begin providing subsidies for individuals with incomes between 138 and 400 percent of the FPL to purchase insurance through the newly designed exchanges (“direct purchase”). Again using data from the SAHIE, Figure A7(d) shows that this provision led to increases in private insurance coverage that were more pronounced in counties with higher shares of their populations who were uninsured, between ages 18 and 64, and had incomes between 138 and 400 percent of the FPL in 2013. We therefore additionally leverage variation in private insurance growth that was driven by variation in the baseline share of the population that stood to benefit from the ACA’s direct purchase subsidies.

We use these instruments to examine the causal effects of changes in health insurance coverage on the concentration of retail clinics. To isolate variation in insurance stemming from these policy changes, we estimate first stages of the form:

$$\begin{aligned}
\{Insured_{ct}, Private_{ct}, Medicaid_{ct}\} = \\
\alpha_1 \cdot Post_t \cdot Employed_c^{2013} + \alpha_2 \cdot Post_t \cdot [138 - 400\% FPL]_c^{2013} \\
+ \alpha_3 \cdot Post_t \cdot [< 138\% FPL]_c^{2013} \cdot Expansion_s \\
+ \alpha_4 \cdot Post_t \cdot [< 138\% FPL]_c^{2013} + \alpha_5 \cdot Post_t \cdot Expansion_s \\
+ \delta \cdot X_{ct} + \gamma_c + \gamma_t + \epsilon_{ct},
\end{aligned} \tag{3}$$

where  $Post_t$  is an indicator denoting years 2014 and onward;  $Expansion_s$  is an indicator denoting whether state  $s$  expanded Medicaid by 2014;  $Employed_c^{2013}$  denotes the share of the population that was employed in 2013 in county  $c$ ;  $[138 - 400\% FPL]_c^{2013}$  and  $[138 - 400\% FPL]_c^{2013}$  denote the share of the population that was uninsured, between the ages of 18 and 64, and either below 138 percent of the FPL or between 138 and 400 percent of the FPL, respectively, in county  $c$  in 2013; and all other variables are defined as in equation (1). Throughout this section, observations are again weighted by county population in 2010, and standard errors are clustered by county.

We estimate equation (3) separately using the county-year share of the population with any health insurance coverage, private insurance coverage, or Medicaid coverage as the dependent variable. Conceptually, the first instrument ( $Post_t \cdot Employed_c^{2013}$ ) shifts individuals into employer-sponsored health insurance, the second instrument ( $Post_t \cdot [138 - 400\% FPL]_c^{2013}$ ) shifts individuals into direct purchase, and the third instrument ( $Post_t \cdot [< 138\% FPL]_c^{2013}$ ).  $Expansion_s$  shifts individuals into Medicaid. In practice, however, our instruments will shift multiple types of insurance simultaneously due to crowd-out.<sup>22</sup> We therefore include all three instruments when considering a given type of insurance.

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<sup>22</sup>Individuals making 100–138 percent of the FPL are eligible for exchange subsidies in all states and Medicaid in states that expanded the program. Medicaid expansions will therefore both increase the share of the population with Medicaid coverage and decrease the share of the population with (direct purchase) private insurance coverage relative to non-expansion states. Moreover, low-income individuals who gain health insurance through their employer will be less likely to enroll in Medicaid.

Again letting  $Clinics_{ct}$  denote the number of open retail clinics per 100,000 people in county  $c$  in year  $t$ , we then estimate the following second-stage regressions:

$$\begin{aligned} Clinics_{ct} = & \left\{ \beta \cdot \widehat{Insured}_{ct}, \beta_1 \cdot \widehat{Private}_{ct} + \beta_2 \cdot \widehat{Medicaid}_{ct} \right\} \\ & + \alpha'_4 \cdot Post_t \cdot [ < 138\% FPL ]_c^{2013} + \alpha'_5 \cdot Post_t \cdot Expansion_s \\ & + \delta' \cdot X_{ct} + \gamma'_c + \gamma'_t + \epsilon'_{ct}, \end{aligned} \quad (4)$$

where  $\widehat{Insured}_{ct}$ ,  $\widehat{Private}_{ct}$ , and  $\widehat{Medicaid}_{ct}$  denote the predicted shares of the population with health insurance coverage of any type, private insurance coverage, and Medicaid coverage at the county-year level from estimation of equation (3), respectively, and all other variables are defined as in equation (3). As in Section IV.A, we estimate this equation separately using either the share of the population with any type of health insurance coverage or the shares of the population with private insurance and Medicaid coverage as the key independent variables to examine both the average effects of health insurance coverage and differences in effects by insurance types.

One concern with our instrumental variables approach is that areas with differing employment and income profiles at baseline might have been on different trends in terms of retail clinic growth. We examine this possibility in Figure A8 by replicating the time-series path of retail clinic growth shown for the whole United States in Figure 1(a) separately across counties that are isolated by the different instrument components. In particular, we plot the average number of retail clinics per 100,000 people at the county-quarter level from 2010 to 2016 across counties in states that did and did not expand Medicaid under the ACA (subfigure (a)), across counties that had an above- versus below-median share of the population under 138 percent of the FPL at baseline by Medicaid expansion status (subfigure (b)), across counties by terciles of employment shares at baseline (subfigure (c)), and across counties with an above- versus below-median share of the population between 138 and 400 percent of the FPL at baseline.<sup>23</sup>

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<sup>23</sup>More precisely, these figures consider the shares of the population that were uninsured, between the ages of 18 and 64, and had incomes either below 138 percent of the FPL or between 138 and 400 percent of the FPL. Because the shares of the population with incomes in these two ranges are highly correlated within counties, we split counties with an above- versus below-median share of the population between 138 and 400 percent of the FPL by whether they additionally had an above- or below-median share of the population

Figure A8 shows that counties that are isolated by our different instrument components were on similar trends in terms of retail clinic growth before the onset of the ACA. However, there is a clear divergence in retail clinic penetration after 2014, with counties with baseline characteristics and Medicaid expansion status that would make them most likely to experience the smallest gains in Medicaid coverage and the largest gains in private insurance coverage under the ACA seeing marked increases in retail clinic concentration.<sup>24</sup> These patterns foreshadow our findings in Section V.

## V Results

### V.A Raw data

We begin by examining patterns in the raw data. Recall from Figure 2(a) that the share of the population with health insurance coverage across the United States (dark, solid line) increased by 7 percentage points from 2010 to 2016, with most of this growth concentrated from 2013 to 2015. Notably, the number of open retail clinics (light, dashed line) was also growing over this period, suggesting that retail clinic growth might be driven by increased insurance provision.

To examine whether growth in retail clinic penetration was concentrated in areas with growth in health insurance coverage, we examine how county-level changes in the number of retail clinics from 2013 to 2015 covary with county-level changes in health insurance coverage over the same period. In particular, Figure 5(a) groups counties into deciles based on changes in the share of the population with any type of health insurance from 2013 to 2015 and plots

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below 138 percent of the FPL at baseline in subfigure (d). Note that our empirical designs always control for the shares of the population in both of these groups, and thus only residual variation in these population shares within counties is exploited.

<sup>24</sup>Note that the income distribution at baseline should affect changes in insurance provision, and thus changes in retail clinic penetration, in both expansion and non-expansion states. In expansion states, counties with higher shares of the population with incomes below 138 percent of the FPL should experience greater increases in Medicaid coverage (and smaller increases in private insurance coverage). In non-expansion states, counties with fewer individuals with very low incomes should see greater increases in private insurance coverage both because of the employer mandate and because of subsidies to purchase insurance through the exchanges. In Figure A8(b), for example, we see the largest increase in retail clinics following the implementation of the ACA in non-expansion states with relatively few individuals with incomes below 138 percent of the FPL; these are places where we expect large increases in employer-sponsored or direct purchase coverage and no large increases in Medicaid coverage, even accounting for “woodwork” effects.

the average change in retail clinics per 100,000 people over the same period among counties in each decile. Perhaps surprisingly, there is no apparent relationship between growth in health insurance coverage and retail clinic growth. Counties saw an average growth of approximately 0.15 retail clinics per 100,000 people from 2013 to 2015, with strikingly similar growth in areas both with and without large gains in the share of the population with any health insurance coverage.

Although there is no association between changes in overall health insurance rates and retail clinic growth, this null effect could be masking heterogeneity by insurance type. Recall from Figure 2(b) that both private insurance and Medicaid coverage grew rapidly from 2013 to 2015, accounting for nearly all of the growth in health insurance coverage over the time period. To examine how retail clinic growth is associated with changes in different types of health insurance coverage, we replicate Figure 5(a) using either changes in private insurance or Medicaid coverage across counties from 2013 to 2015. As county-level changes in private insurance and Medicaid coverage are somewhat negatively correlated over this time period (see Figure 2(c)), we show both the raw association between changes in retail clinic concentration and changes in each type of insurance coverage as well as these patterns conditional on changes in all other coverage types.

Figures 5(b) and (c) show these relationships. Comparing the two subfigures, a striking pattern emerges: retail clinic growth is strongly positively correlated with growth in private insurance coverage and strongly negatively correlated with growth in Medicaid coverage. These patterns hold even conditional on changes in the other type of health insurance coverage, indicating that the opposing patterns are not simply driven by the negative correlation between changes in the two insurance types. Looking first to Figure 5(b), we see that counties that experienced the largest growth in private insurance coverage following the ACA saw an average increase of over 0.20 retail clinics per 100,000 people compared to an average increase of less than 0.10 clinics per 100,000 people in counties with the lowest growth in the share of the population with private insurance coverage. The association between changes in Medicaid coverage and retail clinic growth is even more pronounced, with counties that experienced high growth in Medicaid coverage seeing almost no increase on average in retail clinic concentration over the time period (Figure 5(c)). In contrast, areas

that experienced low growth in Medicaid coverage saw increases in retail clinic penetration that were comparable—even conditional on changes in private insurance coverage—to those experienced in areas with the highest growth in private insurance coverage.

These opposing responses of clinic concentration to growth in health insurance coverage of different types can further be seen in Figure A9. This figure plots the average number of retail clinics per 100,000 people at the county-quarter level from 2010 to 2016 by terciles of growth in private insurance (subfigure (a)) and Medicaid coverage (subfigure (b)) from 2013 to 2015. In both subfigures, we see that counties in different terciles of growth in the shares of the population covered by each insurance type under the ACA were on similar trends in terms of retail clinic growth before 2014. After the ACA’s implementation, however, counties with the largest increases in private insurance coverage and the smallest increases in Medicaid coverage diverge from other counties and experience sizable growth in retail clinic penetration. While level differences in retail clinic penetration in the pre-period could be driven either by baseline differences in insurance provision or local characteristics such as income that correlate with differences in insurance coverage, the timing of the changes in retail clinic concentration—which accord with the onset of the ACA—suggests that the patterns observed in Figure 5 are driven by changes in insurance provision rather than changes in underlying socio-demographics.

## V.B Two-way fixed effects

To control for general time trends and differences across locations, we estimate equations (1) and (2). As shown in column (1) of Table 2, there is no statistically significant effect of county-level growth in the share of the population with any type of health insurance coverage on the concentration of retail clinics. This finding counters the common belief that supply-side responses will necessarily accompany sizable insurance expansions.

However, as suggested by the patterns in the raw data, this null result masks important heterogeneity by insurance type. As shown in column (2) of Table 2, growth in private insurance coverage leads to significant increases in clinic growth, whereas growth in Medicaid coverage leads to significant reductions. The estimates suggest that growth in private insurance coverage of 5 percentage points—the average increase experienced by sample counties

over our time period—leads to a relative increase of 0.049 retail clinics per 100,000 people, or an increase of 8.0 percent relative to the mean. Moreover, growth in Medicaid coverage of 4 percentage points—the average increase experienced by counties in our sample over our time period—leads to a relative reduction of 0.051 retail clinics per 100,000 people, or a decrease of 8.4 percent relative to the mean.

These results suggest that retail clinics may have a preference for private insurance and a distaste for Medicaid. To examine this potential distaste for Medicaid further, we estimate an analogue of equation (2) in which we interact the county-level shares of the population with either private insurance or Medicaid coverage with an indicator denoting whether the county experienced an above-median increase in Medicaid coverage from 2013 to 2015. This indicator largely proxies for counties in states that expanded Medicaid but further incorporates the fact that some counties in Medicaid expansion (non-expansion) states nevertheless experienced small (large) increases in Medicaid coverage.

As shown in column (3) of Table 2, increases in private insurance coverage only led to relative growth in the concentration of retail clinics in areas that experienced below-median increases in Medicaid coverage. Moreover, increases in Medicaid coverage only led to relative reductions in the concentration of retail clinics in areas that experienced above-median increases in the share of the population covered by Medicaid. These patterns underscore retail clinics' apparent distaste for Medicaid and the lexicographic nature of their preferences: retail clinics appear to first avoid counties with high increases in Medicaid coverage and then locate in counties with large increases in private insurance coverage among this subset.<sup>25</sup>

## V.C Instrumental variables

The patterns in the raw data and the two-way fixed effects analyses suggest that growth in private insurance leads to increases in the concentration of retail clinics whereas growth in

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<sup>25</sup>An alternative way to see this is to estimate equation (2) separately among states that did and did not expand Medicaid. As shown in Table A3, increases in private insurance coverage only led to significant increases in the concentration of retail clinics in states that did not expand Medicaid. This is despite the fact that many counties in Medicaid expansion states also experienced sizable increases in private insurance coverage: counties in states that expanded Medicaid experienced an average increase in private insurance coverage of 2.2 percentage points over our sample period (range: -12.7 to 14.8 percentage points) compared to an average increase of 5.8 percentage points among counties in states that did not expand Medicaid (range: -5.2 to 23.5 percentage points).

Medicaid coverage dampens clinic penetration. While the timing of these patterns suggests that they are driven by changes in insurance provision rather than changes in other county-level characteristics that might also affect retail clinic presence, we isolate changes in health insurance coverage induced by policy changes to ensure that we are measuring the causal effects of health insurance provision. This instrumental variables approach has the additional benefit of correcting for measurement error in county-level insurance rates, which is likely to be significant since these data are based on self-reports from a survey of approximately 1 percent of the population.

As outlined in Section IV.B, we exploit provisions of the ACA to instrument for changes in health insurance coverage. First-stage results from estimation of equation (3), which show the impacts of these instruments on changes in the shares of the population with any health insurance coverage and health insurance coverage of different types, are shown in the top panel of Table 3. Looking first to column (1), we see that all three instruments strongly predict county-level changes in the share of the population with any health insurance coverage. This is because the majority of growth in insurance coverage over our sample period was driven by changes in private insurance and Medicaid coverage, and the instruments strongly predict changes in these insurance types: As shown in column (3), areas with higher employment and a greater share of the population between 138 and 400 percent of the FPL at baseline saw greater increases in private insurance coverage following the implementation of the ACA. Moreover, as shown in column (4), growth in Medicaid coverage was significantly higher following the ACA in counties that had a higher share of the population below 138 percent of the FPL at baseline and were in states that expanded Medicaid.

The two-stage least squares estimates from estimation of equation (4), which show the impacts of instrumented insurance changes on growth in the concentration of retail clinics, are provided in the bottom panel of Table 3. As shown in column (2), there is no statistically significant effect of the share of the population with any health insurance coverage on retail clinic concentration. However, as suggested by our previous analyses, this null effect masks important heterogeneity by insurance type: as shown in column (5), growth in private insurance coverage leads to large growth in clinic penetration, whereas growth in Medicaid coverage leads to large reductions in the concentration of retail clinics.

The effects by type of health insurance coverage are not only statistically significant but also economically meaningful. In particular, the estimates in column (5) indicate that growth in private insurance coverage of 5 percentage points—the average increase experienced by counties in our sample over our time period—leads to an increase of 0.16 retail clinics per 100,000 people, or about 25 percent relative to the mean. The effects for Medicaid are even more pronounced, with growth in Medicaid coverage of 4 percentage points—the average increase experienced by sample counties over our time period—leading to a reduction of 0.20 retail clinics per 100,000 people, or over 30 percent relative to the mean. Notably, the instrumental variable estimates are larger than the corresponding two-way fixed effects estimates for both private insurance and Medicaid coverage, suggesting that the estimates in Table 2 were attenuated by measurement error.

## V.D Openings versus closings

Variation in the number of retail clinics over our sample period is driven both by entries and exits; as noted in Section II, there were over 50 entries and nearly 22 exits on average in each quarter from 2010 to 2016 (see Figure A1). To examine whether the effects of health insurance coverage on the concentration of retail clinics are driven by entries, exits, or both, we estimate analogues of our two-way fixed effects and instrumental variables specifications using either the number of clinic entries or exits at the county-year level as the dependent variable. As entries and exits are flow measures rather than stocks, we specify the right-hand side of each equation in first differences when considering these outcomes. We further control for the number of retail clinics per 100,000 people in the previous period to account for the fact that openings (closings) are less (more) common in markets with many retail clinics.<sup>26</sup>

Table 4 presents results from these analyses.<sup>27</sup> We again begin by considering the effects of health insurance coverage of any type. As shown in columns (1) and (2), there are no significant effects of overall health insurance rates on either the entry (top panel) or exit (bottom panel) of clinics. The point estimates go in the anticipated directions, however,

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<sup>26</sup>The estimating equations for these analyses are presented in Appendix D.

<sup>27</sup>First-stage results for the instrumental variables designs are provided in Table A4. The results confirm that annual changes in our instruments are strong predictors of annual changes in insurance growth.

with the coefficient on insurance coverage being positive in the entry specifications and negative in the exit specifications. Moreover, the point estimates from the instrumental variables analyses (column (2)) are again larger than the two-way fixed effects specifications (column (1)).

Columns (3) and (4) in Table 4 turn to examining the effects of different types of health insurance coverage. Looking first to the results for private insurance, the top panel of Table 4 shows that growth in private insurance coverage leads to large increases in clinic entry. In particular, the instrumental variables estimate in the top panel of column (4) indicates that growth in private insurance coverage of 0.86 percentage points—the average annual change among sample counties over our sample period—leads to an annual increase of 0.026 retail clinic entries per 100,000 people, or over 35 percent relative to the average entry rate. Moreover, growth in private insurance coverage leads to significant reductions in clinic exit: as shown in the bottom panel of column (4), private insurance growth of 0.86 percentage points reduces annual clinic exit by 0.011 per 100,000 people, a 35 percent reduction relative to the average exit rate. These results show that the positive effects of private insurance coverage on clinic growth are driven both by increased clinic entry and reduced clinic exit.

Turning to the results for Medicaid coverage, we see that the dampening of clinic growth in areas with increases in Medicaid coverage is driven predominately by increased exit. As shown in the bottom panel of column (4), increased Medicaid coverage of 0.69 percentage points—the average annual change in sample counties over our sample period—leads to an increase in annual clinic exits of 0.013 per 100,000 people, an increase of over 40 percent relative to the average exit rate. These results call into question the viability of retail clinics in markets with increasing Medicaid coverage and suggest that such clinics are not well positioned to help absorb additional health care demand stemming from such expansions.

## V.E Identification and robustness

We conduct a number of additional analyses to probe the validity of our empirical design and the robustness of our findings. We begin by presenting output from two-step balancing regressions that show how our empirical strategies address endogeneity concerns stemming from the relationship between local socio-demographics, health insurance rates, and retail

clinic entry patterns. To implement these two-step balancing regressions, we first predict retail clinic concentration using local socio-demographics.<sup>28</sup> We then examine how insurance provision at the county-year level correlates with this measure of predicted, rather than actual, retail clinic penetration across various specifications. For each insurance type, the top three rows of Figure 6 show associations with predicted retail clinic concentration from estimation of equation (2) without county fixed effects (“Cross-section”), equation (2) with county fixed effects (“County FEs”), and equation (4) (“2SLS”), respectively.<sup>29</sup> For comparison, the fourth rows show the effects of each insurance type on actual retail clinic penetration from estimation of equation (4) first shown in Table 3 above (“Baseline estimate”).

As shown by the cross-sectional specification in Figure 6, the share of the population with Medicaid is negatively associated with variation in retail clinic penetration driven by local socio-demographics. Although the relationship between Medicaid coverage and predicted retail clinic concentration is attenuated when we consider within-county changes in predicted clinic penetration and health insurance provision, the relationship remains significant in the two-way fixed effects design. Moreover, the positive association between private insurance coverage and predicted retail clinic concentration becomes slightly stronger when within-county changes are considered. Reassuringly, however, our instrumental variables approach weakens these relationships, with the association between each insurance type and variation in retail clinics driven by local socio-demographics becoming insignificant when we isolate changes in insurance provision driven by changes in policy.<sup>30</sup> Moreover, the magnitudes of the associations in these balancing regressions are very small compared to our main results,

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<sup>28</sup>We include all county-year socio-demographics listed in Table 1 under the headings “Basic demographics” and “Income and education” as explanatory variables in this first-stage regression. This regression has an F-stat of 98.0 and an adjusted  $R^2$  of 0.286.

<sup>29</sup>Since the outcome in these analyses is predicted using the vector of time-varying, county-level controls  $X_{ct}$ , we exclude these controls from equations (2) and (4) when estimating the second-stage regressions.

<sup>30</sup>Figure A10 displays results from one-step balancing regressions in which we examine how insurance provision at the county-year level correlates with select socio-demographics across our specifications. As shown by the cross-sectional specifications, differences in the shares of the population with different types of health insurance coverage across counties predict differences in population density, share White, median income, and the employment rate. Notably, however, many of these associations are attenuated when we consider within-county changes in demographics and health insurance provision, and these relationships are attenuated even further in our instrumental variables approach. While reassuring, these balancing regressions do not account for the association between each potential confounder and retail clinic penetration, and thus we prefer the two-step version that weights each potential confounder by its importance in predicting our outcome of interest.

providing reassuring evidence that our findings are driven by changes in insurance provision rather than changes in local socio-demographics that might independently influence the entry and exit decisions of clinics.

We also examine the sensitivity of our estimates to several empirical choices. As outlined in Section IV, we include time-varying, county-level controls in our baseline specifications and weight county-level observations by county population in 2010. Reassuringly, Figure A11 shows that our baseline estimates first presented in Tables 2 and 3 are quantitatively robust to excluding time-varying socio-demographic controls and to giving all observations equivalent weight in estimation. This is true both for our estimated effects of the share of the population with private insurance coverage (subfigure (a)) and for the share of the population covered by Medicaid (subfigure (b)).

We conduct two additional robustness analyses for our instrumental variables designs. First, recall that we consider states as having expanded Medicaid in our primary specification if they expanded their Medicaid programs to include individuals making up to at least 138 percent of the FPL by 2014. This includes 20 states that did so in 2014 and five states (California, Connecticut, Delaware, Minnesota, and New Jersey) that did so in part between 2010 and 2013. Because the timing of our instrumental variables analysis—which considers 2014 onward as the “post” period for Medicaid expansions—will be less accurate for these early expanders, we replicate our analysis dropping counties in these five early expansion states. As shown in the final row of each subfigure in Figure A11, the results are very similar when we exclude early expanders.

Second, recall that our primary specification uses variation in the share of the population that was employed in 2013 to capture variation in the share of the population that was affected by the ACA’s employer mandate. Because the employer mandate only required employers with at least 50 full-time employees to provide health insurance coverage, and because small- to mid-size firms were the least likely to offer coverage to their employees before the ACA (KFF, 2013), we replicate our analysis using shares of the population employed by firms of different sizes when constructing the instrument. We approximate the shares of the population working in firms with 50-99, 50-249, 50-499, 50-999, and 50+ employees in 2013 by multiplying county-level employment shares from the one-year ACS by state-level shares

of employees working in establishments of these sizes from the QCEW.

As shown in Figure A12, the results are very similar when we use either the overall employment rate or the share of the population employed in firms with 50+ employees. This similarity in findings is consistent with the fact that most workers are employed by large firms (QWI, 2013). Moreover, the results for private insurance are slightly more pronounced when use the share of the population employed by firms with 50-99 or 50-249 employees. This is to be expected since most firms with over 200 employees already offered health insurance coverage before the ACA, and thus they were largely unaffected by the employer mandate (KFF, 2013).

## VI Extensions

### VI.A What drives the negative supply-side effects of Medicaid?

In this section, we show that the negative supply-side effects of Medicaid are likely driven by the program's low reimbursement rates. We further show that there is no evidence to support other potential mechanisms for these perverse supply-side responses, such as expanded Medicaid coverage failing to generate increased demand for on-demand health care clinics.

**Low provider reimbursement rates** We begin by considering the role of Medicaid reimbursement rates. As outlined in Section III, the positive (negative) supply-side effects of private insurance (Medicaid) coverage should be concentrated in areas that were not accepting Medicaid at baseline. Because clinics should be more likely to accept Medicaid in areas with higher reimbursement rates under the program, it follows that the supply-side responses to growth in private insurance and Medicaid coverage should be more pronounced in areas with low Medicaid reimbursement rates.<sup>31</sup>

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<sup>31</sup>More precisely, the theory predicts that clinics should be more likely to accept Medicaid in areas with higher reimbursement rates under Medicaid *relative* to other insurers. Combining information on Medicaid fee-for-service to private insurance reimbursement ratios for office visits in 32 states in 2010 from a report from the Government Accountability Office (GAO, 2014) with data on Medicaid reimbursement rates in the same year from Alexander and Schnell (2019), we find that relative Medicaid payments are highly correlated with Medicaid fee-for-service payment levels ( $\rho = 0.73$ ). In what follows, we use Medicaid payment levels

To examine whether this prediction is borne out in the data, we use information on state-level Medicaid reimbursement rates for office visits from Alexander and Schnell (2019). As shown in Figure A13(a), there was wide variation across the United States in the amount that providers were reimbursed under Medicaid for a low-complexity office visit (CPT 99201) in 2010.<sup>32</sup> Splitting the sample by terciles of these Medicaid reimbursement rates, we replicate Figures 5(b) and (c) separately for states in the bottom and top terciles of baseline Medicaid reimbursements. In particular, for each payment tercile, we group counties into deciles based on changes in the share of the population with private insurance or Medicaid coverage from 2013 to 2015 and plot the average change in retail clinics per 100,000 people over the same period among counties in each decile. As before, we focus on changes in a given type of insurance conditional on changes in other types of health insurance coverage.

Figure 7 presents these results. Despite losing some precision when focusing on a subset of states, we see in Figure 7(a) that the positive association between growth in private insurance coverage and clinic penetration is most pronounced in counties in the bottom tercile of Medicaid payments (i.e., locations in which clinics are least likely to accept Medicaid). Consistent with the theory outlined in Section III, there is no apparent relationship between growth in private insurance coverage and changes in retail clinic penetration among counties in states with the highest Medicaid payments. As shown in Figure 7(b), similar patterns are observed for Medicaid coverage, with the negative effects of Medicaid on the concentration of retail clinics being the most pronounced in locations with low Medicaid reimbursement rates at baseline.

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since we have this information for all states, although the results in this section are very similar if we instead use payment ratios.

<sup>32</sup>We verify that retail clinics are more likely to accept Medicaid in states with higher Medicaid reimbursement rates. In particular, Figure A13(b) shows that states with higher reimbursements under Medicaid in 2015 (the latest year available in the reimbursement data) were much more likely to have some form of Medicaid accepted at their CVS MinuteClinics in 2020. For this exercise, we hand-collected data on Medicaid acceptance among CVS MinuteClinics—which accounted for an average of 49.6 percent of all retail clinics in each year over our sample period—in 2020 by navigating to CVS’s “Insurance Check” website, selecting an insurance carrier and plan from the dropdown menu, filling in a zip code, and then recording all clinics within 20 miles of the chosen zip code that accepted the selected carrier and plan. Repeating this for all combinations of carriers, plans, and zip codes, we recovered a comprehensive list of the locations in which CVS MinuteClinics accepted at least one Medicaid plan in 2020. Among the 34 states with CVS MinuteClinics, these data indicate that at least some form of Medicaid coverage was accepted at the company’s clinics in 23 (67.6 percent). We thank Danielle Handel and Jimmy Kim for help with this exercise.

**Limited demand-side responses** The results in Figure 7 suggest that provider prices play an important role in generating the observed supply-side responses. We conduct additional analyses to investigate the role of other mechanisms that might also contribute to the negative supply-side effects of Medicaid coverage that we observe. Notably, because individuals moving from being uninsured to having Medicaid may increase their use of more traditional health care delivery outlets (Finkelstein et al., 2012; Taubman et al., 2014; Garthwaite et al., 2019), Medicaid expansions might not generate growth in demand for retail clinics that is necessary to promote entry.<sup>33</sup> Relatedly, if federally qualified health centers (FQHCs)—community-based health centers that predominately serve low-income populations—are more likely to enter areas with growth in Medicaid coverage, then Medicaid patients may not need to rely on retail clinics due to the growing presence of FQHCs.

We conduct two sets of analyses to examine these possibilities. First, using county-level information from the HRSA on the number of primary care physicians and nurse practitioners (“primary care providers”) per capita, we split the sample by terciles of primary care providers per capita in 2010 and replicate Figures 5(b) and (c) separately among counties in the bottom and top terciles. If the negative supply-side effects of Medicaid are driven by patients substituting away from retail clinics to other types of care when they gain access to Medicaid, then the perverse supply-side responses should be more pronounced in areas with a higher concentration of alternative resources (i.e., areas in which there is more scope for substitution). Second, recognizing that FQHCs were expanding rapidly over our sample period, we re-estimate our main analyses controlling for the number of FQHCs per capita at the county-year level from the HRSA. If retail clinics avoid areas with increasing Medicaid coverage because they are forced to compete with FQHCs for low-income patients in such areas, then the coefficient on Medicaid coverage should be attenuated when controlling for FQHC presence.

These results are shown in Figures A14 and A11, respectively. Looking first to Figure

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<sup>33</sup>Even if Medicaid expansions generate increased demand for retail clinics, the simultaneous increases in demand for more traditional health care delivery outlets might create competition for nurse practitioners. This competition could in turn lead to staffing difficulties at retail clinics. However, we would expect similar labor market effects to occur after private insurance expansions, and yet the concentration of retail clinics expands following growth in private insurance coverage. The negative effects of Medicaid expansions are therefore unlikely to be driven by difficulties in hiring induced by insurance expansions.

[A14](#), we see that the negative supply-side effects of Medicaid coverage are largest in areas with the least primary care providers per capita.<sup>34</sup> Because areas with few providers at baseline offer the least scope for use of traditional health care delivery mechanisms, increased demand for retail clinics following Medicaid expansions should, if anything, be most pronounced in such areas. Moreover, looking to the final row of each subpanel in Figure [A11\(b\)](#), we see that our results are essentially unchanged when we control for the number of FQHCs per capita at the county-year level.<sup>35</sup> These analyses suggest that limited demand-side responses are unlikely to explain the negative supply-side effects of Medicaid coverage that we observe.

## VI.B Are the observed supply-side responses (in)efficient?

An outstanding question is whether the heterogeneous supply-side responses to different types of health insurance expansions that we document are efficient from the perspective of the social planner. If private insurance expansions induce clinics to enter previously underserved areas, whereas Medicaid expansions induce clinics to exit (or stop entering) areas that had an excess supply of health care resources at baseline, then such responses might simultaneously address existing access barriers while limiting the scope for unnecessary service provision. However, if clinics enter areas with sufficient baseline resources following private insurance expansions and exit areas with insufficient baseline resources following Medicaid expansions, then such responses will exacerbate inequities in health care access and may lead to additional and unnecessary service use in well-resourced areas. Concerns over excess service provision at retail clinics have been particularly pronounced given the clinics' convenience and relatively low prices ([Ashwood et al., 2016](#)).

To examine whether the supply-side effects of health insurance that we observe promote

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<sup>34</sup>This analysis will be closely related to the findings in Figure [7\(b\)](#) if provider location decisions are affected by Medicaid reimbursement rates. In particular, if areas with low reimbursements rates are both less likely to have Medicaid accepted at local clinics and have fewer local providers, then we might observe that the negative supply-side effects of Medicaid coverage are concentrated in areas with few local providers simply because of the positive correlation between provider concentration and Medicaid acceptance that runs through reimbursement rates. However, Figure [A14\(b\)](#) shows that effect heterogeneity by primary care providers per capita at baseline is very similar when controlling for baseline Medicaid reimbursement rates.

<sup>35</sup>Figures [A14](#) and [A11](#) further present analogous results for private insurance. We present these results for consistency, but note that these tests are predominately intended to examine drivers of the negative supply-side effects of Medicaid coverage (rather than the positive supply-side effects of private insurance coverage). The results for private insurance look very similar across terciles of baseline primary care providers per capita and are unaffected by controlling for FQHC presence.

allocative efficiency, we examine how the effects vary by the need for additional health care resources at baseline. In particular, we use information on whether each county was designated a primary care shortage area in 2010 by the HRSA. These designations are determined using information on the number of providers per capita, distance to the nearest source of care, local poverty rates, and measures of infant health. We split the sample into counties that were and were not designated as full primary care shortage areas at baseline and replicate Figures 5(b) and (c) separately among these two groups of counties.<sup>36</sup> As before, we focus on changes in a given type of insurance conditional on changes in other types of health insurance coverage.

Figure A15 presents results from this analysis.<sup>37</sup> Looking first to the left subplots, we see that the positive supply-side effects of private insurance expansions are predominately concentrated in non-shortage areas. In contrast, the negative supply-side effects of Medicaid expansions (right subplots) are observed across both shortage and non-shortage areas. These results show that the heterogeneous supply-side responses to different types of health insurance expansions are unlikely to address existing access barriers and may exacerbate unnecessary service provision in well-resourced areas.

### VI.C Are the results specific to retail clinics?

To examine whether our results are likely to extend to other providers in the primary care market, we examine the location patterns of urgent care centers. Like retail clinics, urgent care centers are on-demand health care clinics that have experienced significant growth in the past two decades. In contrast to retail clinics, however, urgent care centers are typically staffed by medical doctors rather than nurse practitioners, treat both minor and moderately severe conditions rather than only minor illnesses, and are often owned and operated by

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<sup>36</sup>The HRSA divides counties into three groups: (1) the whole county is designated as a primary care shortage area (“full shortage”), (2) one or more parts of the county is designated as a primary care shortage area (“partial shortage”), or (3) none of the county is designated as a primary care shortage area (“no shortage”). In 2010, 42 (40) (18) percent of U.S. counties were designated as full (partial) (non-) primary care shortage areas. A similar distribution is observed among our sample of 555 counties, with 37, 49, and 14 percent of counties being designated as full, partial, and non-primary care shortage areas, respectively. Since only 76 sample counties are designated as non-shortage areas, we group partial shortage and non-shortage areas and compare outcomes relative to full shortage areas.

<sup>37</sup>As in Figure A14, we show the results both conditional on baseline reimbursement rates under Medicaid (Figure A15(b)) and without this additional control (Figure A15(a)).

hospital systems rather than major retail outlets.

As outlined in Section II, data covering the locations of all 12,721 urgent care centers operating in the United States in 2021 come from the NUCR database. As these data contain no information on urgent care centers that opened and closed before 2021, it is difficult to replicate our primary analyses for these clinics. However, we can examine how the association between the local provision of health insurance and the concentration of urgent care centers in the cross-section compares to that observed among retail clinics to investigate whether these clinic types exhibit similar location patterns. In particular, we examine how the number of retail clinics per 100,000 people in 2016 and the number of urgent care centers per 100,000 people in 2021 covary at the county level with the share of the population covered by different types of insurance in the 2012–2016 and 2016–2020 ACS, respectively. As in the within-county plots shown in Figure 5, we consider the correlation between a given type of insurance coverage and clinic penetration both unconditional and conditional on other types of insurance coverage (with the share of the population that is uninsured serving as the omitted category).

As shown in Figure 8(a), the county-level concentration of retail clinics in 2016 was strongly increasing in the share of the population covered by private insurance and strongly decreasing in the share of the population covered by Medicaid. The raw gradients are striking: counties with the highest rates of private insurance coverage or the lowest rates of Medicaid coverage had nearly 1.2 retail clinics per 100,000 people, whereas counties at the other end of each spectrum were largely unserved by retail clinics. Although these relationships are somewhat attenuated when controlling for the shares of the population covered by other types of health insurance, pronounced gradients in clinic penetration persist among counties with similar insurance profiles other than the share of the population covered by private insurance or Medicaid.

Moreover, as shown in Figure 8(b), the relationship between local insurance composition and the concentration of urgent care centers in 2021 closely mirrors the patterns observed among retail clinics in 2016. While counties with the lowest (highest) share of patients covered by private insurance (Medicaid) had less than 3.5 urgent care centers per 100,000 people in 2021, counties with the highest (lowest) shares of patients covered by private

insurance (Medicaid) had nearly 30 percent more clinics per capita. These gradients are again attenuated when conditioning on the shares of the population with other types of health insurance coverage, in part due to the fact that conditioning on other insurance profiles absorbs much of the variation in the share of the population with private insurance coverage across counties. Nevertheless, the location patterns in the raw data indicate that urgent care centers—like retail clinics—are disproportionately located in areas with high rates of private insurance and low rates of Medicaid coverage.

## VII Discussion and conclusion

Seminal work by [Arrow \(1963\)](#) argued that health insurance expansions should lead the supply side to expand. The economics behind this insight is simple: Health insurance expansions reduce the prices paid by consumers at all levels of service provision, thereby shifting the demand curve outwards. This shifting of the demand curve leads to upward movement along the supply curve—that is, the supply side expands—to arrive at a new equilibrium.

We show that this understanding of health insurance expansions is incomplete. To understand how health insurance expansions affect the supply side, one must also take into account how such expansions affect the prices paid to providers. In markets with a mix of patients covered by insurance that pays either administered or market-based pricing, a common feature of modern health insurance markets in both the United States and abroad, health insurance expansions can lead the supply side to contract if such expansions (1) shift patients into programs with administered prices and (2) these administered prices are sufficiently low such that firms preferred to serve patients not covered by the program at baseline.

We begin by confirming the predictions of [Arrow \(1963\)](#) for health insurance expansions that reduce the prices paid by consumers but are likely to weakly increase the prices received by providers. Leveraging growth in private insurance stemming from the ACA, we find that expansions of private insurance increase the concentration of retail clinics. These effects are large and indicate that an increase in private insurance coverage of 5 percentage points—the average increase experienced by sample counties following the ACA—lead to an increase in

clinic concentration of nearly 25 percent relative to the mean.

Moreover, additional analyses show that these positive supply-side effects of health insurance are likely the result of outward shifts in demand. In particular, although we consider employer-sponsored and direct purchase insurance jointly as “private” insurance coverage throughout, consumers who directly purchase their health insurance are more likely to be covered by high-deductible health plans (HDHPs) than consumers who receive their insurance through their employer.<sup>38</sup> If patients with HDHPs are unlikely to have reached their deductibles, and thus are effectively uninsured for the services that they receive, then the outward shift in demand—and the subsequent positive supply-side effects of health insurance—should be less pronounced for expansions of direct purchase versus employer-sponsored coverage. Examining the effects of direct purchase and employer-sponsored coverage separately, we find that the positive supply-side effects of private insurance are driven predominately by employer-sponsored coverage.<sup>39</sup> Consistent with [Arrow \(1963\)](#), this suggests that the positive supply-side effects of health insurance are mediated by the generosity of cost-sharing for patients.

However, not all insurance expansions weakly increase the prices received by providers. In the United States, Medicaid tends to pay providers less than private coverage ([Alexander and Schnell, 2019](#)), and these lower reimbursement rates are compounded by administrative hassles that providers face when billing the program ([Dunn et al., 2021](#)). Despite being

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<sup>38</sup>In 2016, over 50 percent of adults with direct purchased coverage were enrolled in HDHPs compared to only 39 percent among adults with employment-based coverage ([NCHS, 2017](#)).

<sup>39</sup>We conduct two sets of analyses to examine whether clinic penetration responds differently to private insurance plans that are directly purchased rather than provided through employers. First, we estimate analogues of our primary specifications that include the county-year level shares of the population with employer-sponsored coverage and direct purchase insurance separately on the right-hand side instead of the share of the population with private insurance. While the coefficient on the share of the population with employer-sponsored coverage is twice the corresponding baseline estimate for all private insurance and statistically significant at the 1 percent level in the instrumental variables specification, the coefficient on the share of the population with direct purchase insurance is negative and not statistically significant at conventional levels (see Table A5). Second, we estimate versions of equations (4) that leverage each of our private insurance instruments separately. Because the first instrument—the baseline employment rate—predominately shifts individuals into employer-sponsored health insurance while the second instrument—the baseline population share with incomes between 138 and 400 percent of the FPL—predominately shifts individuals into direct purchase coverage, the estimated effects of private insurance coverage from versions of equation (4) that only include the first or second private insurance instrument can be loosely thought of as showing the effects of employer-sponsored and direct purchase insurance, respectively. The results from this analysis further suggest that the positive supply-side effects of private insurance coverage are driven predominately by employer-sponsored coverage (see Table A6).

very generous for patients, with limited to no cost-sharing, we find that in the case of on-demand health care clinics—key contributors to the expansion of health care systems in recent years—recent growth in Medicaid coverage caused the supply side to contract. The effects are large, with the negative effects on clinic penetration in counties with the average increase in Medicaid coverage being similar in magnitude, but opposite in sign, to the positive effects on the concentration of clinics in counties with average growth in private insurance coverage.

The supply-side effects that we document are likely inefficient from the perspective of the social planner. While growth in clinics following private insurance expansions is larger in areas with sufficient baseline resources, reductions in clinic penetration following growth in Medicaid coverage are more pronounced in areas with fewer providers per capita. This suggests that supply-side responses to insurance expansions have the potential to contribute to unnecessary service use in well-resourced areas while further limiting access in areas with an already limited supply of providers. We note, however, that the effects on consumer welfare remain uncertain. Although an increase in health care access might improve patient welfare in areas with growth in private insurance coverage, insurers might increase premiums for consumers in response to increased service use. A careful examination of how the effects of health insurance expansions on the entry and location patterns of firms affect consumer welfare is a fruitful area for future research.

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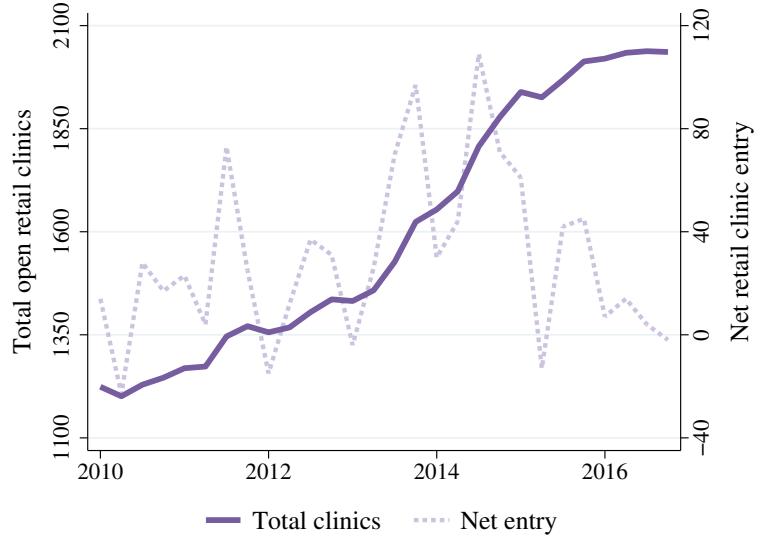
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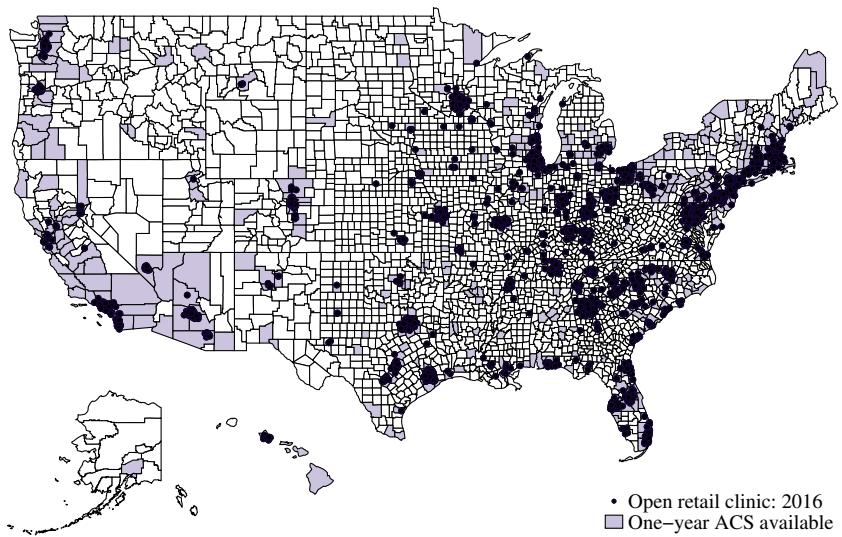
## VIII Figures

Figure 1: Retail clinics across the United States

(a) Quarterly total and net entry: 2010–2016

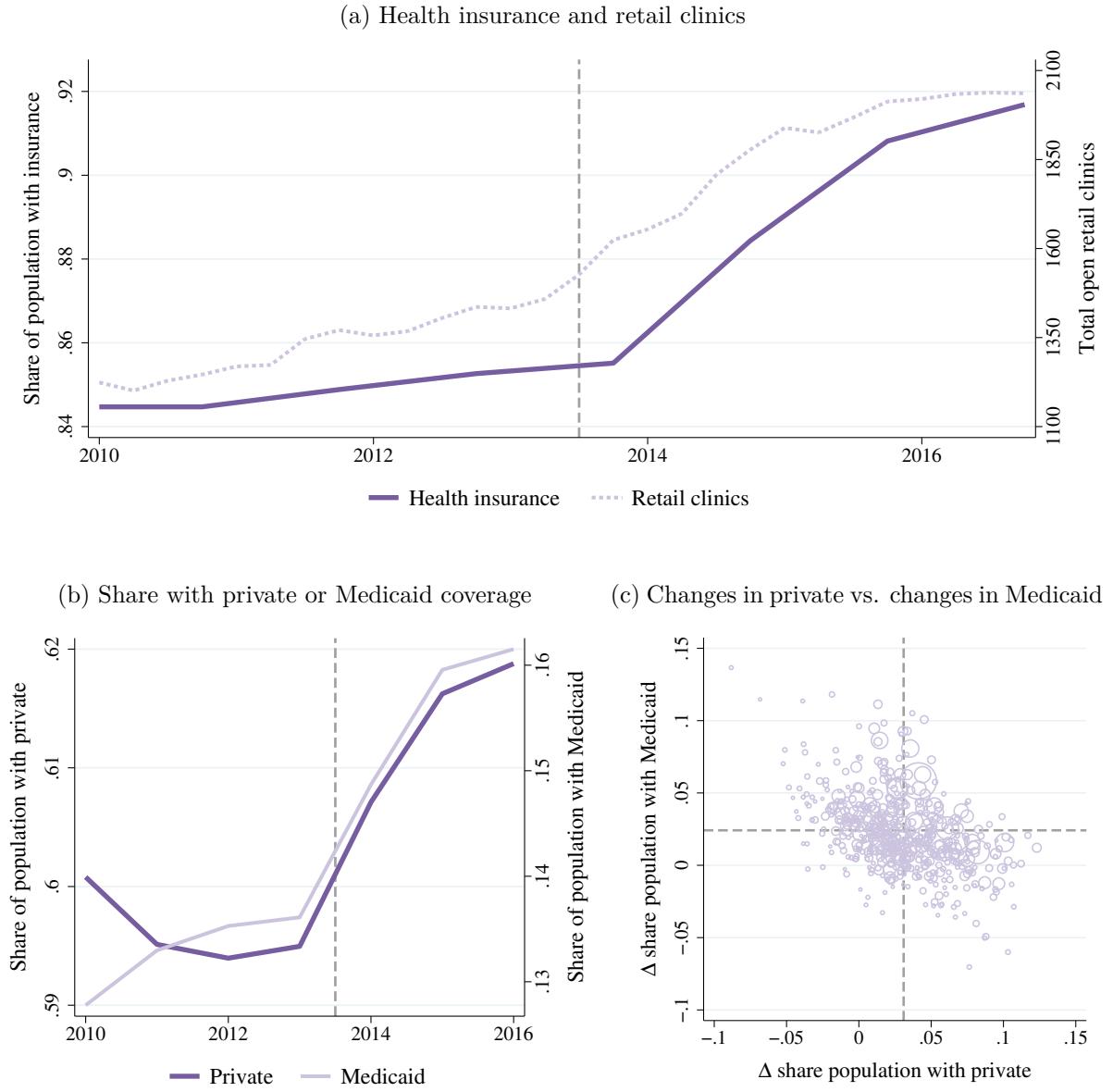


(b) Locations: 2016



Notes: The above figures show the number and locations of retail clinics across the United States. Subfigure (a) shows the total number of retail clinics (dark, solid line) and net entry (light, dashed line) quarterly from 2010 to 2016. “Net entry” refers to the total number of openings net of the total number of closings in a given quarter; refer to Figure A1 for openings and closings separately over the same period. Subfigure (b) shows the locations of retail clinics in 2016 (geo-coded dots). Subfigure (b) further displays counties with data available in every one-year ACS from 2010 to 2016 (shaded counties); counties must have a population of 65,000 or more to be included in the one-year ACS in a given year. Data on retail clinics come from Merchant Medicine.

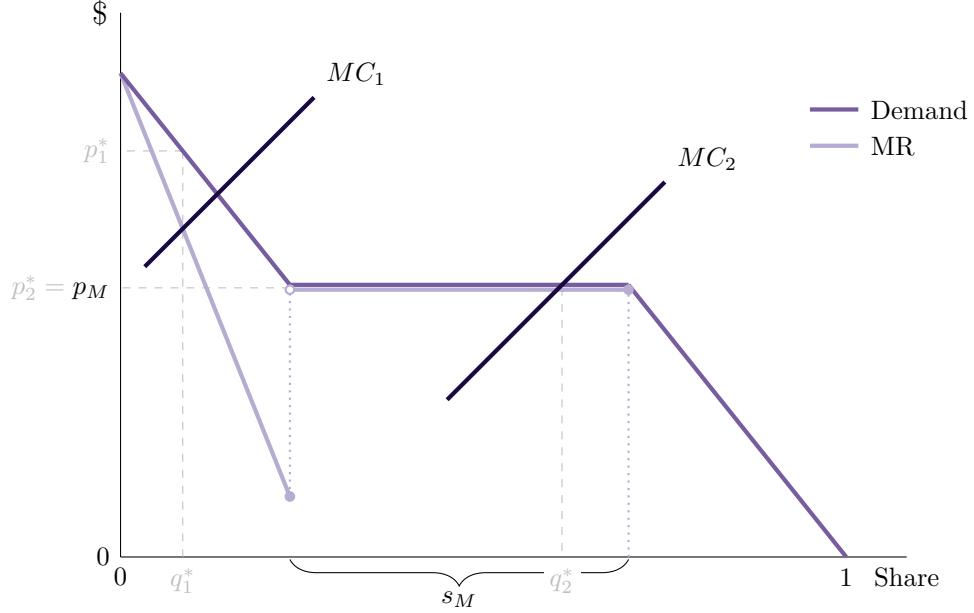
Figure 2: Changes in health insurance coverage: 2010–2016



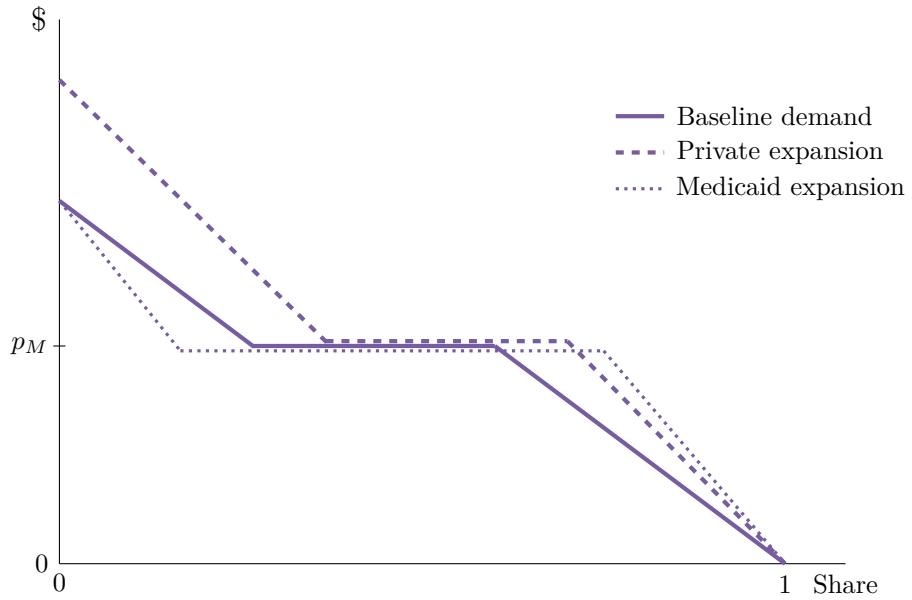
Notes: The above figures show changes in the share of the population with different types of health insurance from 2010 to 2016. Subfigure (a) displays the annual share of the population with health insurance of any type (dark, solid line) and the quarterly number of retail clinics (light, dashed line). Subfigure (b) displays the annual share of the population with private insurance (dark, thick line) or Medicaid coverage (light, thin line). Subfigure (c) shows how county-level changes in the share of the population covered by Medicaid from 2010 to 2016 covary with county-level changes in the share of population with private insurance over the same period. The size of the markers in subfigure (c) denotes county-level population in 2010; the dashed lines denote the population-weighted median of changes in Medicaid coverage and private insurance from 2010 to 2016. Private insurance includes employer-sponsored coverage and direct purchase. Data on retail clinics come from Merchant Medicine; data on health insurance come from the one-year ACS.

Figure 3: Clinic in market with both administered and market-based prices

(a) Firm's baseline problem

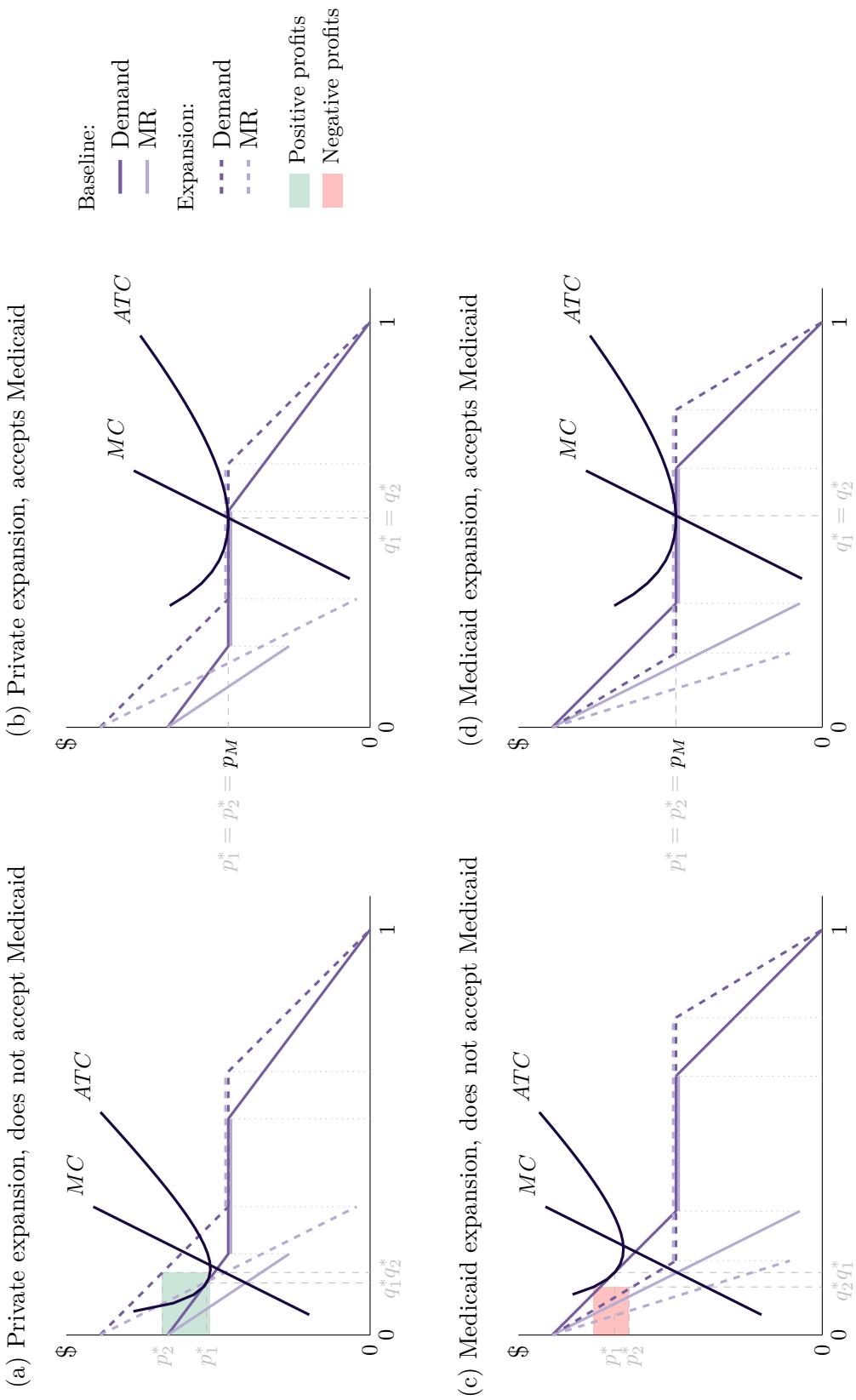


(b) Impact of insurance expansions on demand



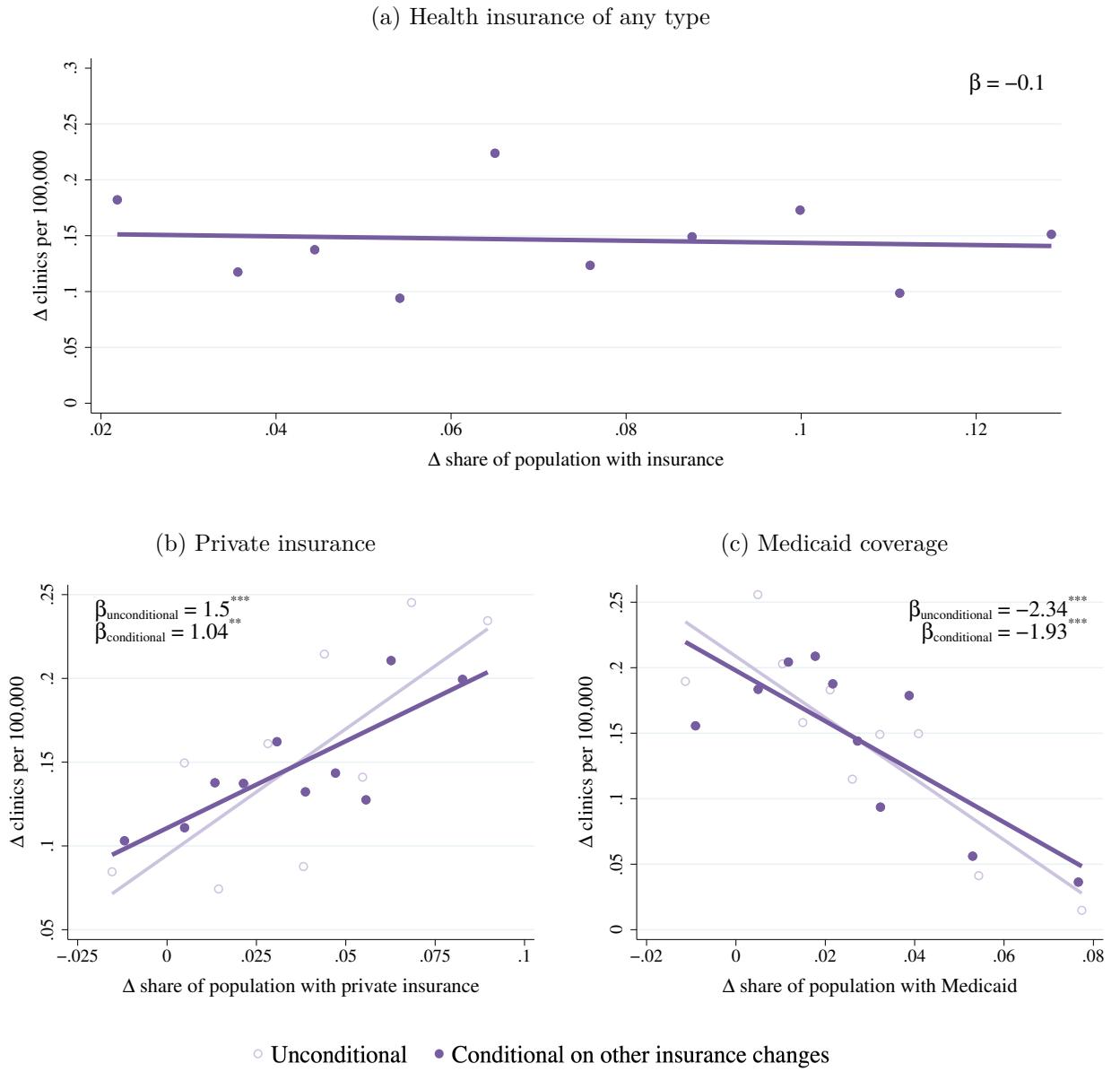
Notes: The above figures consider a firm in a market with both administered and market-based prices. Subfigure (a) displays the total demand curve (dark purple line) and associated marginal revenue curve (light purple line); the demand curve is perfectly elastic at the administered Medicaid price ( $p_M$ ) with length equivalent to the share of the population covered by Medicaid ( $s_M$ ). Subfigure (a) additionally shows how prices and quantities are determined when there is a single intersection between marginal revenue and marginal costs; the cases in which there are two intersections or no intersections are shown in Figure A4. Subfigure (b) shows how the demand curve changes under expansions of private insurance (dashed line) and Medicaid coverage (dotted line).

Figure 4: Effects of insurance expansions on clinic profits



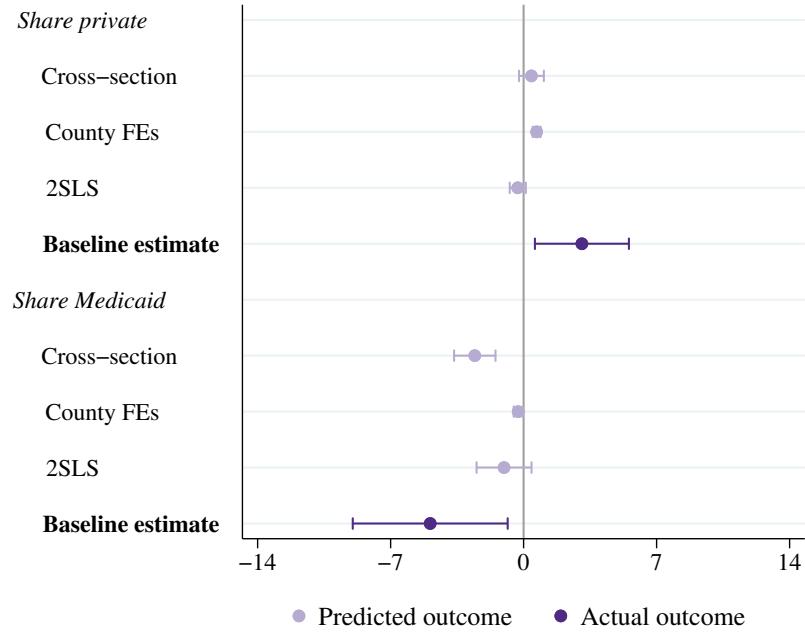
Notes: The above figures show how expansions of private insurance (subfigures (a) and (b)) and Medicaid coverage (subfigures (c) and (d)) affect firm profits, both when the firm accepted Medicaid patients at baseline (subfigures (b) and (d)) and when the firm did not (subfigures (a) and (c)). As shown in subfigure (a), private expansions increase firm profits when the firm was not accepting Medicaid patients at baseline; in contrast, Medicaid expansions decrease profits for such firms (subfigure (c)). As shown in subfigures (b) and (d), firm profits are unaffected by both private and Medicaid expansions if the firm accepted Medicaid patients at baseline and the marginal cost curve intersected the marginal revenue curve on the perfectly elastic component at baseline. Figure A5 shows the impacts of private and Medicaid expansions when there is no intersection between marginal costs and marginal revenue at baseline.

Figure 5: Changes in retail clinic presence versus changes in health insurance: 2013–2015



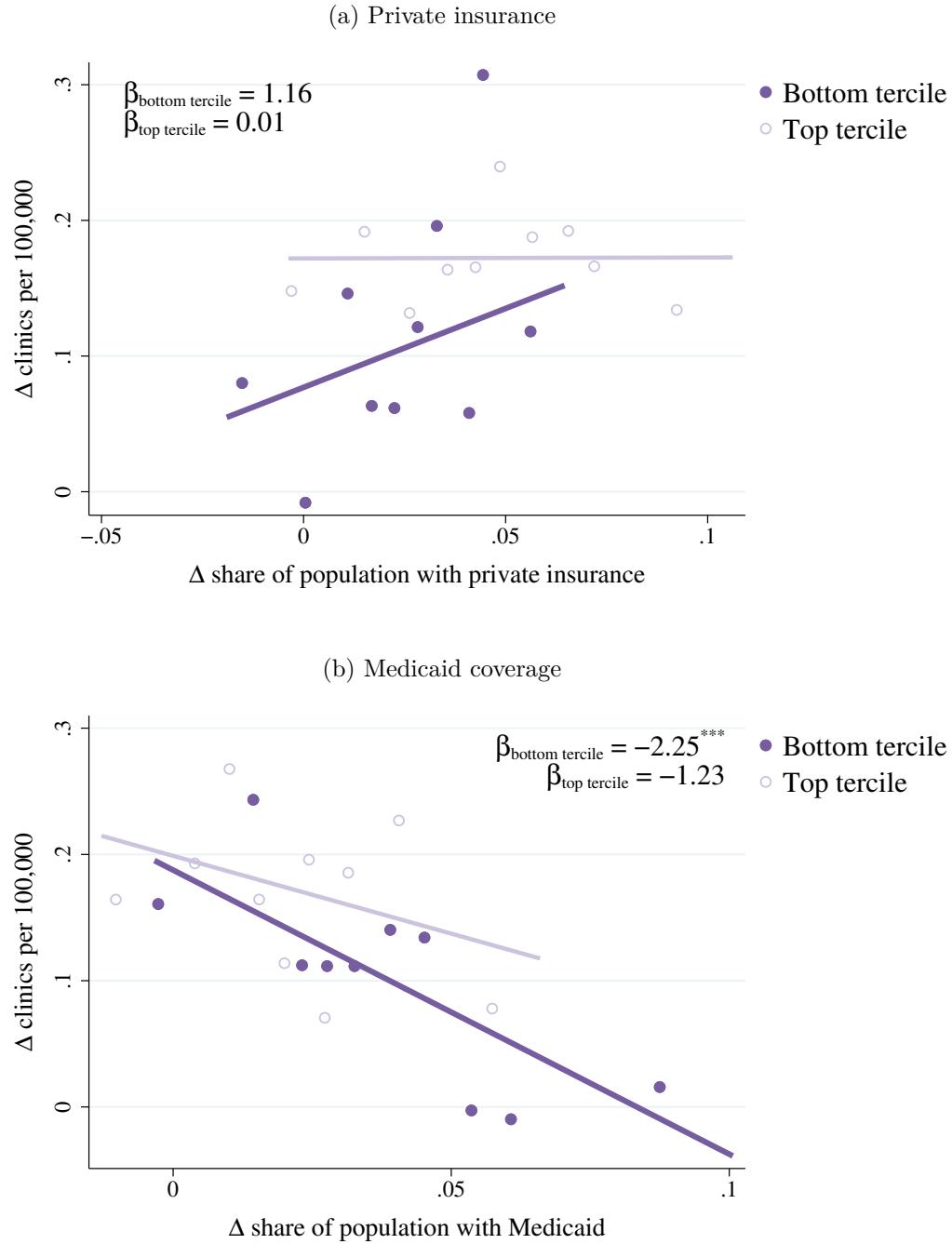
Notes: The above figures show how county-level changes in retail clinics per 100,000 people from 2013 to 2015 covary with county-level changes in the share of the population with health insurance of any type (subfigure (a)), private insurance coverage (subfigure (b)), and Medicaid coverage (subfigure (c)) over the same period. In subfigures (b) and (c), both the unconditional relationships (light lines, hollow dots) and the relationships conditional on changes in other types of health insurance (dark lines, solid dots) are shown. In all subfigures, counties are grouped into deciles accounting for approximately equal shares of the population based on the variable denoted on the x-axis. Private insurance includes employer-sponsored coverage and direct purchase. Data on retail clinics come from Merchant Medicine; data on health insurance come from the one-year ACS.

Figure 6: Effects of insurance on retail clinic penetration predicted using potential confounders



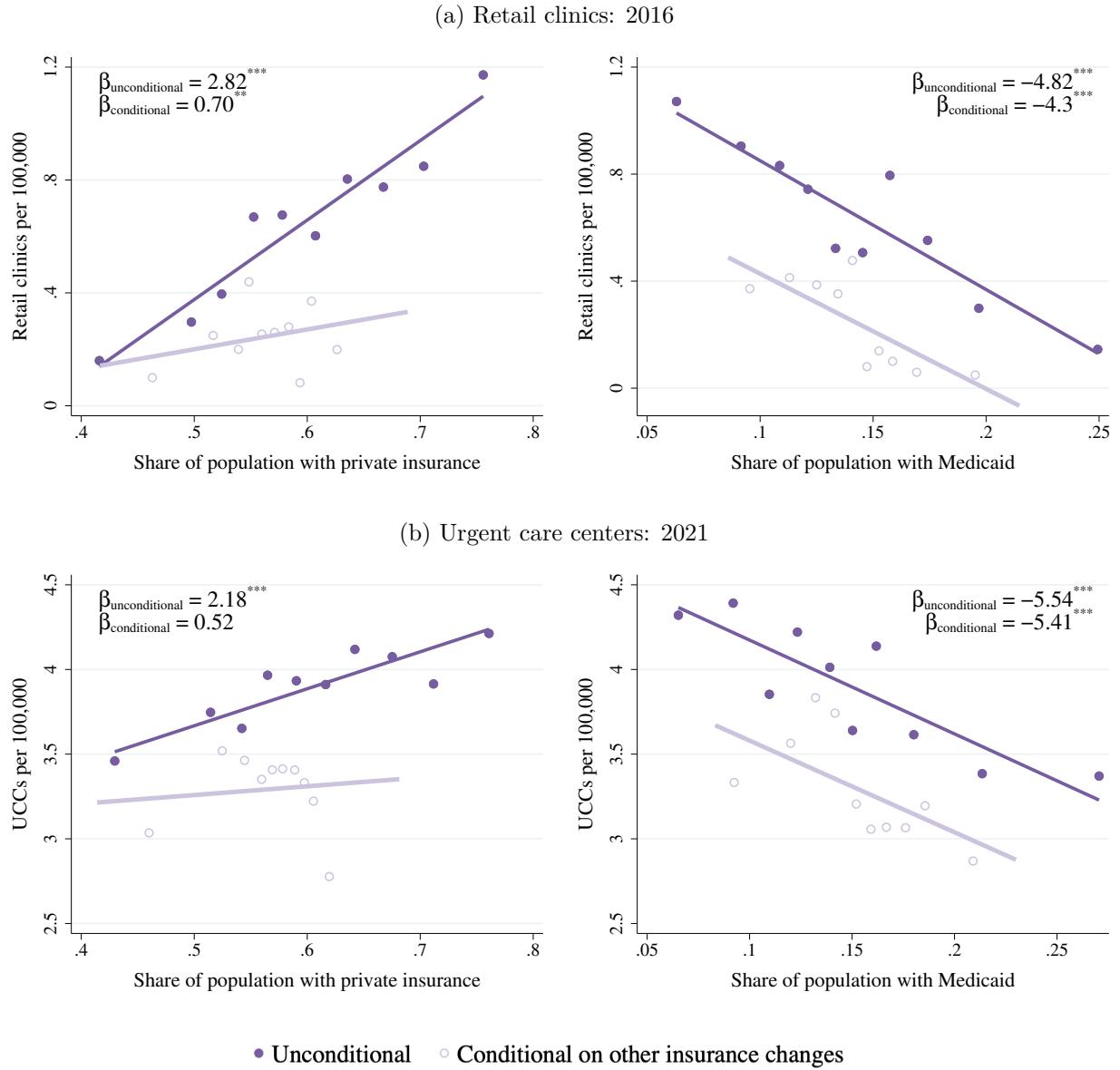
Notes: The above figure shows output from estimation of two-step balancing regressions. We first predict retail clinic concentration using local socio-demographics. We then estimate the specification denoted on the y-axis using this measure of predicted retail clinic penetration as the dependent variable. “Cross-section” refers to estimation of equation (2) without county fixed effects, “county FEs” refers to estimation of equation (2), and “2SLS” refers to estimation of equation (4). Since the outcome in these analyses is predicted using the vector of time-varying, county-level controls  $X_{ct}$ , we exclude these controls from equations (2) and (4) when estimating the second-stage regressions. The share of the population with private insurance and Medicaid coverage are always included in the same regression. Private insurance includes employer-sponsored coverage and direct purchase. Data come from the one-year ACS.

Figure 7: Changes in clinic presence versus changes in insurance by baseline Medicaid rates



Notes: The above figures show how county-level changes in retail clinics per 100,000 people from 2013 to 2015 covary with county-level changes in the share of the population with private insurance coverage (subfigure (a)) and Medicaid coverage (subfigure (b)) over the same period. These relationships are conditional on changes in other types of health insurance and are shown separately among counties in states with Medicaid reimbursement rates for office visits of low complexity (CPT 99201) in the bottom tercile (light lines, hollow dots) and the top tercile (dark lines, solid dots) across all states in 2010. In both subfigures, counties are grouped into deciles accounting for approximately equal shares of the population based on the variable denoted on the x-axis. Private insurance includes employer-sponsored coverage and direct purchase. Data on retail clinics come from Merchant Medicine, data on health insurance come from the one-year ACS, and data on Medicaid reimbursement rates come from [Alexander and Schnell \(2019\)](#).

Figure 8: Retail clinic and urgent care center presence by health insurance coverage



Notes: The above figures show how the county-level number of retail clinics per 100,000 people in 2016 (subfigure (a)) and urgent care centers per 100,000 people in 2021 (subfigure (b)) covary with the county-level share of the population with private insurance coverage (left subplot in each subfigure) and Medicaid coverage (right subplot in each subfigure) in 2012–2016 (subfigure (a)) and 2016–2020 (subfigure (b)). All subfigures show both the unconditional relationship (dark lines, solid dots) and the relationship conditional on the shares of the population with other types of health insurance (light lines, hollow dots). When conditioning on other insurance profiles, the share of the population that is uninsured is the omitted category. Counties are grouped into deciles accounting for approximately equal shares of the population based on the variable denoted on the x-axis. Private insurance includes employer-sponsored coverage and direct purchase. Data on retail clinics come from Merchant Medicine, data on urgent care centers come from the NUCR database, and data on health insurance come from the five-year ACS.

## IX Tables

Table 1: County-level summary statistics: retail clinics and socio-demographics

	Number of retail clinics in 2016		
	One or more (1)	None (2)	P-value (3)
<b>a. Retail clinics</b>			
<i>2016</i>			
Open clinics	5.51	0	
Clinics per 100,000	1.11	0	
<i>2010–2016</i>			
Openings	3.66	0.10	
Closings	1.39	0.32	
Share ever clinic	1.00	0.19	
<b>b. County characteristics (2016)</b>			
<i>Basic demographics</i>			
Total population	610,456	221,138	0.000
Population density (per sq. mile)	3,397	715	0.001
Share White	0.67	0.77	0.000
Share Black	0.15	0.10	0.000
Share Hispanic	0.21	0.17	0.022
Share under 18	0.23	0.23	0.101
Share aged 18–64	0.63	0.61	0.000
<i>Income and education</i>			
Median income	59,270	50,398	0.000
Share poverty	0.13	0.15	0.000
Share employed	0.62	0.58	0.000
Share high school	0.24	0.28	0.000
Share some college	0.28	0.31	0.000
Share college plus	0.36	0.28	0.000
<i>Health insurance</i>			
Share insured	0.92	0.92	0.605
Share private	0.63	0.59	0.000
Share Medicaid	0.16	0.17	0.007
Share Medicare	0.07	0.07	0.413
Expanded Medicaid by 2014	0.58	0.51	0.215
Number of counties	321	234	

Notes: The above table presents information on the concentration of retail clinics (panel (a)) and local socio-demographics and insurance status (panel (b)) at the county-year level. Column (1) provides averages across counties with one or more open retail clinics in 2016, column (2) provides averages across counties with no open retail clinics in the same year, and column (3) provides *p*-values showing whether the values in columns (1) and (2) are statistically different. Only counties with data available in every one-year ACS from 2010 to 2016 are included; counties must have a population of 65,000 or more to be included in the one-year ACS. Data on retail clinics come from Merchant Medicine; data on county-level characteristics come from the one-year ACS. Refer to Table A1 for analogous statistics across all counties using data from the five-year ACS.

Table 2: Changes in insurance and retail clinic penetration: OLS

	Retail clinics per 100,000		
	(1)	(2)	(3)
Share insurance	0.242 (0.412)		
Share private		0.979** (0.453)	1.214** (0.488)
$\times \mathbb{1}\{\Delta Medicaid > median\}$			-0.938 (0.599)
Share Medicaid		-1.279*** (0.448)	0.540 (0.766)
$\times \mathbb{1}\{\Delta Medicaid > median\}$			-1.792** (0.759)
County fixed effects	X	X	X
Year fixed effects	X	X	X
Demographic controls	X	X	X
Observations	3,869	3,869	3,869
$R^2$	0.898	0.900	0.901
Mean dependent variable	0.612	0.612	0.612
Share private $\times (1 + \mathbb{1}\{\Delta Medicaid > med.\})$			0.276 (0.570)
Share Medicaid $\times (1 + \mathbb{1}\{\Delta Medicaid > med.\})$			-1.252*** (0.451)

Notes: The above table shows the association between retail clinics per 100,000 people and the share of the population with different types of health insurance from estimation of equations (1) and (2). Observations are at the county-year level from 2010 to 2016. All specifications include county fixed effects, year fixed effects, and all time-varying socio-demographic controls listed in Table 1. Standard errors are clustered by county.

Table 3: Effects of insurance on retail clinic penetration: 2SLS

Dependent variable:	Any health insurance		By insurance type	
	Share insured (1)	Retail clinics per 100,000 (2)	Share private (3)	Share Medicaid (4)
<b>a. First stage</b>				
$Post_t \times Employee_c^{2013}$	0.132*** (0.027)		0.198*** (0.022)	-0.056*** (0.017)
$Post_t \times [138 - 400\% FPL]_c^{2013}$	0.545*** (0.086)		0.369*** (0.066)	0.089 (0.071)
$Post_t \times Expansion_s \times [ < 138\% FPL]_c^{2013}$	0.296*** (0.111)		-0.035 (0.083)	0.321*** (0.050)
<b>b. Two-stage least squares</b>				
Share insurance		0.927 (1.185)		
Share private				3.142** (1.303)
Share Medicaid				-5.085** (2.062)
County fixed effects	X	X	X	X
Year fixed effects	X	X	X	X
Demographic controls	X	X	X	X
Observations	3,869	3,869	3,869	3,869
$R^2$	0.950		0.981	0.964
Mean dependent variable	0.885	0.612	0.613	0.145
First stage F-stat		160.5		112.3

Notes: The above table shows the effects of the share of the population with different types of health insurance on the number of retail clinics per 100,000 people from estimation of equation (4) (panel (b)). First-stage estimates showing the relationship between our instruments and the share of the population with different types of health insurance from estimation of equation (3) are provided in panel (a). Observations are at the county-year level from 2010 to 2016. All specifications include county fixed effects, year fixed effects, and all time-varying socio-demographic controls listed in Table 1. Standard errors are clustered by county. Cragg-Donald Wald F-statistics are reported.

Table 4: Effects of insurance on retail clinic penetration: openings versus closings

	Any insurance		By type	
	OLS (1)	2SLS (2)	OLS (3)	2SLS (4)
<b>a. Openings per 100,000</b>				
Share insurance	0.156 (0.194)	2.227 (1.474)		
Share private			0.316 (0.211)	3.067* (1.799)
Share Medicaid			-0.120 (0.238)	-0.335 (2.665)
<b>b. Closings per 100,000</b>				
Share insurance	-0.072 (0.118)	-0.567 (0.470)		
Share private			-0.163 (0.131)	-1.286** (0.606)
Share Medicaid			0.112 (0.142)	1.908* (1.044)
County fixed effects	X	X	X	X
Year fixed effects	X	X	X	X
Demographic controls	X	X	X	X
Observations	3,313	3,313	3,313	3,313
$R^2$ : openings	0.059		0.060	
$R^2$ : closings	0.089		0.090	
Mean openings	0.073	0.073	0.073	0.073
Mean closings	0.032	0.032	0.032	0.032
First stage F-stat		36.7		8.5

Notes: The above table shows the effects of the share of the population with different types of health insurance on the number of retail clinic entries (panel (a)) and exits (panel (b)) per 100,000 people from estimation of equation (A1) (columns (1) and (3)) and equation (A3) (columns (2) and (4)). Table A4 provides first-stage estimates showing the relationship between our instruments and the first difference of the share of the population with different types of health insurance from estimation of equation (A2). Observations are at the county-year level from 2010 to 2016. All specifications include year fixed effects and the first difference of all time-varying socio-demographic controls listed in Table 1. These specifications further control for the number of retail clinics per 100,000 people in the previous period to account for the fact that openings (closings) are less (more) common in markets with many retail clinics. Standard errors are clustered by county. Cragg-Donald Wald F-statistics are reported.

# Online Appendix

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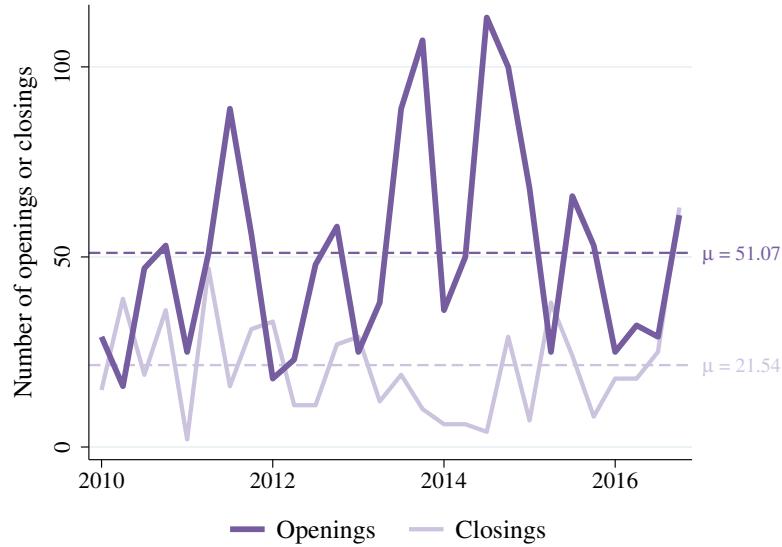
The Expansionary and Contractionary  
Supply-Side Effects of Health Insurance

*Geddes and Schnell (2023)*

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## A Supplementary figures

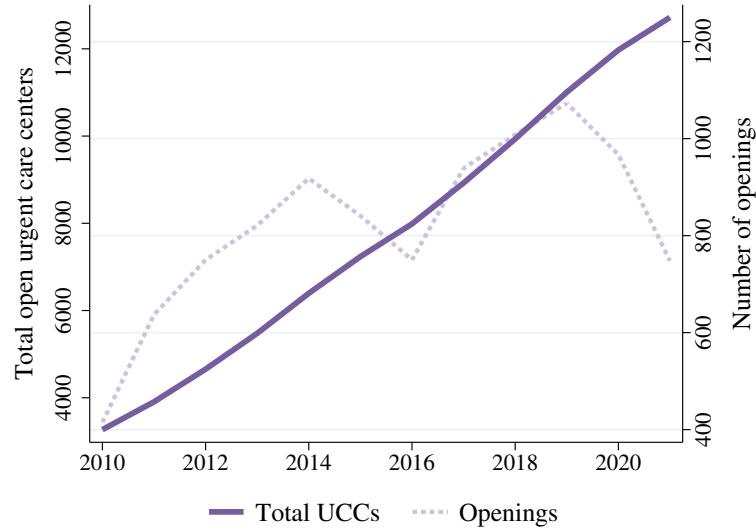
Figure A1: Retail clinic openings and closings: 2010–2016



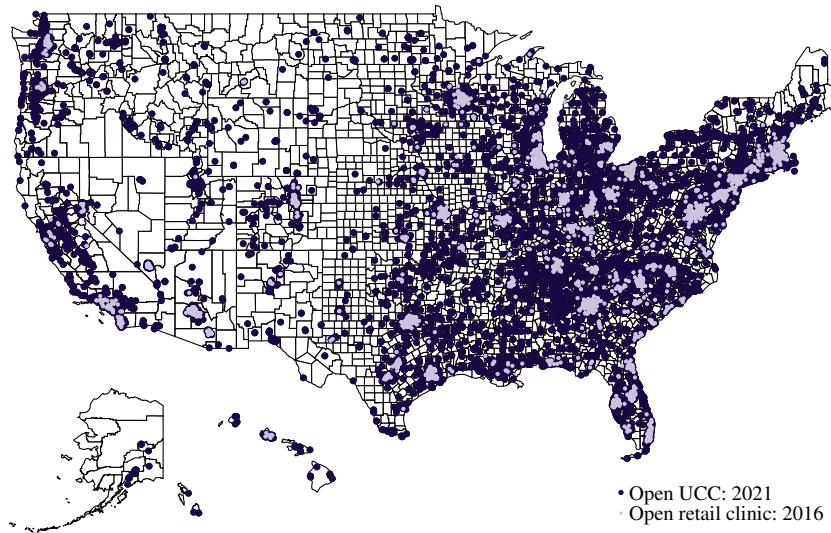
Notes: The above figure shows the number of retail clinic openings (dark, thick line) and closings (light, thin line) quarterly from 2010 to 2016. The dashed horizontal lines denote the quarterly averages of each measure over the entire sample period. Refer to Figure 1 for the total number of open retail clinics and net entry (the number of openings net of the number of closings in a given quarter) over the same period. Data come from Merchant Medicine.

Figure A2: Urgent care centers across the United States

(a) Total open and annual entry: 2010–2021

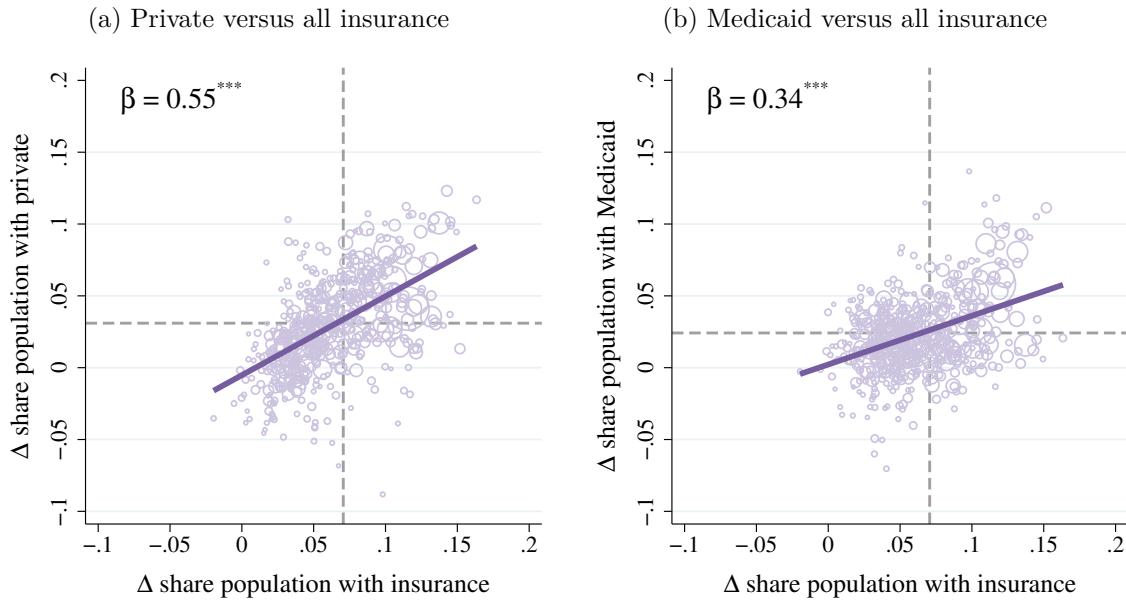


(b) Locations: 2021



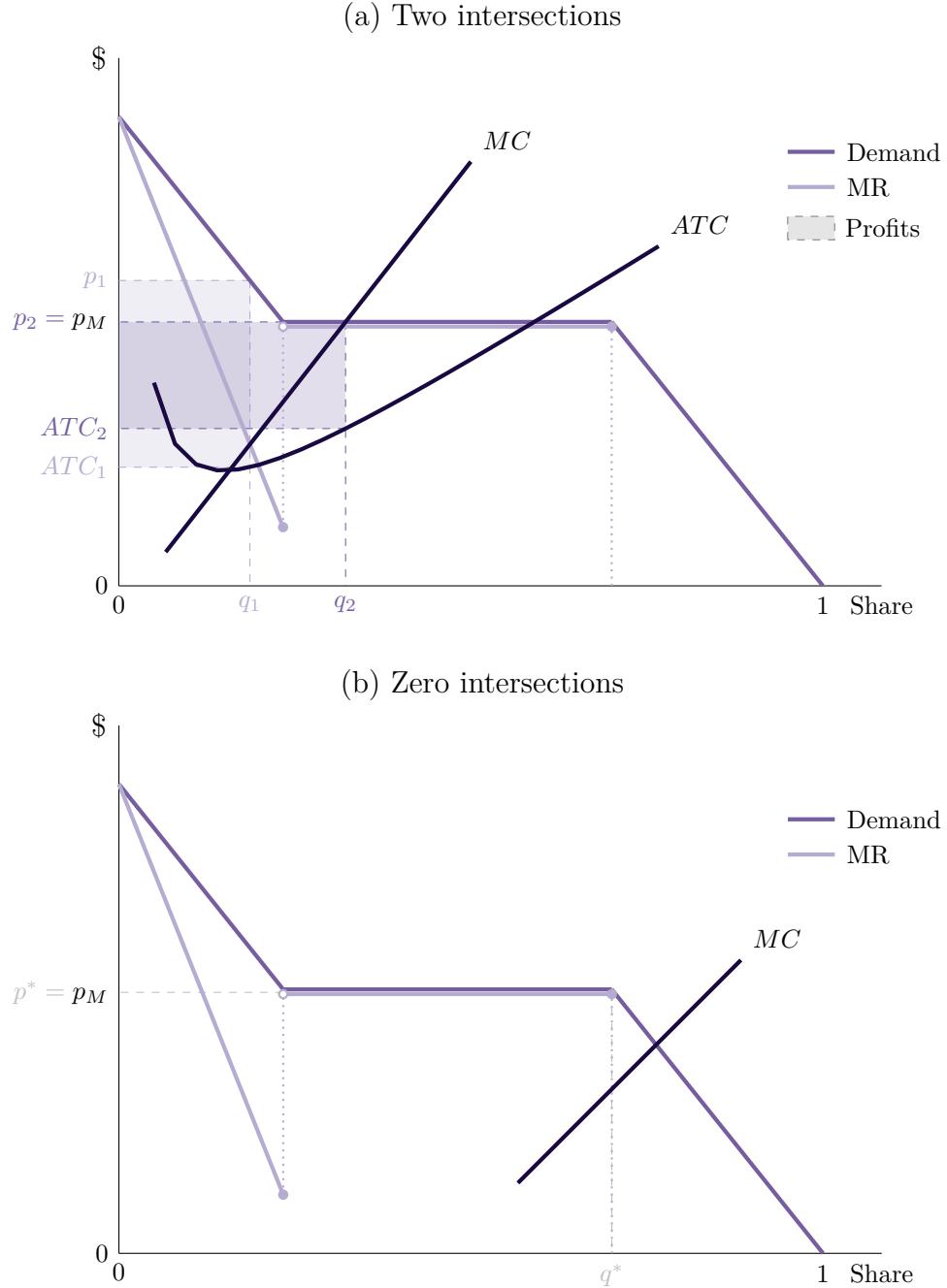
Notes: The above figures show the number and locations of urgent care centers across the United States. Subfigure (a) shows the total number of open urgent care centers (dark, solid line) and annual entry conditional on survival to 2021 (light, dashed line) from 2010 to 2021. Subfigure (b) shows the locations of urgent care centers in 2021 (geo-coded, dark dots) and retail clinics in 2016 (geo-coded, light dots). Data on urgent care centers come from the NUCR database; data on retail clinics come from Merchant Medicine.

Figure A3: Correlation between changes in insurance types: 2013–2015



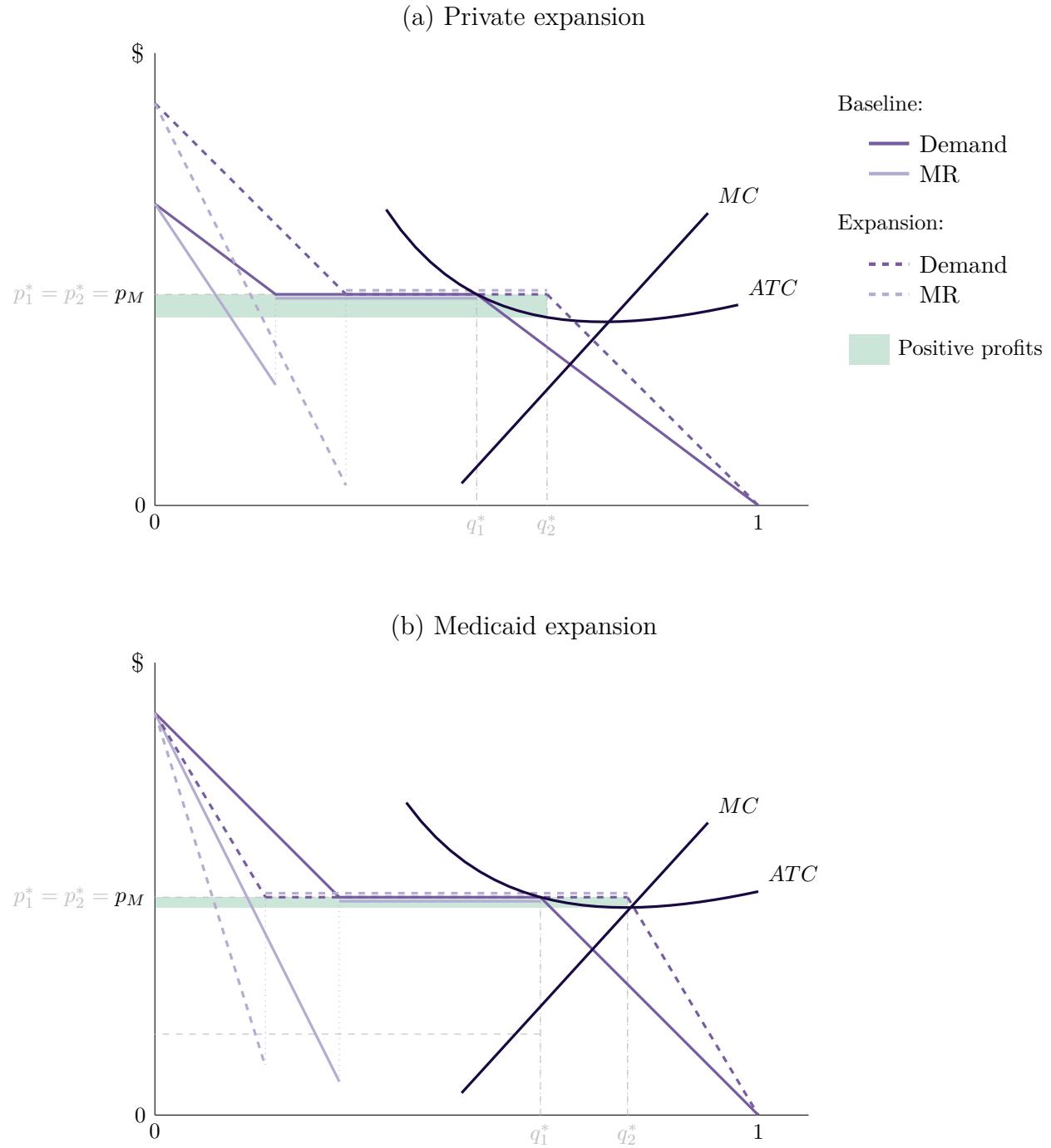
Notes: The above figures show how county-level changes from 2013 to 2015 in the share of the population with private insurance (subfigure (a)) and the share of the population covered by Medicaid (subfigure (b)) covary with county-level changes in the share of population with health insurance coverage of any type over the same period. The size of the markers denotes county-level population in 2010; the solid line denotes the best fit line. Private insurance includes employer-sponsored coverage and direct purchase. Data come from the one-year ACS. Refer to Figure 2(c) for county-level changes in the share of the population with Medicaid coverage versus county-level changes in the share of the population with private insurance over the same period.

Figure A4: Additional solutions to firm's baseline problem



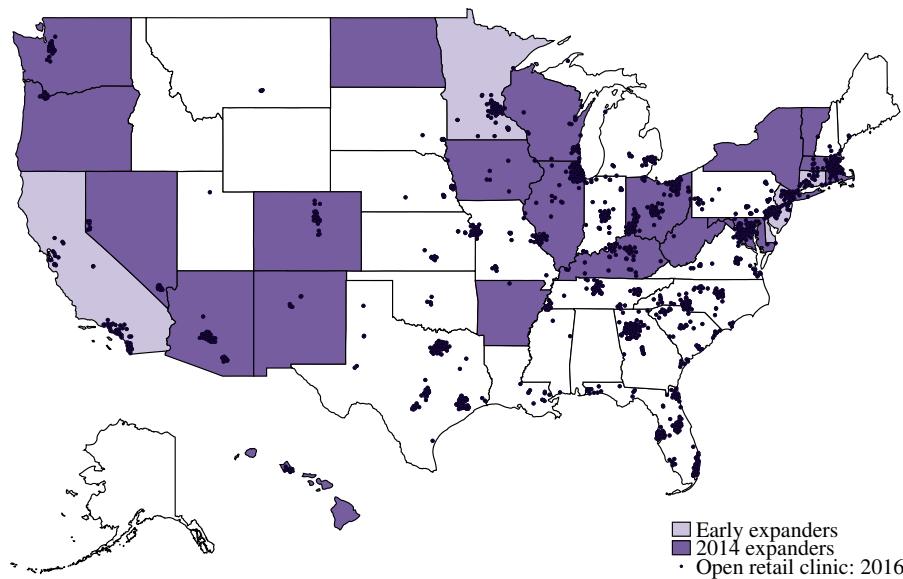
Notes: The above figures show how prices and quantities are determined when marginal revenue and marginal costs intersect twice (subfigure (a)) or not at all (subfigure (b)). As outlined in Figure 3, the total demand curve (dark purple line) and associated marginal revenue curve (light purple line) in each subfigure are for a firm in a market with both administered and market-based prices. The demand curve is perfectly elastic at the administered Medicaid price ( $p_M$ ) with length equivalent to the share of the population covered by Medicaid. As shown in subfigure (a), the firm must consider average total costs to compare profits at each potential set of prices and quantities when there are two intersections between marginal revenue and marginal cost. As shown in subfigure (b), the firm sets  $p^* = p_M$  and sees all patients willing to pay at least  $p_M$  (i.e., the firm does not need to restrict capacity) when the marginal cost curve lies entirely below the positive portion of the marginal revenue curve.

Figure A5: Effects of expansions when no intersection between marginal revenue and costs



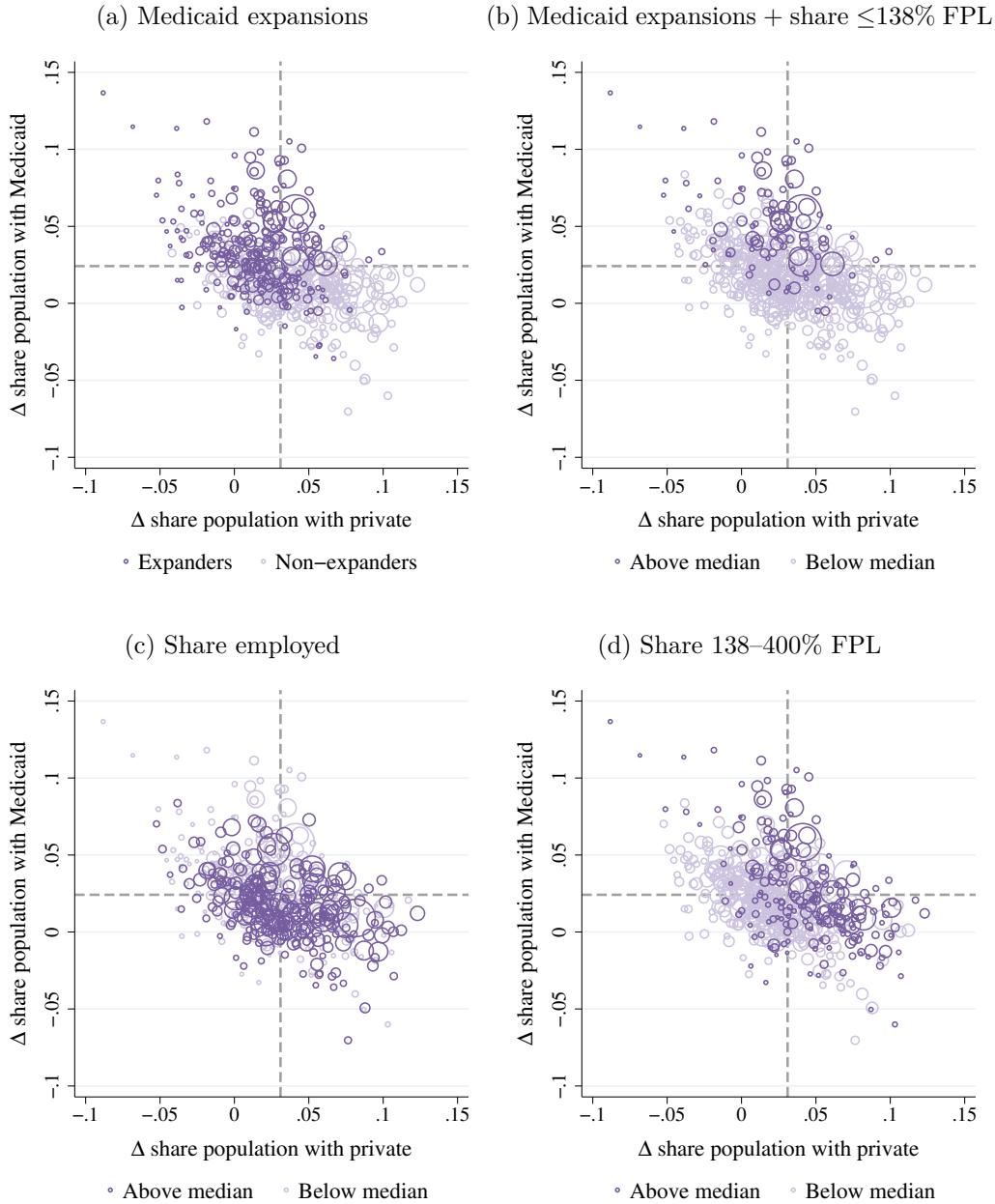
Notes: The above figures show how expansions of private insurance (subfigure (a)) and Medicaid coverage (subfigure (b)) affect firm profits when there is no intersection between marginal costs and marginal revenue at baseline. As shown in the subfigures, both private and Medicaid expansions tend to increase firm profits when the marginal cost curve lies entirely below the positive portion of the marginal revenue curve at baseline.

Figure A6: Medicaid expansion status and retail clinic locations



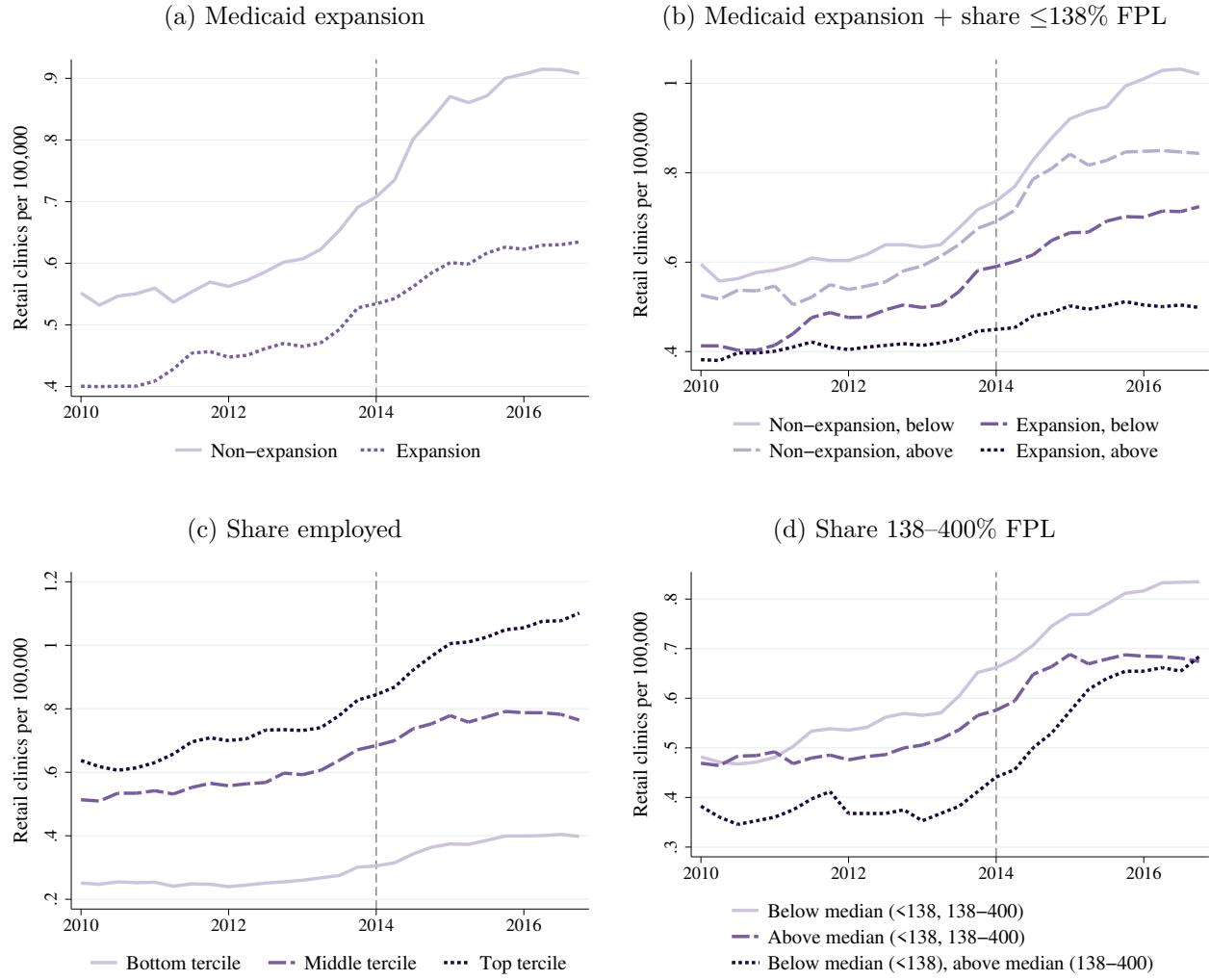
Notes: The above figure shows the locations of open retail clinics in 2016 (geo-coded dots) and state-level Medicaid expansions by 2014 (shaded states) across the United States. Data on retail clinics come from Merchant Medicine; data on Medicaid expansions come from the Kaiser Family Foundation.

Figure A7: Changes in insurance types (2013–2015) by instrument components



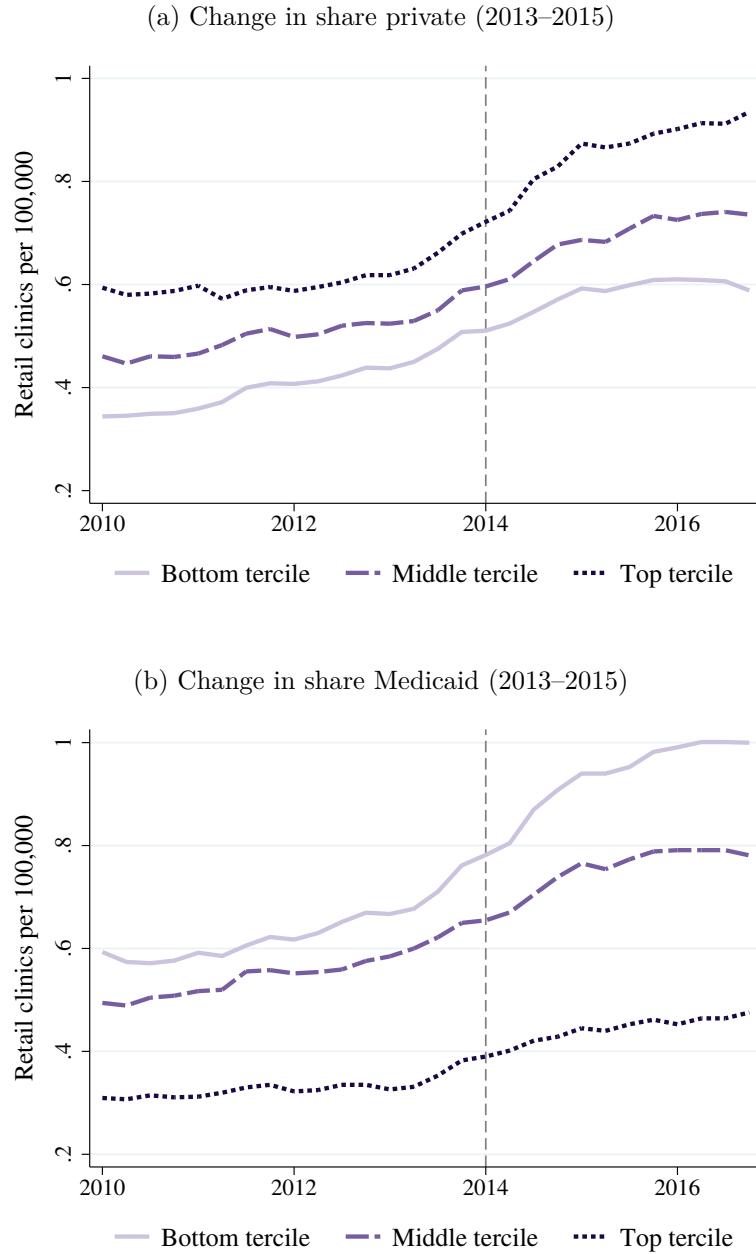
Notes: The above figures show how the different instrument components isolate county-level changes in the share of the population covered by Medicaid (y-axis) and private insurance (x-axis) from 2013 to 2015. In all subfigures, the size of the markers denotes county-level population in 2010, and the dashed lines denote the population-weighted median of changes in Medicaid coverage and private insurance from 2013 to 2015. In subfigure (a), the dark (light) circles denote counties in states that expanded (did not expand) Medicaid by 2014. In subfigure (b), the dark circles denote counties that both had an above-median share of the population under 138 percent of the federal poverty level (FPL) in 2013 and are in states that expanded Medicaid by 2014. In subfigure (c), the dark (light) circles denote counties with an above-median (below-median) share of the population employed in 2013. In subfigure (d), the dark (light) circles denote counties with an above-median (below-median) share of the population between 138 and 400 percent of the FPL in 2013. Private insurance includes employer-sponsored coverage and direct purchase. Data on retail clinics come from Merchant Medicine; data on health insurance come from the one-year ACS.

Figure A8: Retail clinic penetration by instrument components: 2010–2016



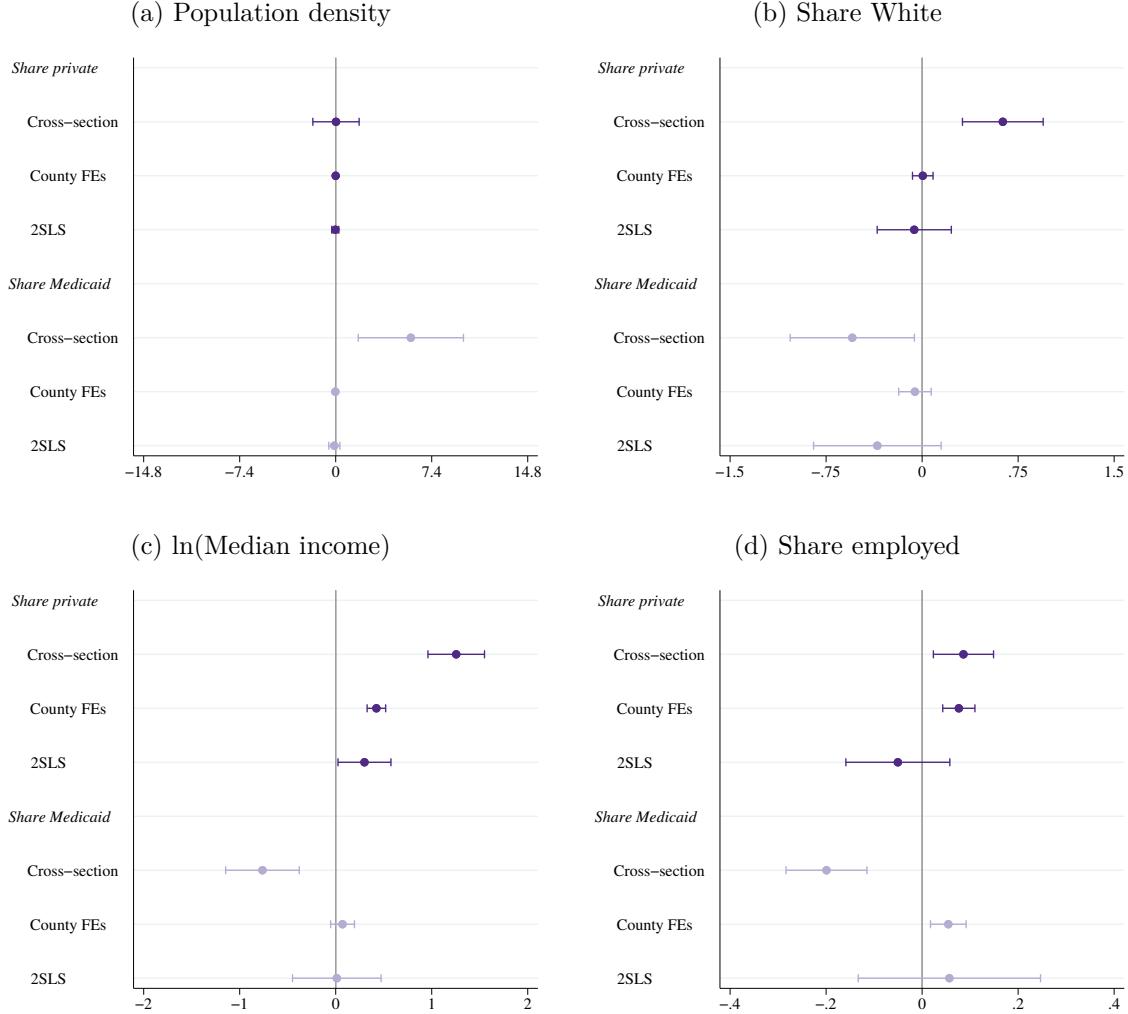
Notes: The above figures show the population-weighted average number of retail clinics per 100,000 people at the county-quarter level from 2010 to 2016 by different instrument components. In subfigure (a), the dark, dotted (light, solid) line considers counties in states that expanded (did not expand) Medicaid by 2014. In subfigure (b), the dark, dotted line considers counties that both had an above-median share of the population under 138 percent of the federal poverty level (FPL) in 2013 and are in states that expanded Medicaid by 2014; the light, solid line denotes counties that had a below-median share of the population under 138 percent of the FPL in 2013 and are in states that did not expand Medicaid by 2014. In subfigure (c), the dark, dotted (light, solid) line considers counties with a top-tercile (bottom-tercile) share of the population employed in 2013. In subfigure (d), the medium, dashed (light, solid) line considers counties with both an above-median (below-median) share of the population under 138 percent of the FPL and between 138 and 400 percent of the FPL in 2013; the dark, dotted line denotes counties with a below-median share of the population under 138 percent of the FPL and an above-median share between 138 and 400 percent of the FPL in 2013. Data on retail clinics come from Merchant Medicine.

Figure A9: Retail clinic penetration by changes in insurance: 2010–2016



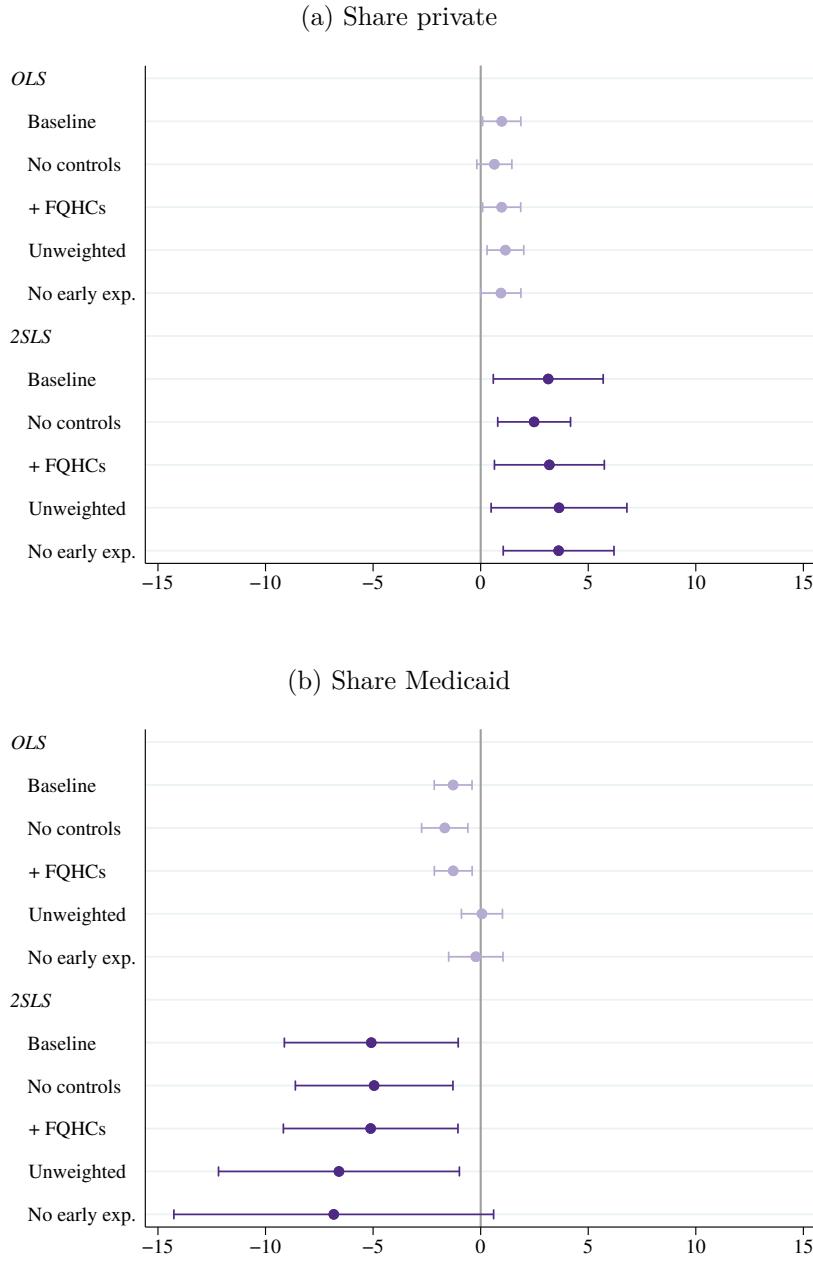
Notes: The above figures show the population-weighted average number of retail clinics per 100,000 people at the county-quarter level from 2010 to 2016 by changes in the share of the population covered by private insurance (subfigure (a)) and Medicaid (subfigure (b)) from 2013 to 2015. In subfigure (a), the dark, dotted (light, solid) line considers counties with a top-tercile (bottom-tercile) change in the share of the population with private insurance coverage from 2013 to 2015. In subfigure (b), the dark, dotted (light, solid) line considers counties with a top-tercile (bottom-tercile) change in the share of the population with Medicaid coverage from 2013 to 2015. Data on retail clinics come from Merchant Medicine; data on health insurance come from the one-year ACS.

Figure A10: Balancing regressions on select potential confounders



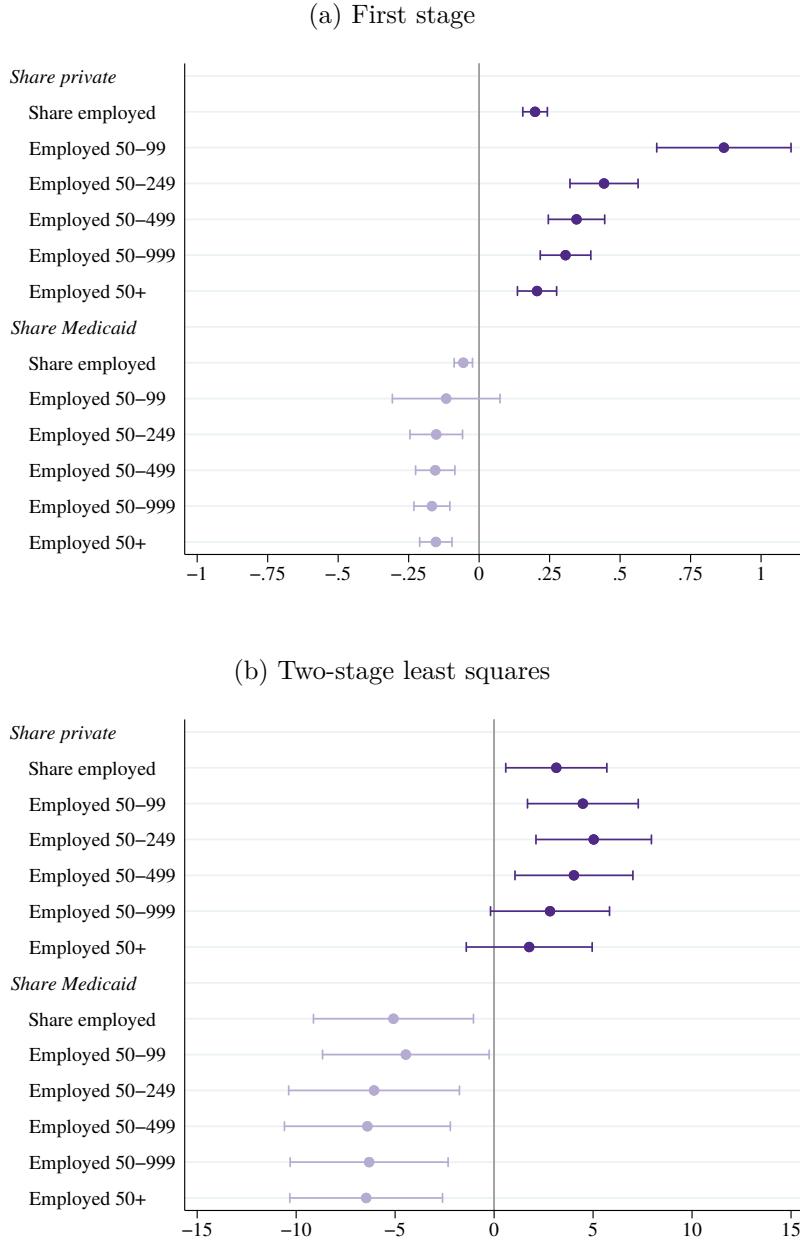
Notes: The above figures show output from estimation of the specifications denoted on the y-axis with different potential confounders as the dependent variable. “Cross-section” refers to estimation of equation (2) without county fixed effects, “county FEs” refers to estimation of equation (2), and “2SLS” refers to estimation of equation (4). The share of the population with private insurance and Medicaid coverage are always included in the same regression. Private insurance includes employer-sponsored coverage and direct purchase. Data come from the one-year ACS.

Figure A11: Effects of insurance on retail clinic penetration: robustness



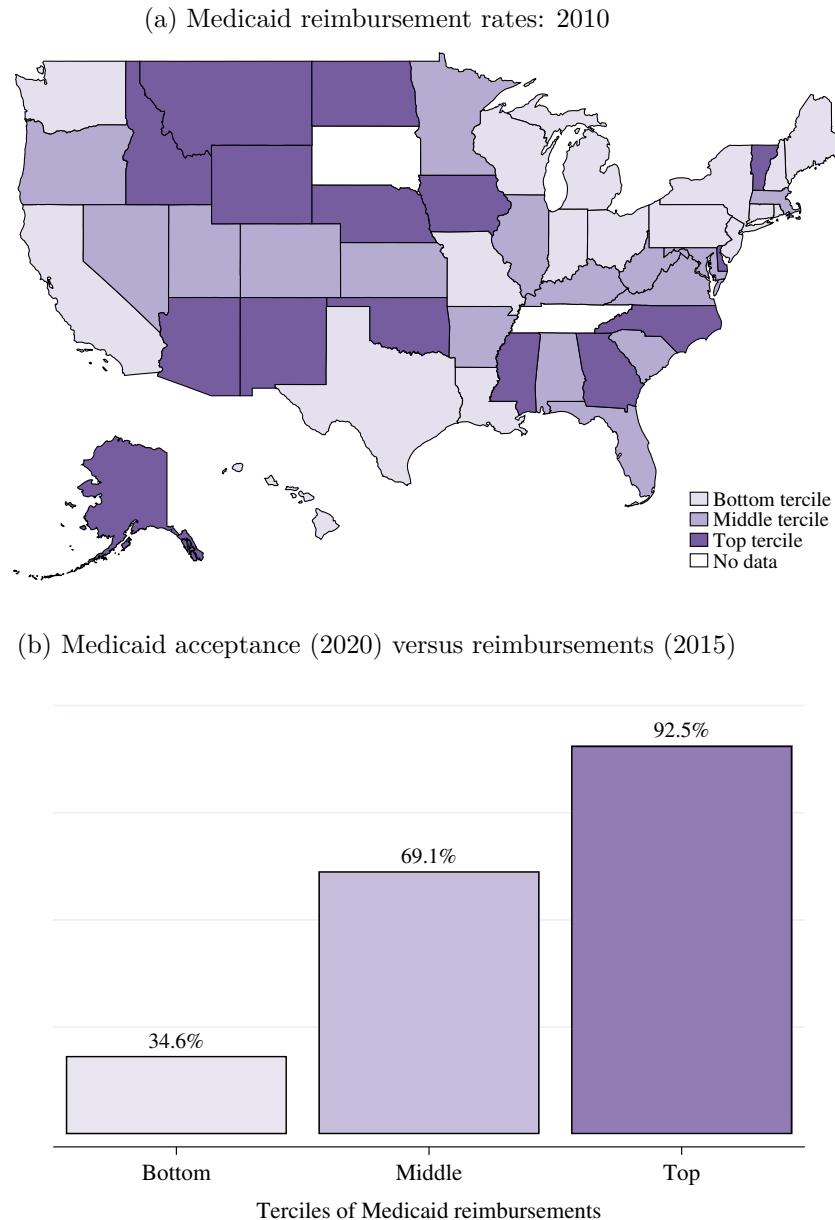
Notes: The above figures show the sensitivity of our baseline estimates to alternative empirical specifications. The top panel of each subfigure (“OLS”) shows output from estimation of equation (2), and the bottom panel (“2SLS”) shows output from estimation of equation (4). As outlined in these equations, the outcome variable is retail clinics per 100,000 people at the county-year level, and the share of the population with private insurance and Medicaid coverage are always included in the same regression. Each row displays results from an alternative specification or sample: “Baseline” refers to our baseline estimates first displayed in column (2) of Table 2 and column (5) of Table 3; “No controls” refers to specifications excluding all time-varying, county-level controls; “Unweighted” refers to specifications in which observations are not weighted by county population in 2010; and “No early exp.” refers to specifications that drop counties in the five states that expanded Medicaid before 2014 from the sample. Private insurance includes employer-sponsored coverage and direct purchase. Data on retail clinics come from Merchant Medicine; data on health insurance come from the one-year ACS.

Figure A12: Effects of insurance on retail clinic penetration: alternative employment shares



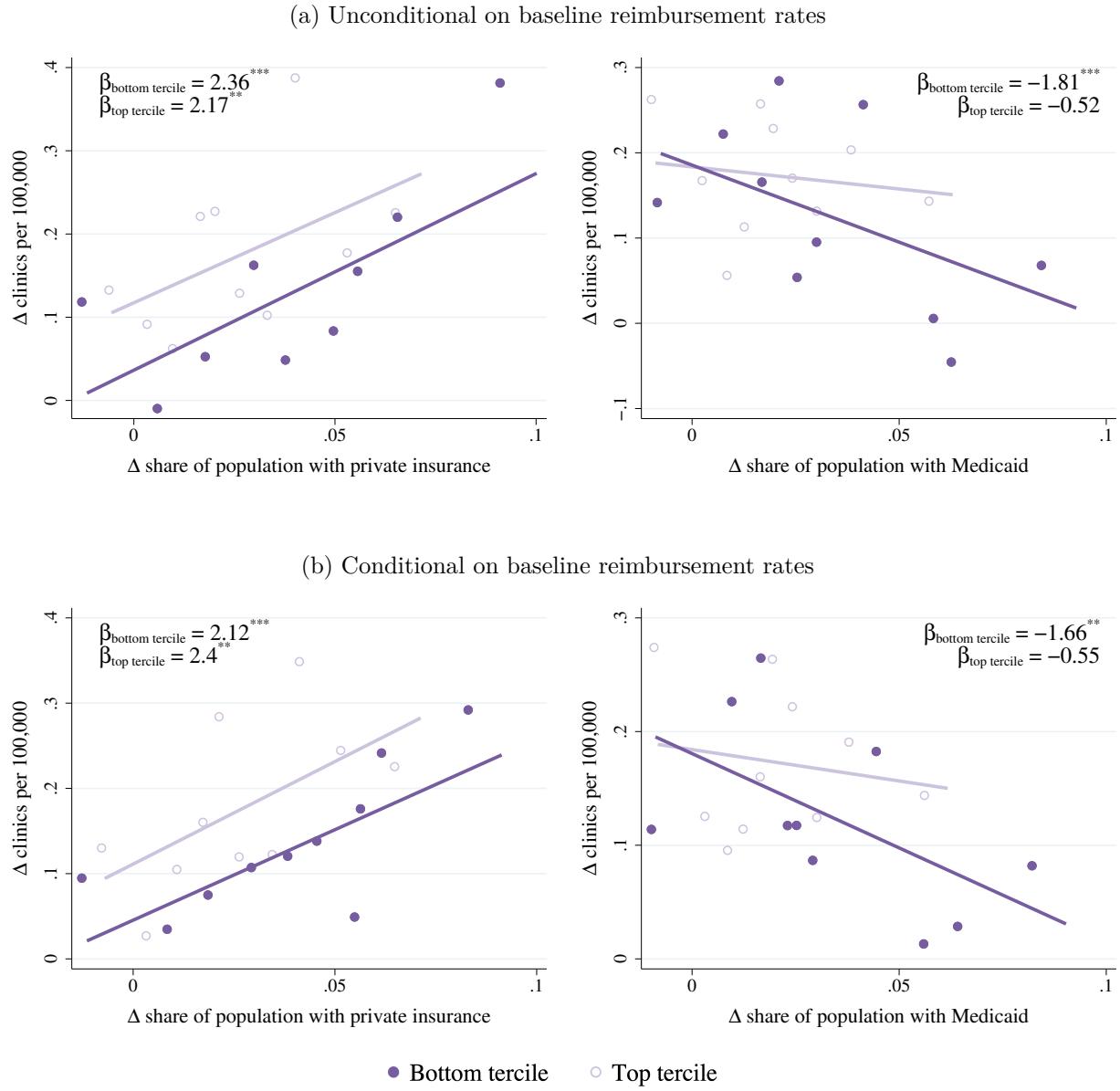
Notes: The above figures show the sensitivity of our baseline estimates to using alternative employment shares when constructing the employment instrument. Subfigure (a) shows output from estimation of equation (3); the outcome variable is either the share of the population with private insurance (top panel) or the share of the population with Medicaid coverage (bottom panel) at the county-year level. Subfigure (b) shows output from estimation of equation (4); the outcome variable is retail clinics per 100,000 people at the county-year level, and the shares of the population with private insurance and Medicaid coverage are always included in the same regression. Each row displays results using an alternative measure of the share employed when constructing the employment instrument: “Share employed” refers to our baseline estimates first displayed in columns (3)–(5) of Table 3 and uses the overall employment rate in 2013, and the remaining rows instead use the share of the population employed in firms with 50–99, 50–249, 50–499, 50–999, or 50+ employees in 2013. We approximate the shares of the population employed by firms of different sizes by multiplying county-level employment shares from the one-year ACS by state-level shares of employees working in establishments of different sizes from the QCEW.

Figure A13: Medicaid reimbursement rates and coverage acceptance



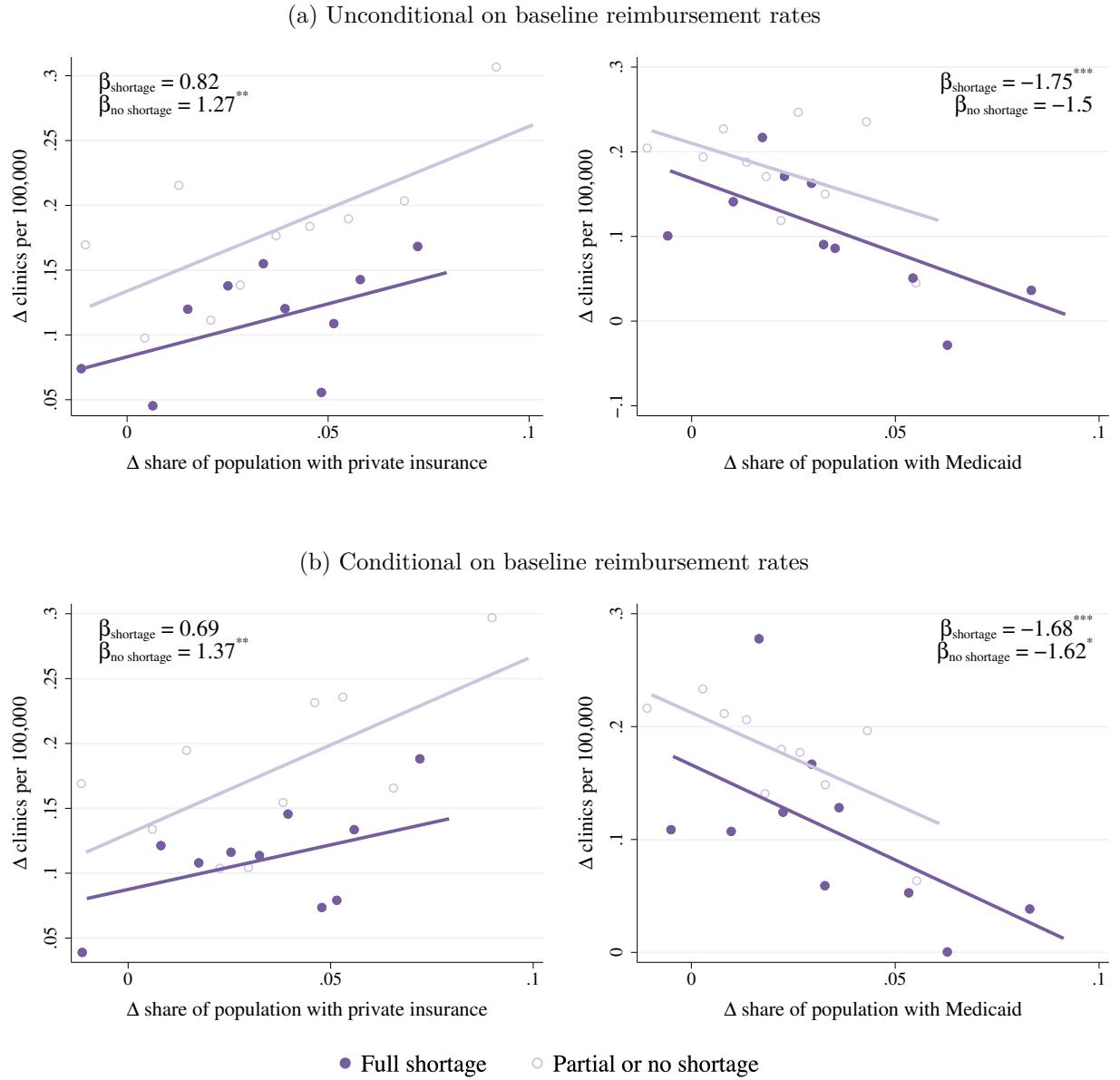
Notes: The above figures show Medicaid reimbursement rates and coverage acceptance by retail clinics across the United States. Subfigure (a) shows state-level Medicaid reimbursement rates for an office visit of low complexity (CPT 99201) in 2010. Subfigure (b) shows the share of states in which CVS MinuteClinics accepted at least one form of Medicaid coverage in 2020 within each tercile of Medicaid reimbursements rates in 2015 (the latest year of data available for Medicaid reimbursements). Only the 34 states with CVS MinuteClinics are considered in subfigure (b). Data on Medicaid reimbursement rates come from [Alexander and Schnell \(2019\)](#), and data on Medicaid acceptance by CVS MinuteClinics was collected by the authors as outlined in footnote 32.

Figure A14: Effects by primary care providers per capita at baseline



Notes: The above figures show how county-level changes in retail clinics per 100,000 people from 2013 to 2015 covary with county-level changes in the share of the population with private insurance coverage (left subfigures) and Medicaid coverage (right subfigures) over the same period. These relationships are shown separately across terciles of the number of primary care providers per capita in 2010; primary care providers include physicians in primary care and nurse practitioners. All subfigures are conditional on county-level changes in other types of health insurance; subplots in subfigure (b) are further conditional on state-level Medicaid reimbursement rates for office visits of low complexity (CPT 99201) in 2010. Counties are grouped into deciles accounting for approximately equal shares of the population based on the variable denoted on the x-axis. Private insurance includes employer-sponsored coverage and direct purchase. Data on retail clinics come from Merchant Medicine, data on health insurance come from the one-year ACS, data on the number of primary care providers per capita come from the HRSA's Area Health Resource Files, and data on Medicaid reimbursement rates come from [Alexander and Schnell \(2019\)](#).

Figure A15: Effects by baseline primary care shortage designations



Notes: The above figures show how county-level changes in retail clinics per 100,000 people from 2013 to 2015 covary with county-level changes in the share of the population with private insurance coverage (left subfigures) and Medicaid coverage (right subfigures) over the same period. These relationships are shown separately among counties that are and are not designated “primary care shortage areas” by the HRSA in 2010. All subfigures are conditional on county-level changes in other types of health insurance; subplots in subfigure (b) are further conditional on state-level Medicaid reimbursement rates for office visits of low complexity (CPT 99201) in 2010. Counties are grouped into deciles accounting for approximately equal shares of the population based on the variable denoted on the x-axis. Private insurance includes employer-sponsored coverage and direct purchase. Data on retail clinics come from Merchant Medicine, data on health insurance come from the one-year ACS, data on primary care shortage areas come from the HRSA’s Area Health Resource Files, and data on Medicaid reimbursement rates come from Alexander and Schnell (2019).

## B Supplementary tables

Table A1: County-level summary statistics by availability in one-year ACS

Retail clinics in 2016:	In one-year ACS		Not in one-year ACS	
	One+ (1)	None (2)	One+ (3)	None (4)
<b>a. Retail clinics</b>				
<i>2016</i>				
Open clinics	5.51	0	1.60	0
Clinics per 100,000	1.13	0	2.39	0
<i>2010–2016</i>				
Openings	3.66	0.10	1.17	0.01
Closings	1.39	0.32	0.16	0.02
Share ever clinic	1.00	0.19	1.00	0.02
<b>b. County characteristics (2012–2016)</b>				
<i>Basic demographics</i>				
Total population	599,659	218,500	81,073	25,371
Population density (per sq. mile)	3,382	699	312	100
Share White	0.68	0.78	0.85	0.84
Share Black	0.15	0.10	0.09	0.09
Share Hispanic	0.21	0.18	0.08	0.08
Share under 18	0.23	0.24	0.24	0.23
Share aged 18–64	0.63	0.62	0.61	0.60
<i>Income and education</i>				
Median income	56,676	48,253	53,306	42,463
Share poverty	0.14	0.17	0.13	0.17
Share employed	0.61	0.57	0.59	0.54
Share high school	0.25	0.28	0.30	0.35
Share some college	0.28	0.31	0.30	0.30
Share college plus	0.35	0.27	0.28	0.20
<i>Health insurance</i>				
Share insured	0.88	0.89	0.90	0.88
Share private	0.62	0.58	0.65	0.56
Share Medicaid	0.14	0.16	0.12	0.15
Share Medicare	0.06	0.07	0.06	0.08
Expanded Medicaid by 2014	0.58	0.51	0.35	0.37
Number of counties	321	234	167	2,420

Notes: The above table presents information on the concentration of retail clinics from Merchant Medicine (panel (a)) and local socio-demographics and insurance status from the 2012–2016 five-year ACS (panel (b)) for counties that are in the one-year ACS (columns (1)–(2)) and counties that are not in the single-year files (columns (3)–(4)). Counties must have a population of 65,000 or more to be included in the one-year ACS. Columns (1) and (3) provide averages across counties with one or more open retail clinics in 2016, and columns (2) and (4) provide averages across counties with no open retail clinics in the same year.

Table A2: County-level summary statistics by urgent care center presence

	Number of UCCs in 2021		
	One or more (1)	None (2)	P-value (3)
<b>a. On-demand health care clinics</b>			
<i>Urgent care centers (2021)</i>			
Open clinics	6.81	0	
Clinics per 100,000	5.54	0	
<i>Retail clinics (2016)</i>			
Share any clinic	0.25	0.01	
Open clinics	1.08	0.02	
Clinics per 100,000	0.36	0.05	
<b>b. County characteristics (2016–2020)</b>			
<i>Basic demographics</i>			
Total population	166,336	12,313	0.000
Population density (per sq. mile)	2,277	120	0.000
Share White	0.70	0.82	0.000
Share Black	0.13	0.10	0.008
Share Hispanic	0.19	0.09	0.000
Share under 18	0.22	0.22	0.061
Share aged 18–64	0.62	0.58	0.000
<i>Income and education</i>			
Median income	57,966	43,045	0.000
Share poverty	0.13	0.16	0.000
Share employed	0.60	0.53	0.000
Share high school	0.26	0.36	0.000
Share some college	0.29	0.31	0.000
Share college plus	0.34	0.19	0.000
<i>Health insurance</i>			
Share insured	0.91	0.90	0.000
Share private	0.62	0.55	0.000
Share Medicaid	0.15	0.17	0.000
Share Medicare	0.08	0.10	0.000
Expanded Medicaid by 2014	0.52	0.37	0.000
Number of counties	1,869	1,274	

Notes: The above table presents information on the concentration of on-demand health care clinics (panel (a)) and local socio-demographics and insurance status (panel (b)) at the county-year level. Column (1) provides averages across counties with one or more open urgent care centers in 2021, column (2) provides averages across counties with no open urgent care centers in the same year, and column (3) provides *p*-values showing whether the values in columns (1) and (2) are statistically different. Data on urgent care centers come from the NUCR database, data on retail clinics come from Merchant Medicine, and data on county-level characteristics come from the 2016–2020 five-year ACS.

Table A3: Changes in insurance and retail clinic penetration: OLS by expansion status

	All counties		Medicaid non-expanders		Medicaid expanders	
	(1)	(2)	(3)	(4)	(5)	(6)
Dependent variable: retail clinics per 100,000	0.242 (0.412)		1.436** (0.557)		-1.525*** (0.406)	
Share insurance		0.979** (0.453)		1.736*** (0.583)	-1.370** (0.577)	
Share private			-1.279*** (0.448)	0.870 (0.785)	-1.465*** (0.467)	
Share Medicaid		X	X	X	X	X
County fixed effects	X	X	X	X	X	X
Year fixed effects	X	X	X	X	X	X
Demographic controls	X	X	X	X	X	X
Observations	3,869	3,869	2,038	2,038	1,831	1,831
R <sup>2</sup>	0.898	0.900	0.888	0.890	0.914	0.914
Mean dependent variable (2013)	0.598	0.598	0.691	0.691	0.528	0.528

Notes: The above table shows the association between retail clinics per 100,000 people and the share of the population with different types of health insurance from estimation of equations **(1)** and **(2)**. These regressions are estimated using all counties (columns (1)–(2)), counties in states that did not expand Medicaid under the ACA (columns (3)–(4)), and counties in states that did expand Medicaid under the ACA (columns (5)–(6)). The results in columns (1) and (2) were first reported in Table 2. Observations are at the county-year level from 2010 to 2016. All specifications include county fixed effects, year fixed effects, and all time-varying socio-demographic controls listed in Table 1. Standard errors are clustered by county.

Table A4: First-stage results from first-difference specification

Dependent variable:	Share insured (1)	Share private (2)	Share Medicaid (3)
$Post_t \times Employed_c^{2013}$	0.064*** (0.016)	0.073*** (0.017)	0.002 (0.012)
$Post_t \times [138 - 400\% FPL]_c^{2013}$	0.279*** (0.053)	0.225*** (0.052)	0.014 (0.037)
$Post_t \times Expansion_s \times [< 138\% FPL]_c^{2013}$	0.147** (0.065)	-0.023 (0.052)	0.172*** (0.036)
Year fixed effects	X	X	X
Demographic controls	X	X	X
Observations	3,314	3,314	3,314
$R^2$	0.515	0.334	0.285
Mean dependent variable	0.017	0.007	0.006

Notes: The above table shows first-stage estimates of the relationship between our instruments and the first difference of the share of the population with different types of health insurance from estimation of equation (A2). Table 4 provides corresponding two-stage least squares results showing the effects of the share of the population with different types of health insurance on the number of retail clinic entries and exits. Observations are at the county-year level from 2010 to 2016. All specifications include year fixed effects and the first difference of all time-varying socio-demographic controls listed in Table 1. These specifications further control for the number of retail clinics per 100,000 people in the previous period to account for the fact that openings (closings) are less (more) common in markets with many retail clinics. Standard errors are clustered by county.

Table A5: Effects of insurance on retail clinic penetration: ESI versus direct purchase

Dependent variable:	OLS		2SLS	
	Retail clinics per 100,000 (1)	Share ESI (2)	Share direct (3)	Share Medicaid (4)
<b>a. First stage</b>				
$Post_t \times Employed_c^{2013}$	0.173*** (0.019)		0.025** (0.011)	-0.056*** (0.017)
$Post_t \times [138 - 400\% FPL]_c^{2013}$	0.127** (0.059)		0.242*** (0.061)	0.089 (0.071)
$Post_t \times Expansion_s \times [< 138\% FPL]_c^{2013}$	0.106* (0.064)		-0.142*** (0.050)	0.321*** (0.050)
<b>b. OLS/ 2SLS</b>				
Share ESI	0.904* (0.486)			7.517*** (2.275)
Share direct purchase	1.235* (0.697)			-7.514* (4.501)
Share Medicaid	-1.267*** (0.444)			-7.045*** (2.479)
County fixed effects	X	X	X	X
Year fixed effects	X	X	X	X
Demographic controls	X	X	X	X
Observations	3,869	3,869	3,869	3,869
$R^2$	0.900	0.979	0.914	0.964
Mean dependent variable	0.6120	0.526	0.086	0.145
First stage F-stat				37.9

Notes: The above table shows the effects of the share of the population with different types of private health insurance coverage (employer-sponsored [ESI] or direct purchase) on the number of retail clinics per 100,000 people from estimation of analogues of equation (2) (column (1)) and equation (4) (column (5)). First-stage estimates showing the relationship between our instruments and the share of the population with different types of health insurance from estimation of analogues of equation (3) are provided in columns (2)–(4). Observations are at the county-year level from 2010 to 2016. All specifications include county fixed effects, year fixed effects, and all time-varying socio-demographic controls listed in Table 1. Standard errors are clustered by county. Cragg-Donald Wald F-statistics are reported.

Table A6: Effects of insurance on retail clinic penetration: instrument components

Dependent variable:	“Private” instrument 1			“Private” instrument 2		
	Share private (1)	Share Medicaid (2)	Retail clinics per 100,000 (3)	Share private (4)	Share Medicaid (5)	Retail clinics per 100,000 (6)
<b>a. First stage</b>						
$Post_t \times Employed_c^{2013}$				0.211*** (0.024)	-0.053** (0.017)	
$Post_t \times [138 - 400\% FPL]_c^{2013}$	0.431** (0.064)	0.072 (0.072)				
$Post_t \times Expansion_s \times [< 138\% FPL]_c^{2013}$	-0.123 (0.080)	0.346*** (0.052)		0.021 (0.085)	0.335*** (0.049)	
<b>b. Two-stage least squares</b>						
Share private				-1.726 (1.997)	6.086*** (1.791)	
Share Medicaid				-3.676 (2.241)	-2.215 (2.478)	
County fixed effects	X	X	X	X	X	X
Year fixed effects	X	X	X	X	X	X
Demographic controls	X	X	X	X	X	X
Observations	3,869	3,869	3,869	3,869	3,869	3,869
$R^2$	0.979	0.963		0.980	0.963	
Mean dependent variable	0.613	0.145	0.6120	0.613	0.145	0.6120
First stage F-stat		74.9				133.9

Notes: The above table shows the effects of the share of the population with different types of health insurance on the number of retail clinics per 100,000 people from estimation of equation (4) (panel (b)) using either only one of our two private insurance instruments. First-stage estimates showing the relationship between these subsamples of our instruments and the share of the population with different types of health insurance from estimation of equation (3) are provided in panel (a). Observations are at the county-year level from 2010 to 2016. All specifications include county fixed effects, year fixed effects, and all time-varying socio-demographic controls listed in Table 1. Standard errors are clustered by county. Cragg-Donald Wald F-statistics are reported.

## C Theoretical extensions

In this section, we consider an extension of the theoretical model in which we relax the assumption that firms can charge only a single price. In particular, we model firms as being able to charge different prices to Medicaid and non-Medicaid patients. That is, while the price paid by Medicaid patients is still fixed administratively at  $p_M$ , the firm can choose to charge a price  $p \neq p_M$  to non-Medicaid patients while simultaneously charging a price of  $p_M$  to those covered by Medicaid (and thereby serving the program's beneficiaries).

### C.1 Two-price model

Recall that in the one-price case outlined in Section III, the firm chooses the total quantity of patients served to maximize total profits. The optimal quantity of patients served is achieved by setting  $p^* = p(q^*) = D^{-1}(q^*)$ , where  $D$  denotes the total demand facing the clinic. The optimal price  $p^*$  in turn dictates whether the firm accepts Medicaid: if  $p^* > p_M$ , the clinic does not serve Medicaid patients, whereas the clinic accepts patients covered by Medicaid if  $p^* \leq p_M$ . Since the firm charges the same price to patients regardless of insurance type, the firm is indifferent between Medicaid and non-Medicaid patients conditional on accepting Medicaid.

In contrast, when the firm charges different prices to Medicaid and non-Medicaid patients, the firm cares both about the total number of patients served and the composition of patients by payer type. But since firms in the on-demand health care market see patients on a first come, first served basis, they can only indirectly influence the composition of patients they ultimately treat with the price they choose to charge non-Medicaid patients. As outlined below, this price is determined by the maximum number of non-Medicaid patients that the firm would want to treat, denoted by  $\tilde{q}$ . This inability of on-demand clinics to perfectly control patient composition makes it difficult to depict the firm's problem and our theoretical results graphically as in Figures 3 and 4 for the one-price case.<sup>40</sup> We therefore instead present

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<sup>40</sup>This is an important distinction in the case of on-demand health care clinics relative to traditional doctors' offices. If clinics are able to directly set the quantity of Medicaid (or non-Medicaid) patients that they see, the problem simplifies, and the firm sets marginal revenues equal for Medicaid and non-Medicaid patients if they choose to accept Medicaid. We then have a kinked, continuous marginal revenue function, and the intuition for how the impacts of health insurance expansions vary depending on whether the firm

the firm's problem and comparative statics in equations below.

In the two-price case, the firm chooses the total number of patients served,  $q$ , and the maximum number of non-Medicaid patients they would want to treat,  $\tilde{q} \leq q$ , to maximize profits. These choices in turn dictate whether the firm accepts Medicaid: if  $q^* = \tilde{q}^*$ , the clinic does not serve the Medicaid market, whereas the clinic accepts patients covered by Medicaid if  $q^* > \tilde{q}^*$ . Letting  $c$  denote the firm's cost function and  $N_M$  denote the total number of Medicaid patients in the market, the firm's maximization problem is given by:

$$\max_{q, \tilde{q}} \bar{p}(q, \tilde{q}) \cdot q - c(q)$$

where

$$\bar{p}(q, \tilde{q}) = \begin{cases} p(\tilde{q}) = p(q) & \text{if } \tilde{q} = q \\ p(\tilde{q}) \frac{\tilde{q}}{\tilde{q} + N_M} + p_M \frac{N_M}{\tilde{q} + N_M} & \text{if } \tilde{q} < q \end{cases}$$

Note that the average price received by the firm depends on whether the firm accepts Medicaid. If the firm does not accept Medicaid (i.e.,  $\tilde{q} = q$ ), then the average price is simply the price that the firm charges non-Medicaid patients ( $p(\tilde{q})$ ). If the firm serves the Medicaid market (i.e.,  $\tilde{q} < q$ ), then the average price per patient is a weighted average between the price charged to non-Medicaid patients ( $p(\tilde{q})$ ) and the administratively fixed Medicaid rate ( $p_M$ ), where the weights are the expected shares of patients that are non-Medicaid patients ( $\frac{\tilde{q}}{\tilde{q} + N_M}$ ) and Medicaid patients ( $\frac{N_M}{\tilde{q} + N_M}$ ), respectively.

## C.2 Insurance expansions

We begin by considering the effects of private insurance expansions. We model a private insurance expansion as an outward shift of the demand function of non-Medicaid patients (i.e.,  $p(\tilde{q} - \lambda)$  for some  $\lambda > 0$ ) and use the envelope theorem to quantify the local effects on firm profits.<sup>41</sup> Mirroring the predictions in Section III and the findings in Figure 7, we show

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accepted Medicaid patients at baseline is similar to the one-price case presented in the main text.

<sup>41</sup>Note that the necessary conditions for the envelope theorem require that the firm is not near the point of indifference between accepting and not accepting Medicaid. It must be the case that locally the first-order conditions are sufficient for firm optimality.

below that private insurance expansions will have larger positive effects on firm profits in places in which Medicaid prices are lower.

Suppose first that the firm does not accept Medicaid. Since  $\tilde{q} = q$  and  $\bar{p}(q, \tilde{q}) = p(q)$ , the firm's profits are given by:

$$\Pi(q; \lambda) = p(q - \lambda) \cdot q - c(q)$$

Let  $\Pi(\lambda) = \Pi(q^*(\lambda); \lambda)$  denote the maximum profit function. By the envelope theorem, we have that  $\Pi'(\lambda) = \Pi_\lambda(q^*(\lambda); \lambda)$ , and thus  $\Pi'(0) = -p'(q^*)q^*$ . Since  $p' < 0$  (i.e., demand slopes downwards), it follows that private insurance expansions lead firm profits to increase, which should in turn induce entry of additional clinics. Moreover, assuming that  $p'$  is locally constant, the positive impacts of private insurance expansions on firm profits are increasing in the number of (non-Medicaid) patients that the firm serves.

Now suppose that the firm also serves the Medicaid market. In this case, the firm's profits are given by:

$$\Pi(q, \tilde{q}; \lambda) = \left[ p(\tilde{q} - \lambda) \frac{\tilde{q}}{\tilde{q} + N_M} + p_M \frac{N_M}{\tilde{q} + N_M} \right] \cdot q - c(q)$$

Again let  $\Pi(\lambda) = \Pi(q^*(\lambda), \tilde{q}^*(\lambda); \lambda)$  denote the maximum profit function. By the envelope theorem, we have that  $\Pi'(\lambda) = \Pi_\lambda(q^*(\lambda), \tilde{q}^*(\lambda); \lambda)$ , and thus  $\Pi'(0) = -p'(\tilde{q}^*) \frac{\tilde{q}^*}{\tilde{q}^* + N_M} q^*$ .<sup>42</sup> As in the case in which firms do not accept Medicaid, private insurance expansions lead firm profits to increase (i.e.,  $\Pi'(0) > 0$ ). Moreover, since  $\frac{\tilde{q}^*}{\tilde{q}^* + N_M} \cdot q^*$  reflects the number of non-Medicaid patients that the firm treats, the positive impacts of private insurance expansions on firm profits are again increasing in the number of patients that the firm serves from the non-Medicaid market.

In the two-price case, private insurance expansions therefore lead firm profits to increase both when the firm does and does not accept Medicaid at baseline. This differs from the one-price case in that private insurance expansions are not predicted to affect the profitability of firms that serve the Medicaid market by setting the price they charge to both Medicaid and non-Medicaid patients at  $p_M$ . However, recall that firms are more likely to price at

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<sup>42</sup>If the firm serves the entire Medicaid market (i.e.,  $q^* = \tilde{q}^* + N_M$ ), then the expression for  $\Pi'(\lambda)$  reduces to the same expression as in the case in which the firm does not accept Medicaid.

$p = p_M$  (thus serving the Medicaid market) in the one-price case when  $p_M$  is higher, thereby leading to the prediction that the positive impacts of private insurance expansions on firm profitability should be larger when  $p_M$  is lower. As shown above, the positive impacts of private insurance expansions on firms profits in the two-price case are increasing in the number of patients that the firm serves from the non-Medicaid market. Since the number of non-Medicaid patients served is generally decreasing in the Medicaid price, the prediction that private insurance expansions should have the largest positive effects on firm profits (and in turn, firm entry) where Medicaid prices are low likewise holds in the two-price case.<sup>43</sup>

Now consider the effects of a Medicaid expansion. As shown in Figure 3(b), a Medicaid expansion causes both an inward shift of the demand function of non-Medicaid patients and an increase in the number of patients covered by Medicaid ( $N_M$ ). The decrease in non-Medicaid demand leads to the opposite predictions as those outlined in the case of private insurance expansions above. That is, the inward demand shift leads firm profits to decline, with the largest negative effects on firm profits occurring where Medicaid prices are low.

Moreover, the increase in  $N_M$  resulting from a Medicaid expansion will serve to either amplify or counterbalance these negative effects on firm profits depending on the relative Medicaid price. To see this, note that an increase in  $N_M$  leads to an increase in the expected share of patients served who are covered by Medicaid, thereby leading the average price per patient ( $\bar{p}$ ) to move toward the Medicaid price. If  $p(\tilde{q}^*) > p_M$ , which is more likely to be the case when  $p_M$  is low, then an increase in  $N_M$  leads  $\bar{p}$  to decline. This further reduces profits following a Medicaid expansion. On the other hand, if  $p(\tilde{q}^*) < p_M$ , then the increase in  $N_M$  leads  $\bar{p}$  to increase and offsets the reduction in profits stemming from the inward demand shift.<sup>44</sup> Mirroring the predictions from the one-price case, we therefore have that Medicaid

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<sup>43</sup>When the firm accepts Medicaid, this follows from the firm's first-order condition for  $\tilde{q}$  (given by  $p'(\tilde{q}^*)\tilde{q}^* + (p(\tilde{q}^*) - p_M)\frac{N_M}{\tilde{q}^* + N_M} = 0$ ). Differentiating with respect to  $p_M$  and limiting to cases in which  $p(\tilde{q}^*) > p_M$  yields

$$\left[ \underbrace{p''(\tilde{q}^*)\tilde{q}^*}_{=0} + \underbrace{p'(\tilde{q}^*) \left( 1 + \frac{N_M}{\tilde{q}^* + N_M} \right)}_{<0} + \underbrace{(p(\tilde{q}^*) - p_M) \frac{-N_M}{(\tilde{q}^* + N_M)^2}}_{<0} \right] \tilde{q}'(p_M) - \underbrace{\frac{N_M}{\tilde{q}^* + N_M}}_{>0} = 0$$

We therefore have that  $\tilde{q}'(p_M) < 0$ . Firms in places with higher Medicaid rates will therefore see relatively fewer non-Medicaid patients and thus will be less affected by private insurance expansions.

<sup>44</sup>An increase in profits stemming from  $N_M$ 's effect on  $\bar{p}$  could be large enough to fully outweigh the

expansions should have the largest negative effects on firm profits (and in turn, firm exit) where Medicaid prices are low.

By focusing on cases in which the envelope theorem applies (i.e., cases in which the first-order conditions are sufficient for optimality), we have limited ourselves to settings in which insurance expansions do not affect the firm's decision over whether to accept Medicaid. Of course, as in the one-price case, insurance expansions can lead firms who were serving both the Medicaid and non-Medicaid markets to only serve non-Medicaid patients and vice versa. Allowing insurance expansions to affect the firm's decision over whether to accept Medicaid does not affect the prediction that private insurance expansions should lead firm profits to increase: since private insurance expansions cause profits to weakly increase at the firm's baseline choice of  $\{q^*, \tilde{q}^*\}$  (by increasing demand among the non-Medicaid market), the firm will only adjust their decision of whether to accept Medicaid if doing so leads their profits to increase further. However, since Medicaid expansions increase the number of patients that firms can serve at the Medicaid price, some firms who did not serve the Medicaid market at baseline may experience an increase in profits following a Medicaid expansion if they switch to serving both non-Medicaid and Medicaid patients. While we therefore emphasize that Medicaid expansions need not always lead firm profits to decrease, the key theoretical result that Medicaid expansions can cause the supply-side to contract (in both the one- and two-price case) nevertheless holds.

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negative profit effects of the inward demand shift. This potential for Medicaid expansions to lead to net profit increases in the two-price case is analogous to the setting outlined in footnote 17 for the one-price case. However, as we typically do not see clinics pricing below  $p_M$ , this case is more of a theoretical possibility than an empirical reality.

## D Additional specifications

In Section V.D, we conduct two set of analyses to examine whether the effects of health insurance coverage on the concentration of retail clinics are driven by entries, exits, or both. This section provides estimating equations for these analyses.

First, we estimate analogues of the two-way fixed effects specifications introduced in Section IV.A using the number of clinic entries or exits at the county-year level as the dependent variable. As entries and exits are flow measures rather than stocks, we specify the right-hand side of each equation in first differences when considering these outcomes. We further control for the number of retail clinics per 100,000 people in the previous period to account for the fact that openings (closings) are less (more) common in markets with many retail clinics. Letting  $Entries_{ct}$  and  $Exits_{ct}$  denote the number of retail clinic entries or exits at the county-year level, respectively, we estimate the following specification:

$$\begin{aligned} \{Entries_{ct}, Exits_{ct}\} = & \{\beta \cdot \Delta Insured_{ct}, \beta_1 \cdot \Delta Private_{ct} + \beta_2 \cdot \Delta Medicaid_{ct}\} \\ & + \eta \cdot Clinics_{ct-1} + \delta \cdot \Delta X_{ct} + \gamma_t + \epsilon_{ct}, \end{aligned} \quad (A1)$$

where  $\Delta$  denotes the first-difference operator and all other variables are defined as in equations (1) and (2). Throughout this section, observations are weighted by county population in 2010, and standard errors are clustered by county.

Second, we estimate analogues of the instrumental variables specifications introduced in Section IV.B using the number of clinic entries or exits at the county-year level as the dependent variable. In these specifications, both the first- and second-stage regressions are in first differences. In particular, we estimate the following first-stage regressions to predict one-year changes in the share of the population with any insurance, private insurance, and

Medicaid coverage:

$$\begin{aligned}
\{\Delta Insured_{ct}, \Delta Private_{ct}, \Delta Medicaid_{ct}\} = \\
\alpha_1 \cdot \Delta \left\{ Post_t \cdot Employed_c^{2013} \right\} + \alpha_2 \cdot \Delta \left\{ Post_t \cdot [138 - 400\% FPL]_c^{2013} \right\} \\
+ \alpha_3 \cdot \Delta \left\{ Post_t \cdot [< 138\% FPL]_c^{2013} \cdot Expansion_s \right\} \\
+ \alpha_4 \cdot \Delta \left\{ Post_t \cdot [< 138\% FPL]_c^{2013} \right\} + \alpha_5 \cdot \Delta \left\{ Post_t \cdot Expansion_s \right\} \\
+ \delta \cdot \Delta X_{ct} + \eta \cdot Clinics_{ct-1} + \gamma_t + \epsilon_{ct}.
\end{aligned} \tag{A2}$$

We then estimate the following second-stage regressions using the predicted insurance changes from equation (A2):

$$\begin{aligned}
\{Entries_{ct}, Exits_{ct}\} = & \left\{ \beta \cdot \widehat{\Delta Insured}_{ct}, \beta_1 \cdot \widehat{\Delta Private}_{ct} + \beta_2 \cdot \widehat{\Delta Medicaid}_{ct} \right\} \\
& + \alpha'_4 \cdot \Delta \left\{ Post_t \cdot [< 138\% FPL]_c^{2013} \right\} + \alpha'_5 \cdot \Delta \left\{ Post_t \cdot Expansion_s \right\} \\
& + \delta' \cdot \Delta X_{ct} + \eta' \cdot Clinics_{ct-1} + \gamma'_t + \epsilon'_{ct},
\end{aligned} \tag{A3}$$

where  $\Delta$  again denotes the first-difference operator, and all other variables in equations (A2) and (A3) are defined as in equations (3) and (4).