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# THE EXPANSIONARY AND CONTRACTIONARY SUPPLY-SIDE EFFECTS OF HEALTH INSURANCE

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**ABSTRACT**

We examine how health insurance expansions influence the entry and location decisions of health care providers. While it is generally expected that reductions in consumer prices following insurance expansions will prompt supply-side growth to meet increased demand (Arrow, 1963), we demonstrate both theoretically and empirically that expansions of insurance with low, regulated provider reimbursements (e.g., Medicaid) can instead lead the supply side to contract. Additionally, we show that expansions of insurance with high, market-based prices (e.g., private coverage) can reduce acceptance of public coverage even as they drive growth in the overall concentration of providers.

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# I Introduction

Insurance expansions increase demand for health care by reducing 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 increase in response to sizable insurance expansions (Finkelstein, 2007; Kondo and Shigeoka, 2013; Hackmann et al., 2025). However, not all health insurance is created equal, and expansions of insurance with generous patient cost-sharing but low reimbursement rates for providers may 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 (Baicker et al., 2023; Einav and Finkelstein, 2023), an understanding of how relative insurance generosity affects supply-side responses to health insurance provision is key.[1]

In this paper, we examine how the supply side responds to health insurance expansions in two-tiered systems. We introduce a theoretical framework demonstrating that expansions of insurance with low, regulated prices (e.g., Medicaid) can reduce the number of providers available to all patients—even those with private coverage—through market-level contractions. Moreover, while expansions of insurance with relatively high, market-based prices (e.g., private insurance) will often lead to supply-side growth, such expansions can reduce the number of providers available to publicly insured patients by leading providers to stop serving the program's beneficiaries. These supply-side responses are closely linked to relative insurance generosity for providers, with the opposing effects of private and public insurance expansions concentrated in markets where regulated payments in the public system are low.

We test these predictions using rich data on provider locations and county-level changes in insurance coverage following the Affordable Care Act, the largest health insurance expansion in the United States in decades. Our primary analysis focuses on retail clinics, which, together with urgent care centers, form the on-demand care market. The on-demand sector

[1] In addition to the provision of health insurance through both public and private systems in the United States, many countries with extensive public health insurance systems—including Australia, Canada, Germany, Israel, Singapore, and the United Kingdom—offer supplemental private health insurance programs.

has driven much of the growth in health care markets in recent decades and now serves almost one-third of the adult population (NCHS, 2022).[2] Using two-way fixed effects and instrumental variable designs, we find that growth in private insurance coverage substantially increases clinic entry, whereas clinic penetration is dampened by increases in Medicaid coverage. These contrasting supply-side effects of private insurance and Medicaid expansions are concentrated in states with low Medicaid reimbursement rates and are also observed for other providers in the outpatient care market, including urgent care centers and physicians in general practice (GPs). Examining spillovers across insurance types, we further find that increases in private insurance coverage lead to reductions in the number of firms accepting Medicaid. These 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.

Our theoretical 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). We consider firms that can opt out of serving Medicaid patients, which applies to outpatient settings but not to hospitals with a legal obligation to provide uncompensated care. As in Garthwaite (2012) and Dillender (2022), 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

[2]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 (NPs) or physician assistants (PAs), 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 NPs and PAs, and often have imaging equipment. Nearly 30 percent of adults reported going to a retail clinic or an urgent care center in 2019, which is significantly higher than the 22 percent of adults that reported going to a hospital emergency room (NCHS, 2022).

reduces profits and induces exit, again among firms that were not accepting Medicaid 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, insurance expansions generally have no effects on the profits of clinics that were serving both markets at baseline. However, as discussed below, they can change the decision to accept Medicaid among firms near the point of indifference between serving both markets versus only the private market. 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 regulated payment under Medicaid is higher. The opposing supply-side responses to growth in private insurance and Medicaid coverage on firm entry patterns 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 assess how retail clinic concentration covaries with county-year changes in health insurance coverage. We combine data from 2010 to 2016 on insurance coverage by type from the one-year American Community Survey (ACS) with information on the entry and exit of the universe of retail clinics from Merchant Medicine. This analysis reveals a surprising result demonstrating that positive supply-side responses need not accompany sizable insurance expansions: despite significant variation in insurance coverage and clinic penetration over our sample period, there was no association between within-county changes in the share of the population with health insurance coverage of any type and clinic growth. In line with the theory, however, this null effect masks pronounced heterogeneity by insurance type. Estimating 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, we find that growth in private insurance coverage is associated with large increases in clinic concentration, whereas growth in Medicaid coverage is 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 greater in counties located in states that expanded Medicaid and had a larger share of residents below 138 percent of the federal poverty level (FPL; i.e., the population eligible for coverage under a Medicaid expansion). Moreover, we introduce two novel instruments that affect private insurance coverage.[3] The first exploits baseline variation in the share of the population between 138 and 400 percent of the FPL (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.[4] 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 sample counties over the sample 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 effects are driven by impacts on both entries and exits, with private insurance growth leading

[3]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.

[4]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.

to increased clinic entry and Medicaid growth leading to increased clinic exit.

These opposing supply-side responses to insurance expansions are likely driven by low reimbursement rates under Medicaid.[5] As outlined above, theory predicts that the supply-side responses to 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 (2024), we show that the positive effects of growth in private insurance coverage on retail clinic penetration are most 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.

In addition to affecting firm concentration, health insurance expansions should also influence the types of coverage that providers accept. The theoretical model predicts that while expansions of private insurance generally encourage supply-side growth, they may lead firms that were on the margin of not accepting Medicaid at baseline to opt out of serving the program's beneficiaries. This occurs because these firms can now operate at full capacity at a higher price in the non-Medicaid market. Conversely, although Medicaid expansions may encourage some firms that did not initially serve the Medicaid market to begin accepting Medicaid coverage, these effects are tempered by reduced profits and the subsequent exit of firms that were only serving the non-Medicaid market at baseline.

We test these predictions regarding the acceptance of public insurance using hand-collected data on Medicaid acceptance by retail clinics across eight states from 2010 to 2016. The results confirm that both the share of retail clinics accepting Medicaid and the

[5] Alternatively, gaining Medicaid might lead patients to shift toward more traditional sources of care, reducing demand for retail clinics and promoting exit. We find little support for this mechanism: the negative supply-side effects of Medicaid are largest in areas with the least health care resources at baseline (i.e., areas with the least scope for substitution), and the results are unaffected when controlling for growth in federally funded clinics serving underserved populations. Moreover, similar location patterns are observed among GPs, a provider type for which expanding insurance coverage has been shown to increase demand.

concentration of Medicaid-accepting clinics decline in response to private insurance growth. Thus, while private insurance expansions often lead to an increase in the total number of firms, they can also reduce access to care for Medicaid patients. Further consistent with the theory, we find limited impacts of growth in Medicaid coverage on the concentration of firms accepting Medicaid. Since Medicaid expansions reduce the overall concentration of firms, these patterns suggest that growth in Medicaid coverage may limit access for privately insured patients while failing to generate sufficient increases in the concentration of providers accepting Medicaid to meet rising demand in the Medicaid market.

Although our primary analysis focuses on retail clinics, we show that our findings likely extend to other providers in the outpatient care market. Notably, the underlying theory applies to any providers that screen patients for insurance coverage and type before providing care—such as urgent care centers and many GPs—and, as a result, do not provide uncompensated care.[6] Using data on the locations of all urgent care centers in 2021, we demonstrate that urgent care centers and retail clinics exhibit similar location patterns, with both types of firms concentrating in areas with high rates of private insurance coverage and low shares of Medicaid beneficiaries.[7] Similar patterns are also observed among GPs, with data from the Centers for Medicare & Medicaid Services (CMS's) Doctors and Clinicians National Downloadable File showing that GP concentration increases with a county's share of privately insured residents and decreases with the share covered by Medicaid. These patterns underscore how providers' location choices vary with insurance composition in ways that may have widespread implications for patient access in two-tiered health insurance systems.

Our work contributes to three literatures. First, we build on the literature examining the supply-side effects of health insurance. Seminal work by Finkelstein (2007) found that the introduction of Medicare in 1965 led to hospital entry and increased adoption of new

[6]Since emergency rooms are legally required to stabilize patients regardless of insurance coverage or type, they will benefit from growth in Medicaid with low, regulated payments if the probability of non-payment is otherwise high (see Section II.B).

[7]These findings relate to work by Magnolfi et al. (2024), who estimate an equilibrium model of market structure and find that hospitals deter 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 such clinics—by avoiding areas with growth in Medicaid and concentrating in areas with high rates of private insurance coverage and existing health care resources—are unlikely to meaningfully address access barriers as many health policy experts have hoped (see, for example, Bechrach and Frohlich, 2016).

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, Hackmann et al. (2025) document 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, showing that firm-level entry responses—anticipated mechanisms through which supply keeps pace with growing demand—depend on the type of coverage being expanded.

Our finding that the type of coverage being expanded is important for shaping supply-side responses relates to a broader literature on how insurance generosity influences provider behavior. Recent work shows that providers' reluctance to accept Medicaid stems from the program's low reimbursement rates for providers relative to other payers (Alexander and Schnell, 2024) and billing hassles that plague the Medicaid system (Dunn et al., 2024). Focusing on technology adoption, Freedman et al. (2015) find that expansions of Medicaid eligibility for pregnant women in the 1980s and 1990s did not lead hospitals to adopt neonatal intensive care units. In the pharmaceutical sector, Garthwaite et al. (2021) document that research and development—which are typically responsive to changes in market size—did not increase in response to recent Medicaid expansions.[8] Similarly, Dillender (2022) finds that job vacancies for health care workers rose after the ACA's Medicaid expansions, but only among firms hiring lower-skilled workers.[9] Each of these studies attributes muted supply-side responses to Medicaid's low reimbursement rates. We contribute to this literature by showing that provider location decisions and firm-level entry responses are also 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, the contractionary effects of Medicaid expansions that we uncover broaden our

[8]Focusing on 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 ineligible for traditional Medicaid.

[9]As outlined in Section II.B, our results should not extend to providers that offer uncompensated care. It is therefore possible that such providers expanded their capacity even as other parts of the market contracted.

understanding of how demand-side subsidies affect the private provision of social services. To maintain choice and promote competition, policymakers often offer consumers subsidies to purchase services such as health care, child care, and education from private providers. Recent work shows that programs allowing consumers to "top up" subsidy payments, as is common in child care markets, can raise prices and harm unsubsidized consumers (Bodéré, 2023). In addition, quantity-based subsidies, such as those offered in some supplemental nutrition programs, can lead to higher prices and reduce program participation (Meckel, 2020; Abito et al., 2022; An et al., 2023; Wang, 2024). We add to this work by showing that expansions of subsidies with prohibitions on "balance billing"—requiring participating providers to offer services at regulated prices—can reduce the number of providers serving the entire market, thereby affecting access for both subsidized and unsubsidized consumers.

This paper proceeds as follows. Section II presents a theoretical framework that delivers predictions about the impacts of changes in insurance provision on clinic entry, exit, and the acceptance of insurance with regulated prices. Section III introduces our primary data sources and documents associations between insurance provision and retail clinic penetration, while Section IV isolates variation in insurance provision driven by policy changes using an instrumental variables design. Extensions, including an empirical examination of the role of reimbursement rates and an analysis of other health care providers, are provided in Section V. Section VI provides a discussion and concludes.

# II Theoretical framework

In this section, we introduce a theoretical model of entry and exit for health care clinics. 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. In contrast to this prior work, which focused on how reimbursement rates under Medicaid and private insurance affect the decision of whether to participate in the Medicaid program conditional on entry, our primary goal is to instead examine how changes in the share of the population covered by different types of health insurance affect firm profitability and equilibrium market structure. As outlined below, the model delivers a number of theoretical insights that both rationalize our main findings and motivate additional empirical exercises.

In our setting, clinics decide whether to enter the market and, if so, which price to charge. We assume 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. Because our primary empirical exercises focus on on-demand clinics, which see patients on a first-come, first-served basis, we assume that clinics cannot choose the share of Medicaid-covered patients they treat, conditional on accepting Medicaid. Firms 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.

## II.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 1(a). Because Medicaid patients typically 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 $s_M$.[10] At all other prices, the firm faces downward-sloping demand from the non-Medicaid population. As shown in Figure 1(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.[11] This reflects the fact that clinics do not need

[10] 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). Moreover, federal law prohibits providers from balance billing Medicaid patients, requiring them to accept Medicaid's payment as payment in full.

[11] In Figure 1(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 Medicaid price is relatively high (i.e., few non-Medicaid patients have willingness to pay greater than $p_M$).

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

Figure 1(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 firms 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 1(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 Medicaid patients 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 A1 shows how prices and quantities are determined in each of these cases. When there are two intersections between marginal revenue and marginal costs (as in Figure A1(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 (Figure A1(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).

## II.B Insurance expansions and firm behavior

We aim to examine how health insurance expansions affect firm profits and in turn, firm entry and exit decisions. In doing so, we further outline how health insurance expansions affect a firm's decision of whether to accept Medicaid. We begin by demonstrating how expansions of Medicaid coverage and private insurance affect the demand facing clinics. As shown in Figure 1(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 1(b), this rotates demand among the non-Medicaid population upward.

**Entry and exit** How do expansions of private insurance and Medicaid coverage 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. We consider two periods, where subscript "1" denotes the period before the expansion and subscript "2" denotes the period after. As shown in Figure 2(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.[12] As shown in Figure 2(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 2(c), an inward shift in the demand curve among non-Medicaid patients leads clinics that did not accept

[12] 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 A2(a), while optimal pricing is unaffected (i.e., $p_2^* = p_1^* = p_M$), optimal quantity—and firm profits—will increase.

Medicaid at baseline to decrease quantity and prices (i.e., $q_2^* < q_1^*$ and $p_2^* < p_1^*$). Profits 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.[13] As shown in Figure 2(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, summarized in the following results:

**Result 1:** *Expansions of private insurance weakly increase firm profitability, thereby leading to additional entry. In contrast, expansions of Medicaid can reduce firm profits and lead to clinic exit.*

**Result 2:** *The supply-side effects of insurance expansions vary depending on whether clinics were accepting Medicaid at baseline (or would have accepted Medicaid had they entered), with the positive (negative) supply-side effects of private insurance (Medicaid) coverage concentrated among firms for which it is not profitable to serve the program's beneficiaries.*

The opposing supply-side effects of private insurance and Medicaid expansions outlined in Result 1 are the focus of Sections III and IV. Moreover, Result 2 motivates our analysis of the heterogeneous supply-side effects of insurance expansions by baseline Medicaid reimbursement rates in Section V.A. 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 from Result 2 that supply-side responses to growth in private insurance and Medicaid coverage should be more pronounced in areas with low Medicaid reimbursement rates.

**Medicaid acceptance** How do health insurance expansions affect a firm's decision of whether to accept Medicaid? Recall that in the one-price case, whether the firm accepts

[13]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 A2(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.

Medicaid depends on how $p^*$ compares to $p_M$, with firms pricing at or below the Medicaid price (i.e., $p^* \leq p_M$) serving the Medicaid market and firms pricing above the Medicaid price (i.e, $p^* > p_M$) not participating in the program.

We again begin by considering the effects of a private insurance expansion. Consider first firms that did not accept Medicaid at baseline. Since the upward rotation in demand leads such firms to increase their price (i.e., $p_2^* > p_1^* > p_M$), firms that did not accept Medicaid at baseline will continue to only serve the non-Medicaid market following a private insurance expansion. However, some firms that did accept Medicaid at baseline will choose to opt out of the program when more patients have private insurance coverage. As shown in Figure A3(a), firms that accepted Medicaid at baseline and experienced an outward shift in the relevant portion of their demand curve following a private insurance expansion will subsequently switch to only serving the non-Medicaid market (i.e., $p_2^* > p_1^* = p_M$). These firms were on the margin of accepting Medicaid at baseline, and thus their decision of whether to serve the program's beneficiaries was sensitive to slight changes in demand. In contrast, firms that accepted Medicaid at baseline and were not on the margin of program participation due to relatively low marginal costs will continue to serve the program's beneficiaries following a private insurance expansion (i.e., $p_1^* \leq p_2^* \leq p_M$).

Now consider the effects of a Medicaid expansion. Since firms that accepted Medicaid at baseline find it optimal to charge the same price both before and after a Medicaid expansion (i.e., $p_2^* = p_1^* = p_M$), their participation in the Medicaid program is likewise unaffected. However, as shown in Figure A3(b), some firms that did not serve the Medicaid market at baseline will find it optimal to lower their price to such an extent that they will begin serving the Medicaid market (i.e., $p_M = p_2^* < p_1^*$). As was the case with private insurance expansions, these firms were on the margin of accepting Medicaid at baseline. Firms with very high marginal costs, for whom the intersection between marginal costs and marginal revenue is far from the perfectly elastic portion of the demand curve, lower their price in response to Medicaid expansions but continue to price above the Medicaid rate, thereby continuing to only serve the non-Medicaid market (i.e., $p_M < p_2^* < p_1^*$).

Our baseline model therefore also delivers predictions concerning the impacts of insurance expansions on Medicaid acceptance, summarized in the following result:

**Result 3:** *Expansions of private insurance weakly reduce Medicaid acceptance, whereas Medicaid expansions weakly increase Medicaid acceptance. These changes in Medicaid acceptance following expansions of private insurance (Medicaid coverage) occur among firms that were on the margin of not accepting (accepting) Medicaid at baseline.*

These opposing impacts of insurance expansions on Medicaid acceptance are the focus of Section V.B, in which we consider how private insurance and Medicaid expansions impact both the concentration and share of open retail clinics accepting Medicaid. When considering these outcomes, Result 3 interacts with Results 1 and 2. Following a private insurance expansion, areas in which only some firms accepted Medicaid at baseline might see the number of firms accepting Medicaid fall (Result 3) even as the total number of firms rises (Results 1 and 2). Moreover, following a Medicaid expansion, the number of firms accepting Medicaid in these areas could fall due to a contraction in the overall number of firms (Results 1 and 2) even as the share of firms serving the program's beneficiaries rises (Result 3).

**Extensions** 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 supply-side effects of insurance expansions that closely mirror those derived above. In particular, we show in Appendix C.2 that expansions of private insurance lead firm profits to increase whereas expansions of Medicaid generally lead firm profits to decline. Moreover, as in the one-price case, we show that areas with low Medicaid reimbursement rates experience the most pronounced supply-side responses to expansions of both private insurance and Medicaid coverage.[14]

Although our primary empirical exercises focus on on-demand clinics, the model can be used to develop insights into the supply-side effects of health insurance in other care settings. Notably, the predictions extend to any care setting in which providers can screen

[14] 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.

patients based on insurance coverage and type before providing care (i.e., providers that do not provide uncompensated care). Moreover, although on-demand clinics treat patients on a first-come, first-served basis, and thus they cannot control payer mix conditional on accepting multiple insurance types, all of the theoretical results hold when providers can "pick" their patients. Thus, we anticipate that providers such as traditional doctors' offices—which 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—to respond similarly to insurance expansions as on-demand clinics.[15] However, since emergency departments are required to stabilize patients regardless of their ability to pay, 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.

# III Insurance provision and clinic concentration

This section provides an initial look at supply-side responses to insurance expansions. After introducing our primary data sources in Section III.A, we outline the relationship between clinic concentration and insurance provision in the raw data and using two-way fixed effects specifications in Sections III.B and III.C, respectively. To ensure that these relationships are not driven by changes in local socio-demographics, we leverage variation in insurance provision driven by policy changes in Section IV. We focus on the retail clinic market in Sections III and IV and then turn to other providers in the outpatient care market whose entry patterns should theoretically exhibit similar relationships with local insurance provision, including urgent care centers and GPs, in Section V.C.

[15]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 ever more pronounced in these other settings. Moreover, 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.

## III.A Retail clinic and insurance data

Information on retail clinics comes from Merchant Medicine, a management consulting firm serving the on-demand health care market.[16] These data are comprehensive and contain the geo-coded locations and operating dates for 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 3(a), the number of retail clinics increased nationally by 66 percent between 2010 and 2016, rising from 1,224 at the start of 2010 to 2,036 by the end of 2016. The market also experienced considerable churn, with an average of more than 50 entries and nearly 22 exits per quarter over this period (see Figure A4).

These clinic data are combined with county-level characteristics from the ACS. The main independent variables of interest are the shares of the population with health insurance coverage overall and by type. 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.[17] 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 3(b)).

As shown in Figure 4(a), the share of the U.S. population with health insurance increased from below 85 percent in 2010 to nearly 92 percent in 2016. This increase was driven primarily by the implementation of the ACA in 2014. Over the same period, the number of open retail clinics (light, dashed line) rose steadily, suggesting a potential link between insurance expansion and clinic growth. Moreover, as shown in Figure 4(b), the increase in health insurance coverage nationally was driven by sizable increases in both private insurance

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

[17] 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 rather than changes in population. All analyses include a time-varying control for population to account for population growth.

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 4(c)).

We also use information on a range of other local socio-demographics from the ACS. As shown in Table 1, retail clinics in 2016 were disproportionately located in dense, urban hubs. This can further be seen in Figure 3(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.[18] This is true both when considering only counties in the one-year ACS (Table 1) and all U.S. counties (Table A1).

## III.B Patterns in the raw data

Although both insurance coverage and retail clinic penetration increased over the period, a causal relationship requires that clinic growth systematically tracked local changes in coverage. To assess this, we examine whether counties with greater gains in insurance coverage from 2013 to 2015 also experienced greater increases in retail clinics. Figure 5(a) groups counties into deciles based on changes in the share of residents with any type of health insurance over this period and plots the corresponding average change in retail clinics per 100,000 people. Strikingly, there is no clear relationship: while counties in the top decile of insurance growth saw average coverage increases of nearly 12 percentage points compared to only 2 percentage points among counties in the bottom decile, counties in both groups experienced increases of approximately 0.15 retail clinics per 100,000 people.

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. To assess 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

[18] Using data from the Area Health Resource files, Barnes et al. (2023) likewise document that counties with any retail clinics in 2016 were more affluent on average than those without.

and Medicaid coverage are somewhat negatively correlated over this time period (see Figure 4(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 notable 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 other types 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 A6. The 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, 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.[19] After the ACA's implementation, however,

[19]This is further shown in Figure A7(a), which plots county-level changes in retail clinics per 100,000 people from 2011 to 2013 against county-level changes in the share of the population with private insurance and Medicaid coverage from 2013 to 2015. While changes in the concentration of retail clinics from 2013 to 2015 are strongly correlated with contemporaneous changes in insurance provision (see Figure 5 and Figure A7(b)), counties with differing changes in insurance coverage under the ACA experienced very similar growth

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.

## III.C Two-way fixed effects specifications

To examine how health insurance coverage covaries with the concentration of retail clinics while controlling for general time trends and differences across locations, we estimate two-way fixed effects specifications. Letting $Clinics_{ct}$ denote the number of open retail clinics per 100,000 people in county $c$ in year $t$, we estimate the following specification:

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

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.

Results from estimation of equation (1) are provided in column (1) of Table 2. As suggested by the raw data, there is no statistically significant relationship between county-level growth in the share of the population with any type of health insurance coverage and 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 likely masks important heterogeneity by insurance type. To examine whether retail clinics respond differently to changes in the share of the population covered by different types of health insurance, we

in retail clinics in the years immediately prior.

estimate the following specification:

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

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).[20] Observations are again weighted by county population in 2010, and standard errors are clustered by county.

Results from estimation of equation (2) are shown in column (2) Table 2. Growth in private insurance coverage is associated with significant increases in clinic growth, whereas growth in Medicaid coverage is associated with significant reductions. The estimates indicate that growth in private insurance coverage of 5 percentage points—the average increase experienced by sample counties over the sample period—is associated with a relative increase of 0.049 retail clinics per 100,000 people, or 8.0 percent relative to the mean. Conversely, growth in Medicaid coverage of 4 percentage points—the average increase experienced by sample counties over the sample period—is associated with a relative reduction of 0.051 retail clinics per 100,000 people, or 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 are only associated with relative growth in the concentration of retail clinics in areas that experienced

[20] When considering effects by insurance type, $X_{ct}$ includes controls for the share of the population with Medicare and the share 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 than private or Medicaid were minimal over the period, results are nearly identical if we exclude these additional controls.

below-median increases in Medicaid coverage. Moreover, increases in Medicaid coverage are only associated with 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 potentially 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.[21]

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) include time-varying controls for socio-demographic characteristics—including median income, the poverty rate, and employment—and the timing of effects shown in Figures A6 and A7 suggests that the impacts are driven by the ACA rather than gradual socio-demographic shifts, it is nevertheless possible that any observed preference for the privately insured (and aversion to Medicaid patients) reflects a broader preference for wealthier populations rather than heterogeneity in supply-side responses to insurance coverage by type. To address this concern, we turn to an instrumental variables design that isolates changes in insurance coverage driven by policy variation.

[21] 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 A2, 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 the 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).

# IV Policy-induced changes in insurance provision

In this section, we isolate changes in health insurance coverage driven by policy changes using an instrumental variables design to measure 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. We introduce our empirical strategy and present impacts on the concentration of clinics in Section IV.A and IV.B, respectively. Section IV.C then considers impacts on entries and exits separately, and Section IV.D discusses identification and robustness.

## IV.A Instrumenting for insurance coverage

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 A8). As shown in Figure A9(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.

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 A9(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 share 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 A9(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.[22]

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 A9(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-stage regressions of the form:

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

where $Post_t$ is an indicator denoting years 2014 and onward; $Expansion_s$ is an indicator

[22] In Section IV.D, we use data from the Quarterly Census of Employment and Wages (QCEW) and show that our results are robust to using proxies for the baseline share of the population that was employed by firms that were most affected by the employer mandate.

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 shares of the population that were 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 ($Post_t \cdot [< 138\%\ FPL]_c^{2013} \cdot Expansion_s$) shifts individuals into Medicaid. In practice, however, our instruments will shift multiple types of insurance simultaneously due to crowd-out.[23] We therefore include all three instruments when considering each insurance type.

First-stage results from estimation of equation (3) 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.

[23]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.

## IV.B Effects on clinic concentration

To examine how policy-induced changes in insurance provision affect the concentration of clinics, we then estimate the following second-stage regressions:

$$\begin{aligned} Clinics_{ct} &= \left\{\beta \cdot \widehat{Insured}_{ct},\ \beta_1 \cdot \widehat{Private}_{ct} + \beta_2 \cdot \widehat{Medicaid}_{ct}\right\} \\ &\quad + \alpha_4' \cdot Post_t \cdot [< 138\%\ FPL]_c^{2013} + \alpha_5' \cdot Post_t \cdot Expansion_s \\ &\quad + \delta' \cdot X_{ct} + \gamma_c' + \gamma_t' + \epsilon_{ct}', \end{aligned} \tag{4}$$

where $Clinics_{ct}$ denotes the number of open retail clinics per 100,000 people in county $c$ in year $t$; $\widehat{Insured}_{ct}$, $\widehat{Private}_{ct}$, 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 III.C, 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.

The two-stage least squares estimates from estimation of equation (4) 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 sample counties over the sample 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 the sample 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.

One concern with this 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 before the onset of the ACA. We examine this possibility in Figure A10 by replicating the time-series path of retail clinic growth shown for the whole United States in Figure 3(a) separately across counties that are isolated by the different instrument components. Reassuringly, the figure shows that such counties were on similar trends in terms of retail clinic growth before 2014. However, there is a clear divergence in retail clinic penetration after the ACA's onset, 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.[24]

## IV.C Openings versus closings

Variation in the number of retail clinics over the sample period is driven both by entries and exits. As noted in Section III.A, there were over 50 entries and nearly 22 exits on average in each quarter from 2010 to 2016 (see Figure A4). To examine whether the effects of health insurance coverage on the concentration of retail clinics are driven by entries, exits, or both, we estimate analogs 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

[24] 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 A10(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.

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.[25]

Table 4 presents results from these analyses.[26] 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, 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. 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 column (4) indicates that growth in private insurance coverage of 0.86 percentage points—the average annual change among sample counties over the 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 the sample period—leads to an

[25] The estimating equations for these analyses are presented in Appendix D.

[26] First-stage results for the instrumental variables designs are provided in Table A3. The results confirm that annual changes in our instruments are strong predictors of annual changes in insurance growth.

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.

## IV.D 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.[27] 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.[28] 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-

[27] 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.

[28] 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.

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.[29] Moreover, the magnitudes of the associations in these balancing regressions are very small compared to our main results, 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 Sections III.C and IV.A, we include time-varying, county-level controls in our baseline specifications and weight county-level observations by county population in 2010. Figure A12 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

[29]Figure A11 displays results from one-step balancing regressions examining how insurance provision at the county-year level correlates with select socio-demographics across 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, 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.

states. As shown in the final row of each subfigure in Figure A12, 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 many large firms already offered 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 A13, 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, as small- to mid-size firms were the least likely to offer coverage before the ACA (KFF, 2013), 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.

# V Extensions

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

Below, 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** As outlined in Section II, 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.[30] 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 (2024). As shown in Figure A14(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.[31]

Splitting the sample by terciles of these Medicaid reimbursement rates, Figure 7 replicates the analysis in Figure 5 separately for states in the bottom and top terciles of baseline Medicaid reimbursements. 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 II, 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 clinics being the most pronounced in locations with low Medicaid reimbursement rates at baseline.

[30] More precisely, the theory predicts that clinics should be more likely to accept Medicaid when reimbursement rates under Medicaid are higher *relative* to other insurers. Combining information on Medicaid fee-for-service (FFS) to private insurance reimbursement ratios for office visits in 32 states in 2010 from the Government Accountability Office (GAO, 2014) with data on Medicaid rates in the same year from Alexander and Schnell (2024), we find that relative Medicaid payments are highly correlated with Medicaid FFS levels ($\rho = 0.73$). In what follows, we use Medicaid payment levels since we have this information for all states, although the results are very similar if we instead use payment ratios for a subset of states.

[31] We verify that retail clinics are more likely to accept Medicaid in states with higher Medicaid reimbursement rates. In particular, Figure A14(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.[32] 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 Health Resources and Services Administration (HRSA) on the number of primary care physicians and NPs ("primary care providers") per capita, we split the sample by terciles of primary care providers per capita in 2010 and replicate Figure 5 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 A15 and A12, respectively. Looking first to Figure

[32]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 NPs. 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.

A15, we see that the negative supply-side effects of Medicaid coverage are largest in areas with the least primary care providers per capita.[33] 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 A12(b), we see that our results are essentially unchanged when we control for the number of FQHCs per capita at the county-year level.[34] These analyses suggest that limited demand-side responses are unlikely to explain the negative supply-side effects of Medicaid coverage that we observe.

## V.B How do insurance expansions affect Medicaid acceptance?

Recall that the theoretical model outlined in Section II delivers predictions both about the impacts of health insurance expansions on the entry and exit of clinics as well as their acceptance of Medicaid. Despite generating increases in the overall number of clinics, private insurance expansions can lead firms to stop accepting Medicaid. Conversely, despite leading to reductions in the overall number of firms, Medicaid expansions can lead firms to start participating in the Medicaid program. These opposing effects on Medicaid acceptance should be concentrated in markets in which firms were on the margin of accepting Medicaid at baseline, with the effects of private expansions being observed in markets in which firms were only marginally accepting Medicaid and the effects of Medicaid expansions being observed in markets in which firms were only marginally not accepting Medicaid.

To examine these predictions empirically, we submitted Freedom of Information Act requests to relevant state agencies in all fifty states and the District of Columbia, requesting information on licensed providers registered as Medicaid providers in each year. These

[33] 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 A15(b) shows that effect heterogeneity by primary care providers per capita at baseline is very similar when controlling for baseline Medicaid reimbursement rates.

[34] Figures A15 and A12 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.

requests yielded annual data covering the period 2010 to 2016 in 13 states: Florida, Georgia, Iowa, Kentucky, Minnesota, Missouri, Montana, Nevada, North Dakota, Texas, Vermont, West Virginia, and Wisconsin. Unfortunately, Georgia does not update its Medicaid provider list when providers stop participating in the program, rendering the data unsuitable for our analysis. Moreover, as we use information on the share of clinics accepting Medicaid at baseline to isolate areas in which firms were near the point of indifference in our analysis below, we further exclude the four states in which there were no retail clinics open before the onset of the ACA (Montana, North Dakota, Vermont, and West Virginia).[35]

To assess how insurance expansions influence Medicaid acceptance, Figure 8 replicates the analysis in Figure 5 using two measures of Medicaid program participation as the outcome: the share of retail clinics accepting Medicaid coverage (left subplots) and the number of retail clinics accepting Medicaid per 100,000 people (right subplots).[36] As the effects of private insurance expansions on Medicaid acceptance should be concentrated in states where firms were marginally accepting Medicaid at baseline, we limit the analysis in Figure 8(a) to the six states in our sample in which more than 25 percent of retail clinics at baseline were accepting Medicaid coverage. Analogously, since the effects of Medicaid expansions on Medicaid acceptance should be concentrated in states where some firms were marginally not accepting Medicaid at baseline, we limit the analysis in Figure 8(b) to the six states in our sample in which fewer than 75 percent of retail clinics were accepting Medicaid coverage at baseline.

Strikingly, Figure 8(a) shows that both the share of retail clinics accepting Medicaid and the concentration of Medicaid-accepting clinics decline with growth in private insurance coverage. This means that even though private insurance leads to overall growth in the concentration of clinics, it results in a contraction in the concentration of clinics available to Medicaid patients. Looking to Figure 8(b), we see that growth in Medicaid coverage has a less pronounced impact on Medicaid acceptance. While the conditional relationships are

[35]Since these four states are very small and collectively represent only 10 counties in the one-year ACS sample, results are very similar when these states are included.

[36]We calculate the change in measures of Medicaid program participation from 2012 to 2015 when analyzing responses to growth in Medicaid coverage. Since many Medicaid expansions were announced in 2012, and clinics likely submit their paperwork for Medicaid billing before actively accepting Medicaid patients, using 2012 rather than 2013 as the base year ensures that we capture increases in Medicaid acceptance stemming from anticipatory paperwork.

upward sloping, which is consistent with Medicaid expansions inducing some firms to begin participating in the program, the impacts are likely weakened by the concurrent reductions in clinics in such areas. That is, while the theory predicts that firms should simultaneously enter and stop accepting Medicaid in markets with growing private insurance coverage, firms should exit and begin accepting Medicaid in markets with growing Medicaid coverage. As it does not matter for patients whether a firm that exited (or decided not to open) would have accepted Medicaid conditional on entry, the positive effects of Medicaid expansions on Medicaid-accepting firms are unsurprisingly less pronounced.

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

Finally, we ask whether our findings extend to other providers in the outpatient care market. To do so, we analyze the locations of urgent care centers, another type of on-demand health care clinic, as well as physicians in general practice (GPs). Many GPs practice in outpatient settings that screen patients based on insurance coverage or ability to pay, and urgent care clinics similarly do not provide uncompensated care. As a result, the theory outlined in Section II is expected to generalize to both types of providers.

Information on the locations of all urgent care centers operating in the United States in 2021 come from the NUCR database. The NUCR data are less comprehensive than the Merchant Medicine data and only include information on the year of entry and geo-coded locations for clinics that remained open in 2021.[37] Given this limitation, 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

[37]Figure A16(b) shows the locations of urgent care centers in 2021. There were nearly 13,000 urgent care centers operating in 2021, only 25 percent of which were open in 2010 (see Figure A16(a)). While growth in urgent care centers had spread beyond the largest metropolitan areas by 2021, Table A4 shows that, like retail clinics, urgent care centers were concentrated in counties that were relatively economically prosperous and racially diverse. Notably, nearly every county with a retail clinic in 2016 saw an urgent care center operating in 2021 (see Figure A16(b)).

by different types of insurance in the 2012–2016 and 2017–2021 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.

As shown in Figure 9(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 9(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.

To examine whether our findings extend to outpatient physicians, we consider the location decisions of GPs.[38] As with the analysis of urgent care centers, we focus on cross-sectional

[38] We define GPs to include physicians in general practice, family medicine, and geriatric medicine. We exclude those in pediatrics and internal medicine, as many of these physicians subspecialize and practice in hospital settings. The results are very similar when geriatricians are further excluded.

patterns when considering GP location decisions for two reasons. First, the ACA temporarily required states to raise Medicaid reimbursement rates to Medicare levels for select primary care services in 2013 and 2014, which may have muted the effects of Medicaid expansions on physicians' location decisions during this period. Second, constructing reliable annual panels of physician locations is challenging.[39] Given these concerns, we use data from CMS's Doctors and Clinicians National Downloadable Files for 2016 and 2021 and examine how the number of GPs per 100,000 people at the county level in each year covaries with the share of the population covered by different types of insurance using data from the 2012–2016 and 2017–2021 ACS, respectively.

As with retail clinics and urgent care centers, Figure 9(c) shows that the concentration of GPs increases with the share of the population covered by private insurance and decreases with the share covered by Medicaid. Although this pattern is evident in both 2016 and 2021, the relationships are somewhat less pronounced in 2021. Across these years, counties with the highest rates of private insurance coverage had 15–30 percent more GPs per capita than those with the lowest, while counties with the highest rates of Medicaid coverage had 15–30 percent fewer.

# VI 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 they affect the prices paid to providers. When provider reimbursements differ

[39] Although data from the National Plan and Provider Enumeration System (NPPES) can in principle be used to track provider locations over time, they often lag actual provider moves, complicating measurement over short windows (Currie et al., 2024). In addition, CMS's Doctors and Clinicians National Downloadable File is only available from 2013 onward, midway through our study period.

by patient coverage type—a common feature of two-tiered health insurance systems in both the United States and abroad—insurance expansions that shift patients into programs with regulated prices can lead the concentration of providers to contract. This supply-side contraction occurs when regulated prices are sufficiently low such that firms preferred to serve patients not covered by the program at baseline. Moreover, despite leading to an increase in the concentration of providers, expansions of insurance with market-based prices can reduce the number of providers available to patients in the public system when regulated prices are such that firms were on the margin of not serving the program's beneficiaries at baseline.

We confirm 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—led to an increase in clinic concentration of nearly 25 percent relative to the mean. Moreover, like retail clinics, the concentration of urgent care centers and physicians in general practice (GPs) increases with the share of the population covered by private insurance across the United States, suggesting that provider location patterns more broadly are responsive to private insurance growth.

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.[40] If patients with HDHPs are unlikely to have reached their deductibles, and thus are facing the full spot prices 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 cov-

[40] 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).

erage separately, we find that the positive supply-side effects of private insurance are driven predominately by employer-sponsored coverage.[41] 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 substantially less than private insurers for the same services (Alexander and Schnell, 2024), and lower reimbursement rates are compounded by administrative hassles that providers face when billing the program (Dunn et al., 2024). Despite being very generous for patients, with limited to no cost-sharing, we find that recent growth in Medicaid coverage caused the number of retail clinics 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. Similar patterns are observed among urgent care centers and GPs, suggesting that the findings may extend to the outpatient market more generally.

An important question is whether the heterogeneous supply-side responses to different types of health insurance expansions that we document are efficient from a social planner's perspective. If private insurance expansions lead clinics to enter underserved areas while Medicaid expansions prompt exits (or deter entry) in areas with an oversupply of resources, these responses may simultaneously address access barriers and limit unnecessary service

[41] 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 analogs 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).

provision. However, if private insurance expansions drive entry into well-resourced areas and Medicaid expansions lead providers to exit underserved areas, these patterns will exacerbate inequities in health care access and increase unnecessary services in resource-rich regions. Examining effects across areas with and without primary care shortages at baseline, we find that clinic growth following private insurance expansions is larger in well-resourced areas, while clinic reductions tied to Medicaid expansions are more pronounced in underserved areas.[42] These findings suggest that the observed supply-side responses to insurance expansions contribute to allocative inefficiency and worsen disparities in access to care.

Taken together, our findings demonstrate that not all health insurance is created equal, and policymakers must account for how the supply side responds differently to various types of insurance when considering expansions. Medicaid expansions are often implemented to help low-income populations access adequate health care. However, if providers not mandated to serve all patients exit areas with higher Medicaid coverage, such policies may unintentionally reduce access to certain types of care for all patients. Similarly, while private insurance expansions in two-tiered systems are often promoted for their ability to provide patients willing and able to pay more with faster or higher-quality care, they can also reduce access for those reliant on public insurance.

Importantly, our results suggest a solution: these contractionary effects in two-tiered systems arise only when regulated reimbursement rates for public insurance are low. Raising these rates to market levels would encourage supply-side growth in response to increased public insurance eligibility while reducing providers' incentives to stop serving publicly insured patients as private insurance coverage expands. Thus, while market failures in health insurance markets justify many forms of government intervention, the disruptive impacts of regulated prices on critical market-level forces cannot be overlooked.

[42] 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 Figure 5 separately among these two groups of counties. Results for these analyses, both unconditional and conditional on baseline Medicaid reimbursement rates, are shown in Figure A17.

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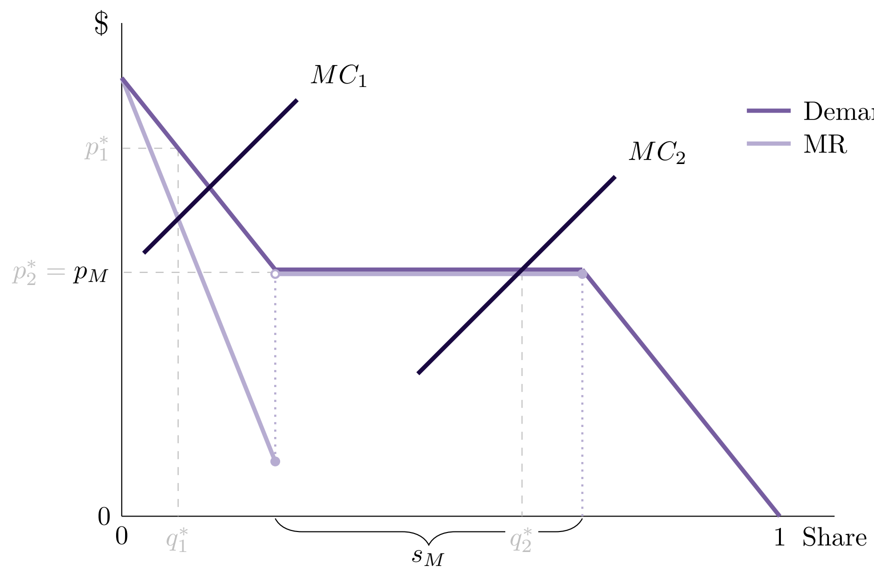
**United States Bureau of the Census**, "Quarterly Workforce Indicators," Technical Report, United States Bureau of the Census [distributor], Washington, DC, USA 2013.

**Wang, X.**, "The Welfare Effects of WIC Purchasing in the Infant Formula Market," SSRN Working Paper 2024.

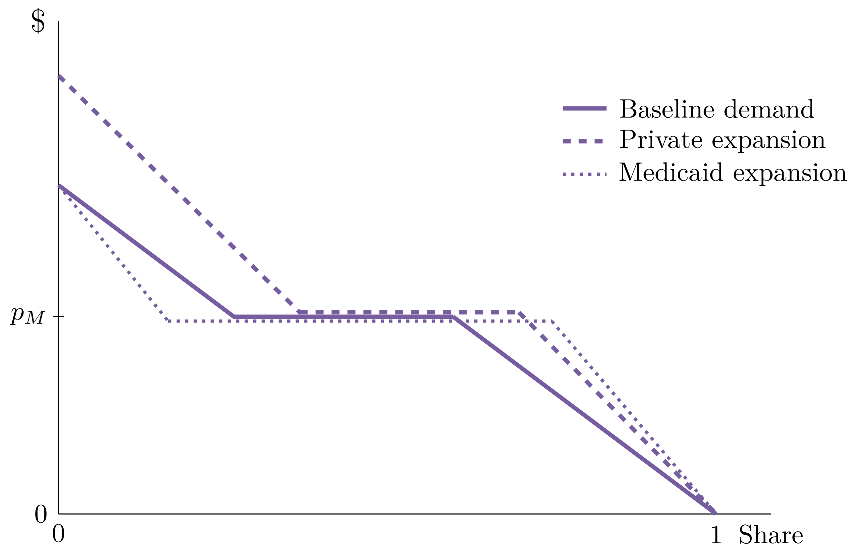
# VII Figures

Figure 1: Clinic in market with both administered and market-based prices

(a) Firm's baseline problem

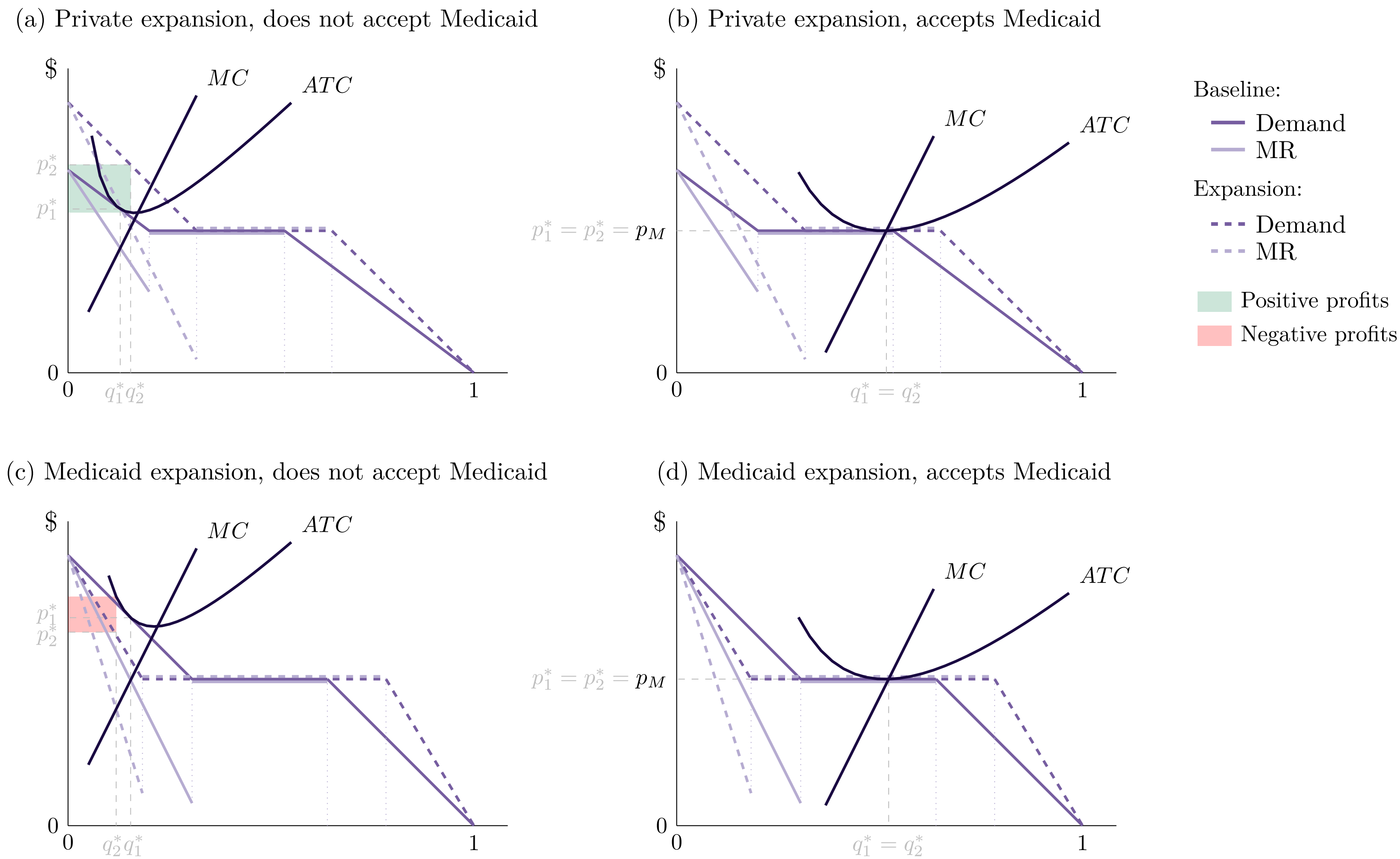



(b) Impacts of insurance expansions on demand




Notes: The above figures consider a firm in a market with both regulated 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 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 zero or two intersections are shown in Figure A1. Subfigure (b) shows how the demand curve changes under expansions of private insurance (dashed line) and Medicaid coverage (dotted line).

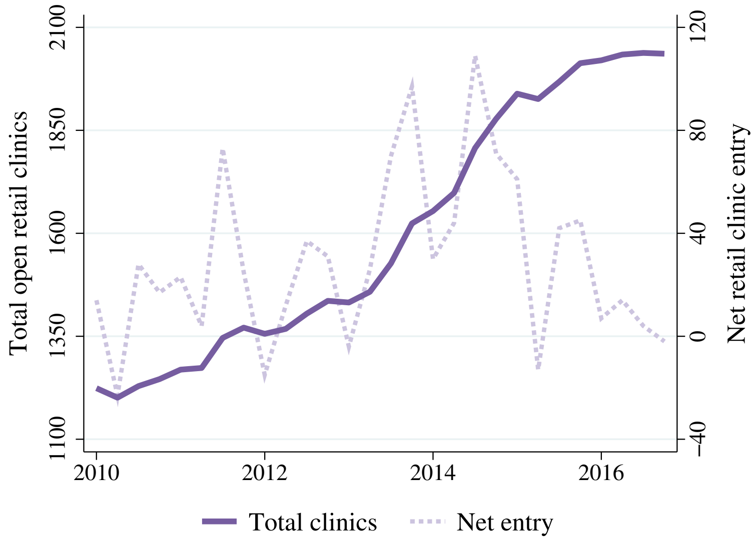
Figure 2: Effects of insurance expansions on clinic profits



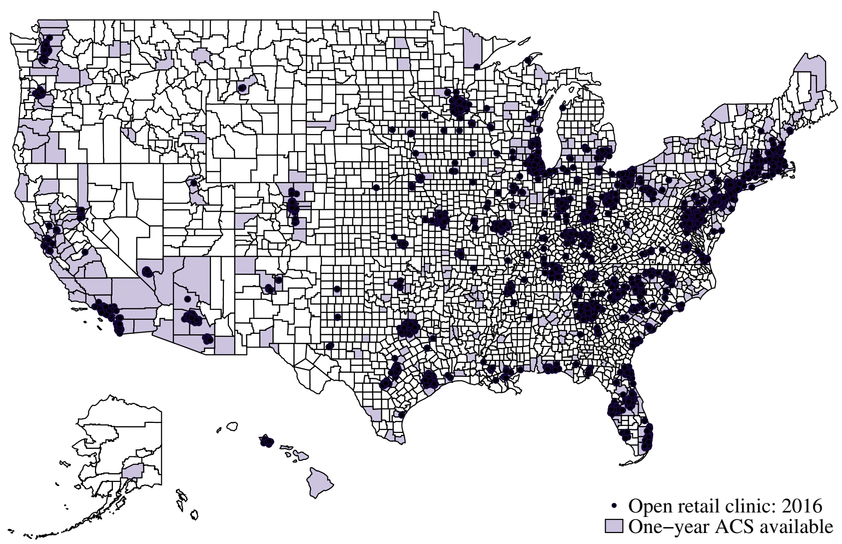

Notes: The above figures show how expansions of private insurance (subfigures (a) and (b)) and Medicaid coverage (subfigures (c) and (d)) affect firm profits, both when the firm accepted Medicaid patients at baseline (subfigures (b) and (d)) and when the firm did not (subfigures (a) and (c)). As shown 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 A2 shows the impacts of private and Medicaid expansions when there is no intersection between marginal costs and marginal revenue at baseline.

Figure 3: Retail clinics across the United States

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



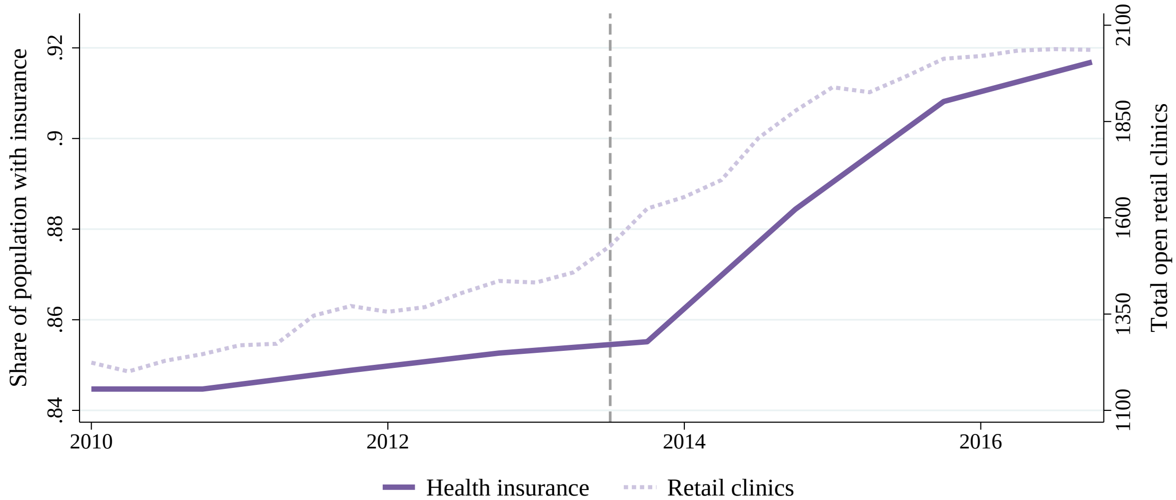

(b) Locations: 2016



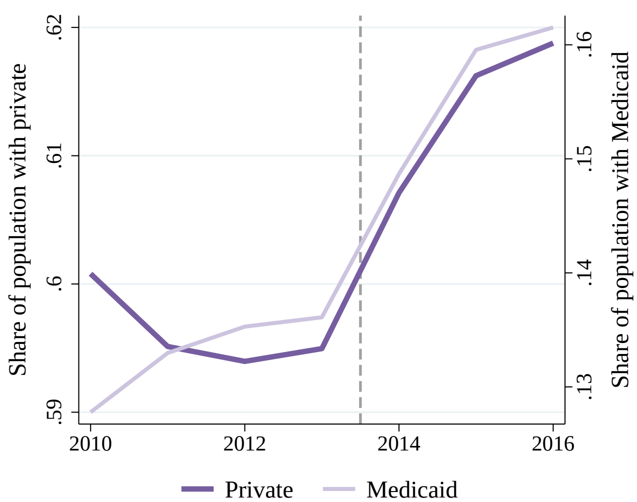

Notes: The above figures show the number and locations of retail clinics across the United States. Subfigure (a) shows the total number of retail clinics (dark, solid line) and net entry (light, dashed line) quarterly from 2010 to 2016. "Net entry" refers to the total number of openings net of the total number of closings in a given quarter; refer to Figure A4 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 4: Changes in health insurance coverage: 2010–2016

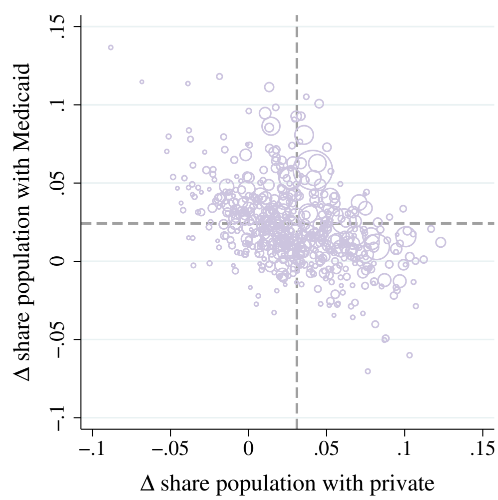
(a) Health insurance and retail clinics




(b) Share with private or Medicaid coverage



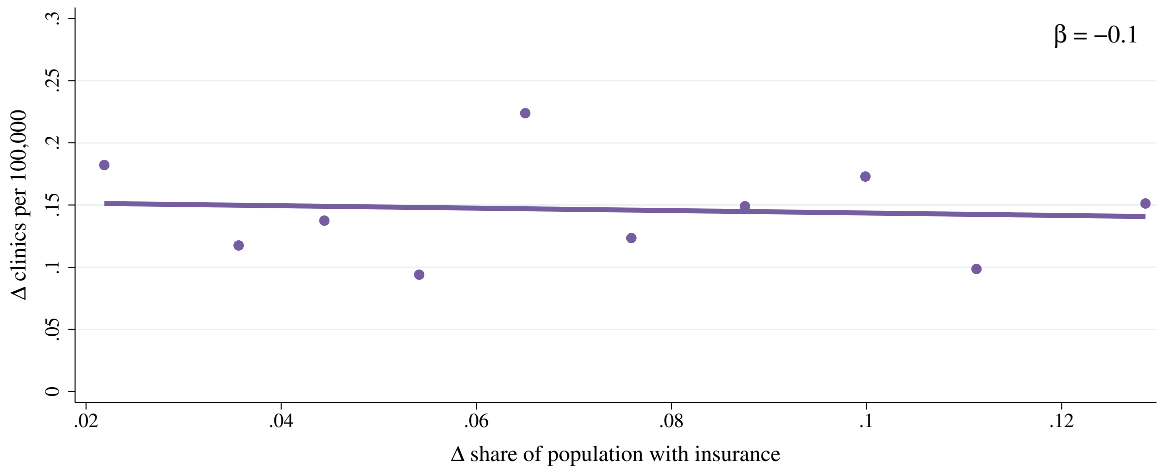

(c) Changes in private vs. changes in Medicaid



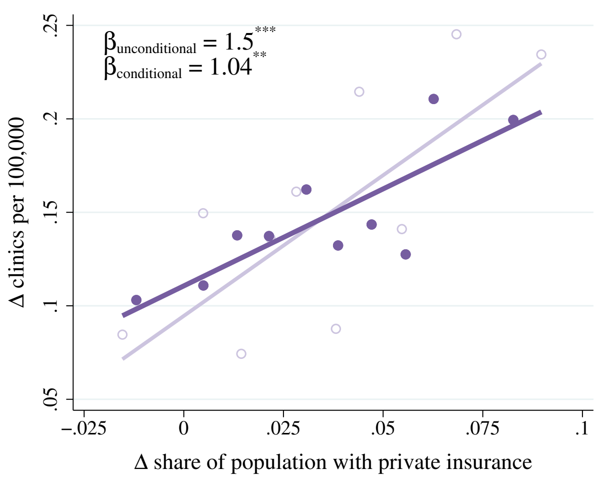

Notes: The above figures show changes in the share of the population with different types of health insurance from 2010 to 2016. Subfigure (a) displays the annual share of the population with health insurance of any type (dark, solid line) and the quarterly number of retail clinics (light, dashed line). 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 5: Changes in retail clinic presence versus changes in health insurance: 2013–2015

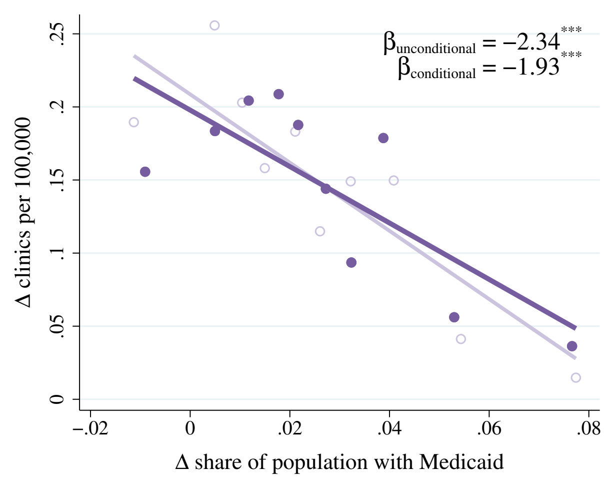
(a) Health insurance of any type




(b) Private insurance



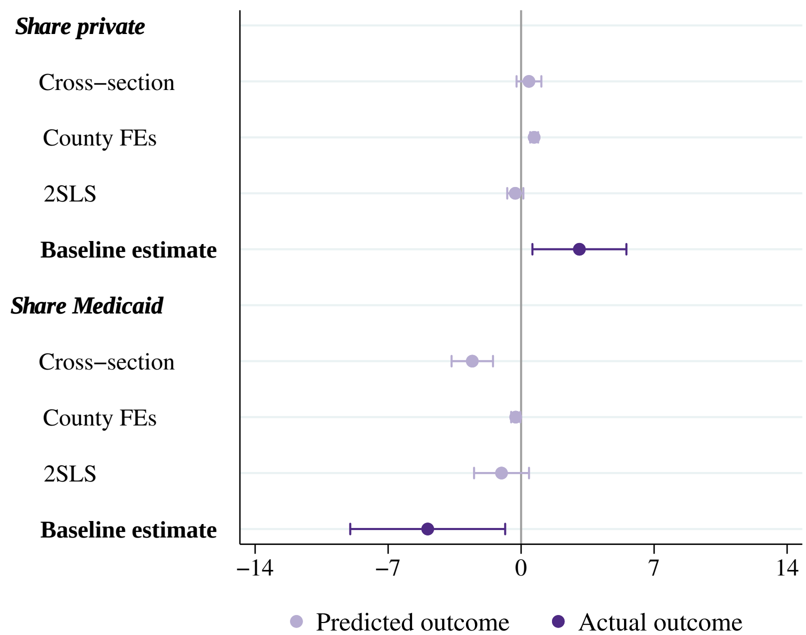

(c) Medicaid coverage




Unconditional • Conditional on other insurance changes

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.

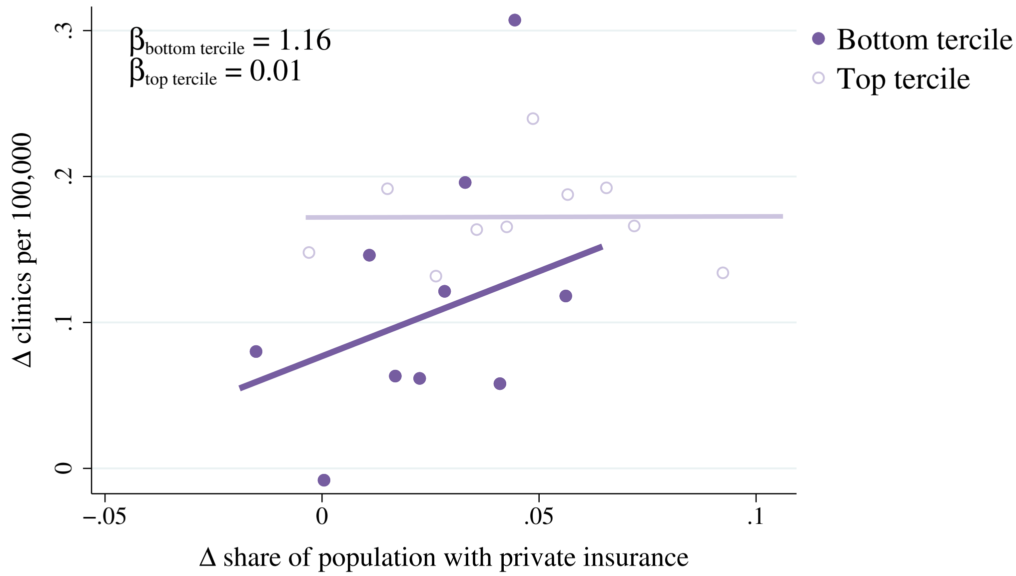
Figure 6: Effects of insurance on retail clinic penetration predicted using potential confounders



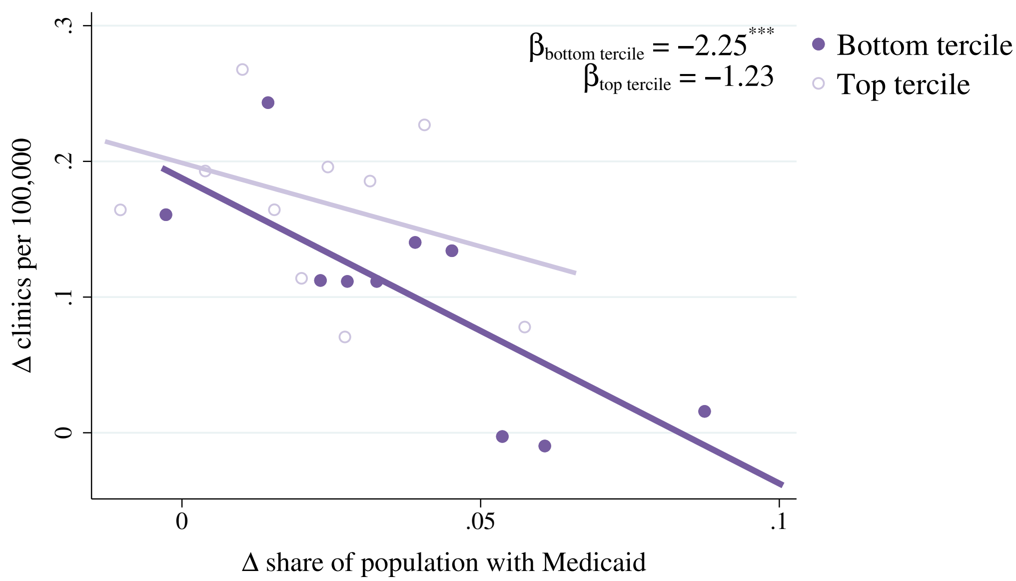

Notes: The above figure shows output from estimation of two-step balancing regressions. We first predict retail clinic concentration using local socio-demographics. We then estimate the specification denoted on the y-axis using this measure of predicted retail clinic penetration as the dependent variable. "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

(a) Private insurance



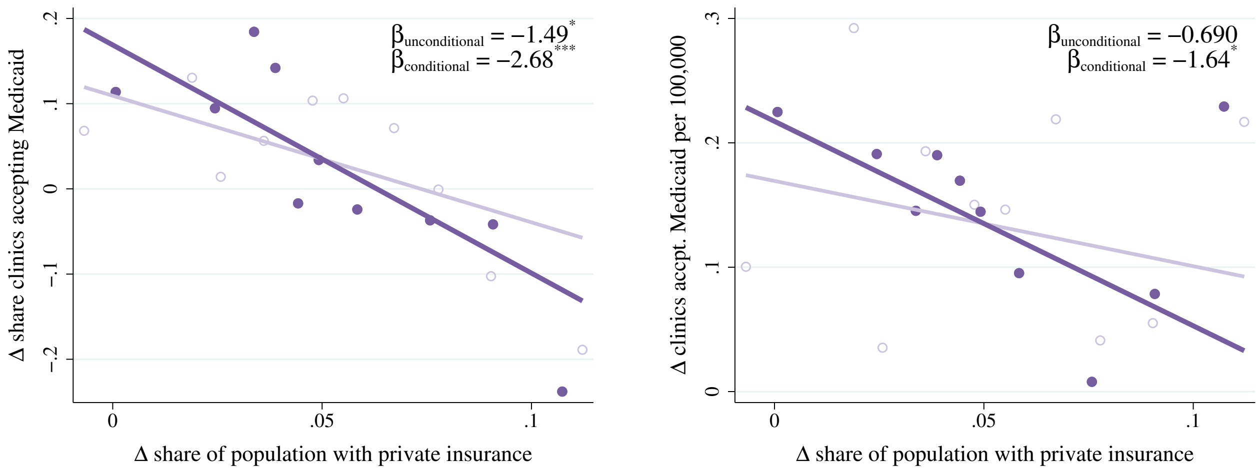

(b) Medicaid coverage



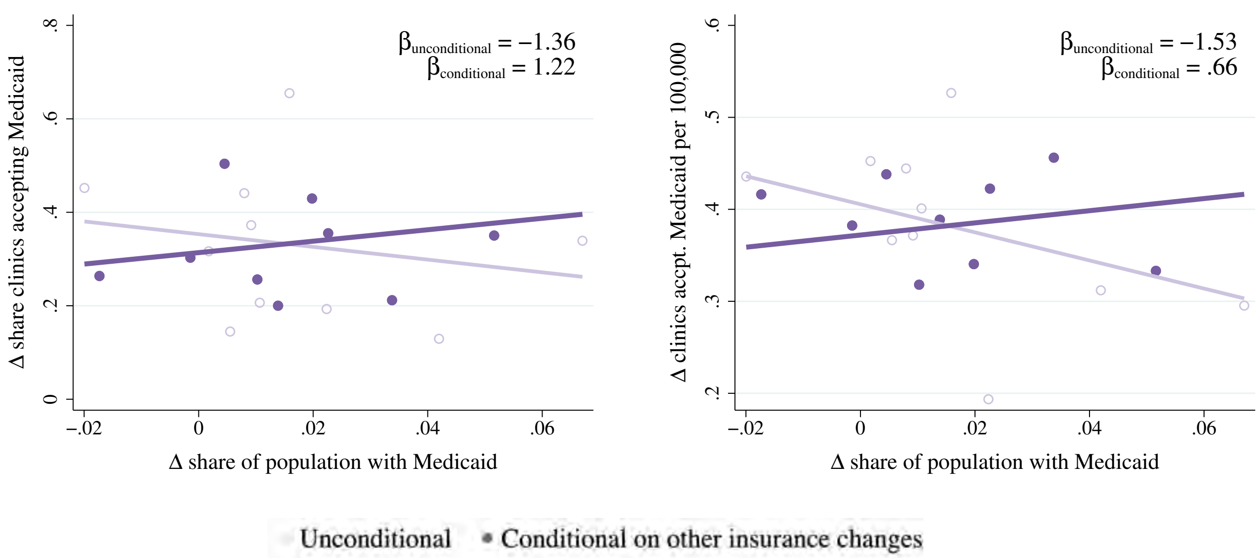

Notes: The above figures show how county-level changes in retail clinics per 100,000 people from 2013 to 2015 covary with county-level changes in the share of the population with private insurance coverage (subfigure (a)) and Medicaid coverage (subfigure (b)) over the same period. These relationships 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, 2024).

Figure 8: Changes in Medicaid acceptance versus changes in insurance

(a) Private insurance



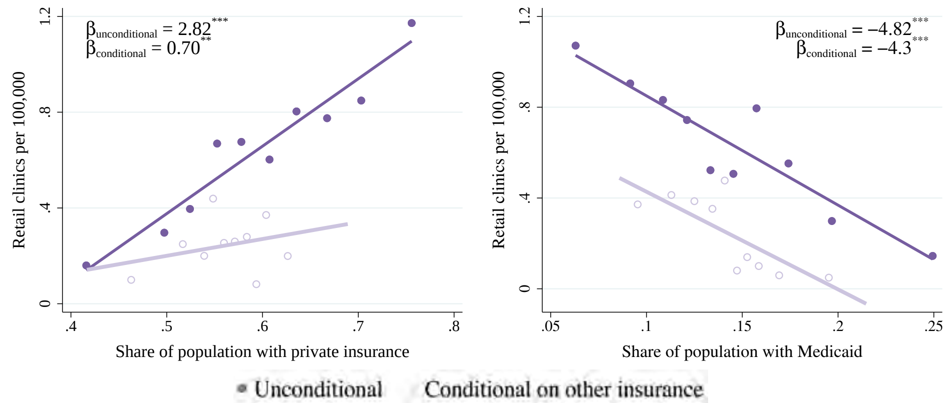

(b) Medicaid coverage



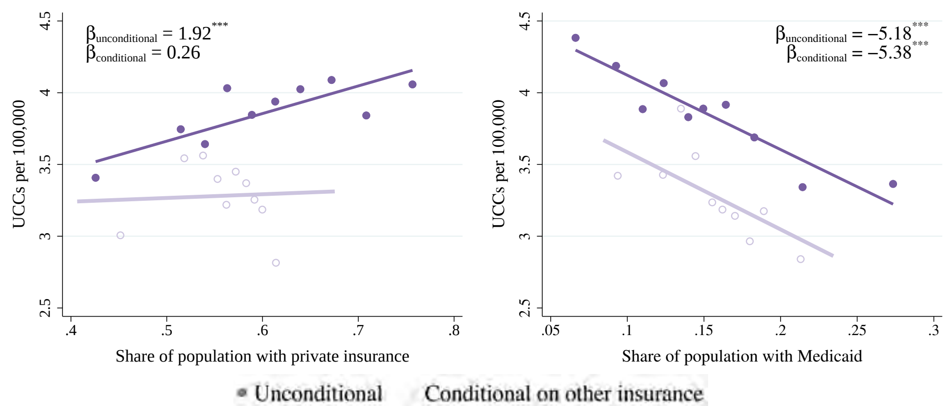

Notes: The above figures show how county-level changes in the share of retail clinics accepting Medicaid (left subplots) and the number of retail clinics that accept Medicaid per 100,000 people (right subplots) covary with county-level changes in the share of the population with private insurance coverage (subfigure (a)) and Medicaid coverage (subfigure (b)) from 2013 to 2015. 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 Medicaid acceptance come from FOIA requests; and data on health insurance come from the one-year ACS.

Figure 9: Cross-sectional relationships between insurance and provider concentration

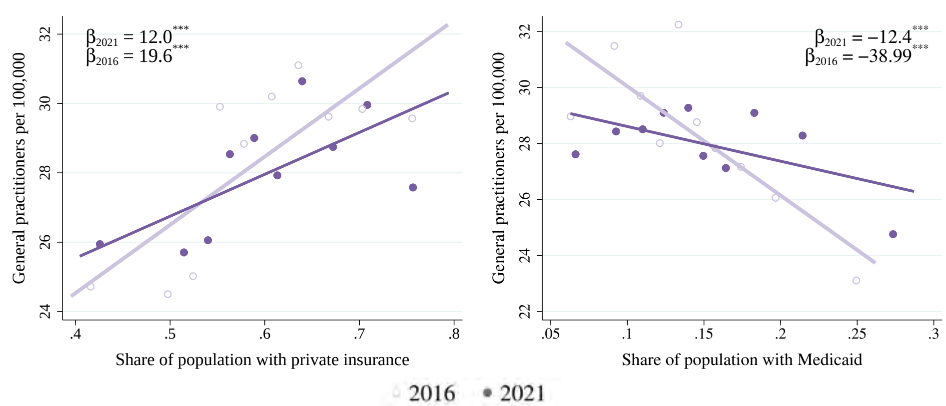
(a) Retail clinics: 2016




(b) Urgent care centers: 2021




(c) Physicians in general practice (GPs): 2016, 2021




Notes: The above figures show how the county-level number of retail clinics in 2016 (subfigure (a)), urgent care centers in 2021 (subfigure (b)), and general practice physicians in 2016 and 2021 (subfigure (c)) per 100,000 people covary with the county-level share of the population with private insurance coverage (left subplots) and Medicaid coverage (right subplots). Subfigures (a) and (b) show 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); subfigure (c) shows the unconditional relationship. Data on retail clinics come from Merchant Medicine, data on urgent care centers come from the NUCR database, data on physicians in general practice come from CMS, and data on health insurance come from the five-year ACS.

# VIII 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) |
| **a. Retail clinics** | | | |
| *2016* | | | |
| Open clinics | 5.51 | 0 | |
| Clinics per 100,000 | 1.11 | 0 | |
| *2010–2016* | | | |
| Openings | 3.66 | 0.10 | |
| Closings | 1.39 | 0.32 | |
| Share ever clinic | 1.00 | 0.19 | |
| **b. County characteristics (2016)** | | | |
| *Basic demographics* | | | |
| 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 |
| *Income and education* | | | |
| 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 |
| *Health insurance* | | | |
| 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** | 1.214** |
| | | (0.453) | (0.488) |
| $\times\ \mathbb{1}\{\Delta Medicaid > median\}$ | | | -0.938 |
| | | | (0.599) |
| Share Medicaid | | -1.279*** | 0.540 |
| | | (0.448) | (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

| | Any health insurance | | By insurance type | | |
|---|---|---|---|---|---|
| Dependent variable: | Share insured (1) | Retail clinics per 100,000 (2) | Share private (3) | Share Medicaid (4) | Retail clinics per 100,000 (5) |
| **a. First stage** | | | | | |
| $Post_t \times Employed_c^{2013}$ | 0.132*** | | 0.198*** | -0.056*** | |
| | (0.027) | | (0.022) | (0.017) | |
| $Post_t \times [138 - 400\%\ FPL]_c^{2013}$ | 0.545*** | | 0.369*** | 0.089 | |
| | (0.086) | | (0.066) | (0.071) | |
| $Post_t \times Expansion_s \times [< 138\%\ FPL]_c^{2013}$ | 0.296*** | | -0.035 | 0.321*** | |
| | (0.111) | | (0.083) | (0.050) | |
| **b. Two-stage least squares** | | | | | |
| 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 | X |
| Year fixed effects | X | X | X | X | X |
| Demographic controls | X | X | X | X | X |
| Observations | 3,869 | 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 | 0.612 |
| 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 | 2SLS | OLS | 2SLS |
| | (1) | (2) | (3) | (4) |
| **a. Openings per 100,000** | | | | |
| Share insurance | 0.156 | 2.227 | | |
| | (0.194) | (1.474) | | |
| Share private | | | 0.316 | 3.067* |
| | | | (0.211) | (1.799) |
| Share Medicaid | | | -0.120 | -0.335 |
| | | | (0.238) | (2.665) |
| **b. Closings per 100,000** | | | | |
| Share insurance | -0.072 | -0.567 | | |
| | (0.118) | (0.470) | | |
| Share private | | | -0.163 | -1.286** |
| | | | (0.131) | (0.606) |
| Share Medicaid | | | 0.112 | 1.908* |
| | | | (0.142) | (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 A3 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

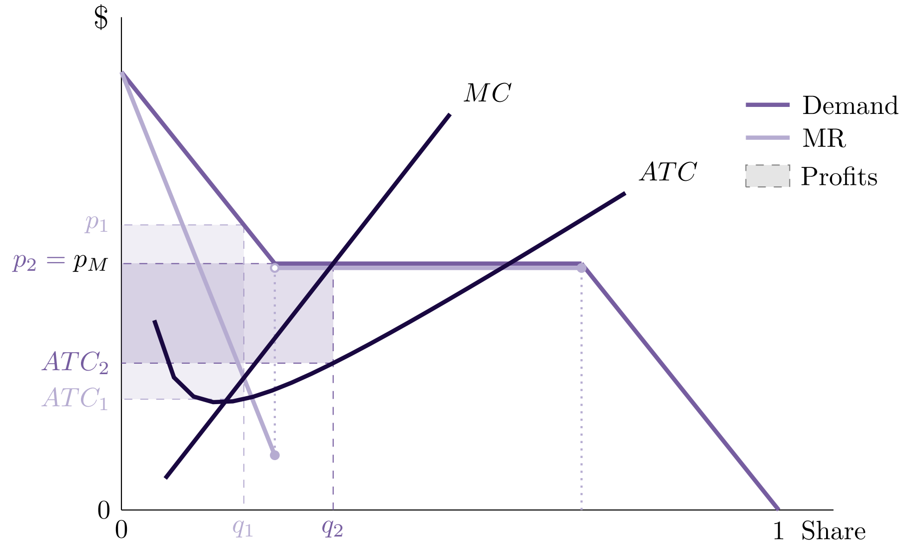
The Expansionary and Contractionary
Supply-Side Effects of Health Insurance

*Geddes and Schnell (2025)*

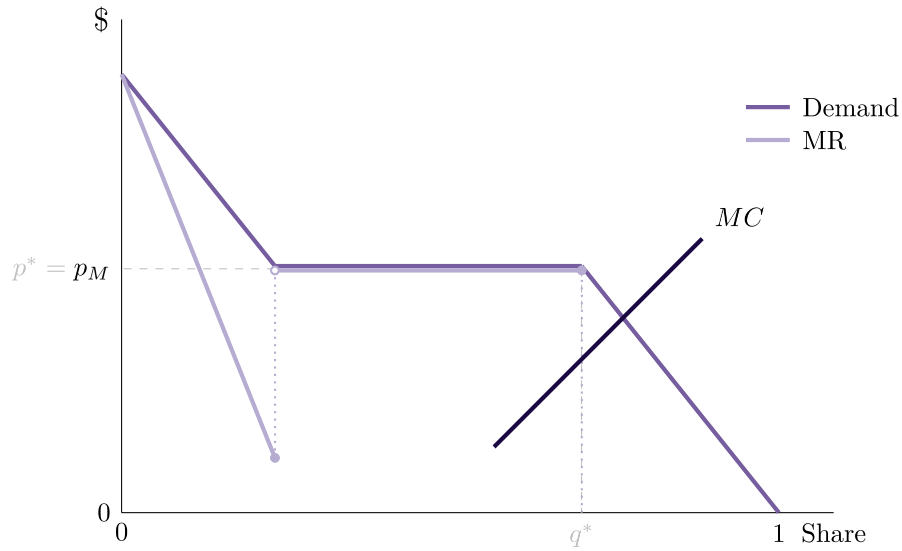
# A Supplementary figures

Figure A1: Additional solutions to firm's baseline problem

(a) Two intersections



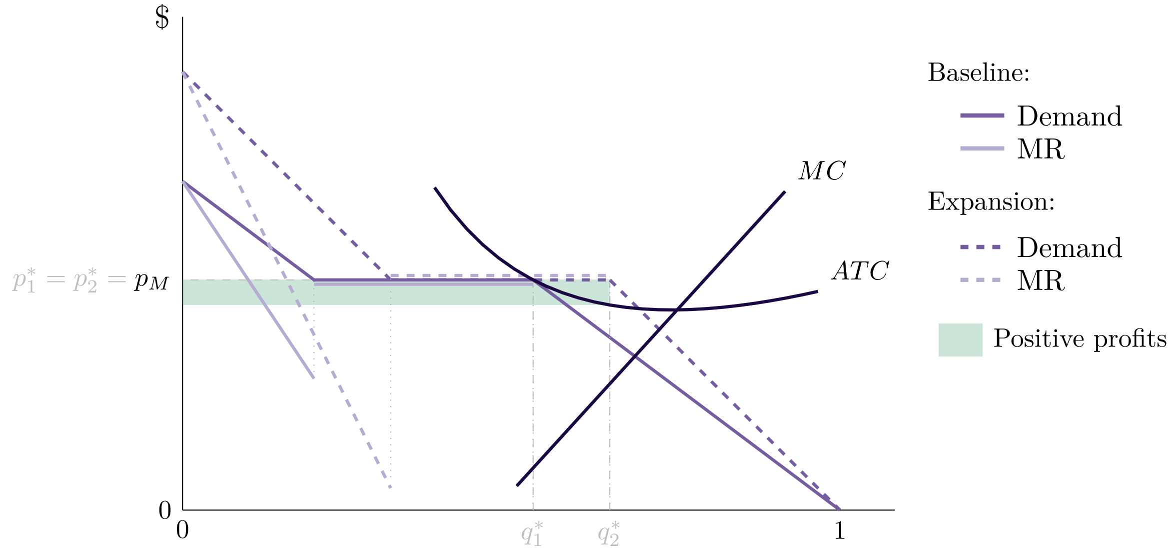

(b) Zero intersections



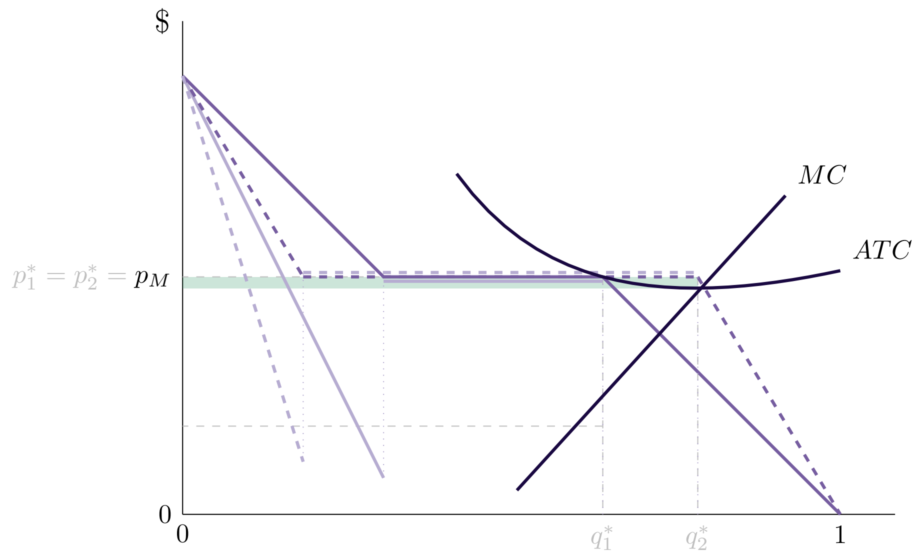

Notes: The above figures show how prices and quantities are determined when marginal revenue and marginal costs intersect twice (subfigure (a)) or not at all (subfigure (b)). As in Figure 1, the total demand curve (dark purple line) and associated marginal revenue curve (light purple line) are for a firm in a market with both regulated and market-based prices. The demand curve is perfectly elastic at the 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 A2: Effects of expansions when no intersection between marginal revenue and costs

(a) Private expansion



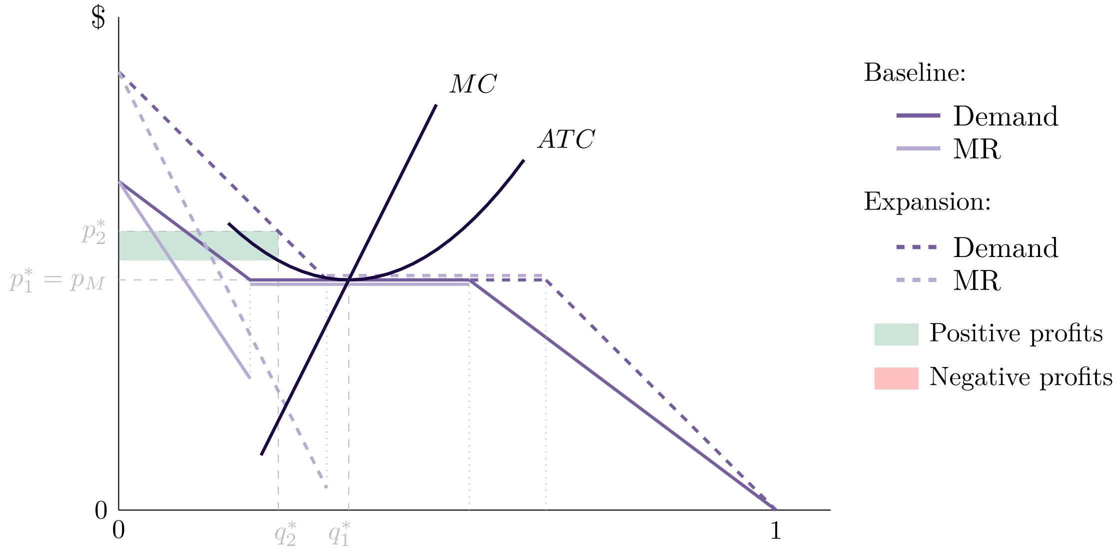

(b) Medicaid expansion



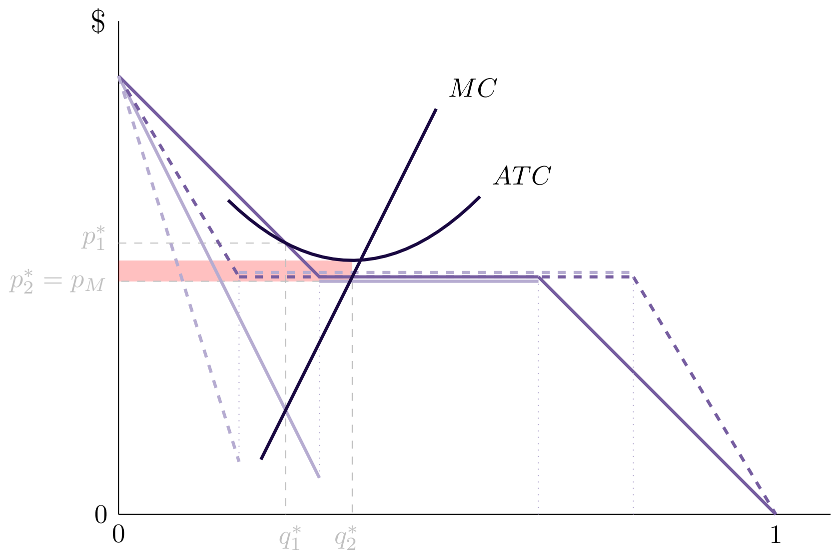

Notes: The above figures show how expansions of private insurance (subfigure (a)) and Medicaid coverage (subfigure (b)) affect firm profits when there is no intersection between marginal costs and marginal revenue at baseline. As shown in the subfigures, both private and Medicaid expansions tend to increase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

Figure A3: Effects of insurance expansions on Medicaid acceptance

(a) Private expansion




(b) Medicaid expansion



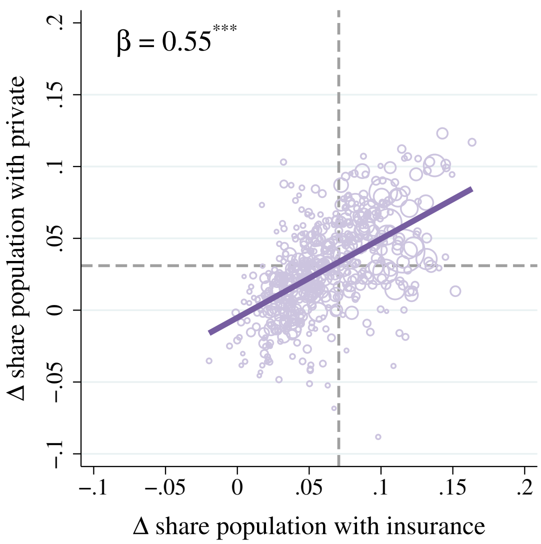

Notes: The above figures show how expansions of private insurance (subfigure (a)) and Medicaid coverage (subfigure (b)) affect firm profits and Medicaid acceptance. Recall from Figure 1(b) and Figures 2(a) and (b) that private expansions increase firm profits for firms that were not accepting Medicaid at baseline and for firms who were accepting Medicaid but were near the point of indifference. As shown in subfigure (a) above, firms that accepted Medicaid at baseline but saw an increase in their profits following a private insurance expansion will stop accepting Medicaid coverage. Further recall from Figure 1(b) and Figure 2(c) that Medicaid expansions decrease firm profits for firms that were not accepting Medicaid at baseline. As shown in subfigure (b) above, firms that did not accept Medicaid at baseline but were near the point of indifference will start accepting Medicaid coverage following a Medicaid expansion.

Figure A4: Retail clinic openings and closings: 2010–2016

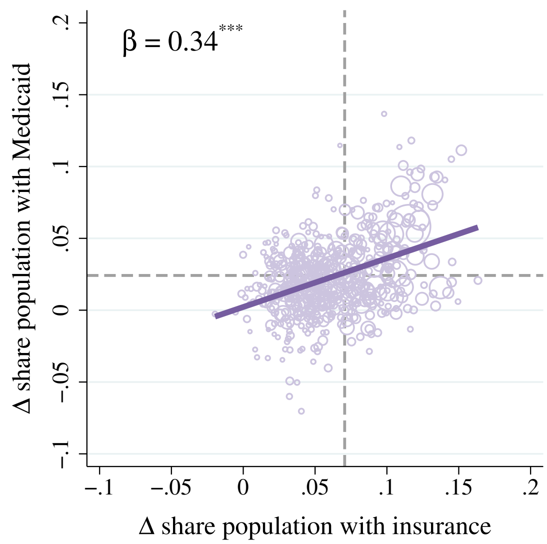

Notes: The above figure shows the number of retail clinic openings (dark, thick line) and closings (light, thin line) quarterly from 2010 to 2016. The dashed horizontal lines denote the quarterly averages of each measure over the entire sample period. Refer to Figure 3 for the total number of open retail 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 A5: Correlation between changes in insurance types: 2013–2015

(a) Private versus all insurance



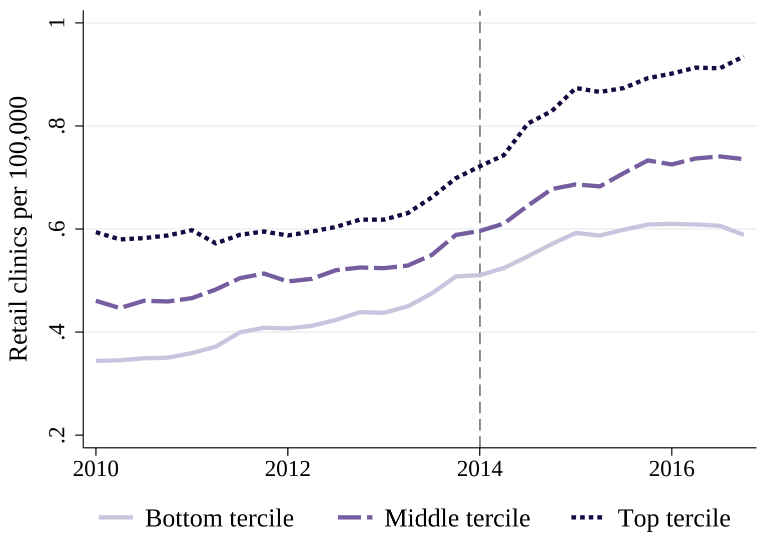

(b) Medicaid versus all insurance



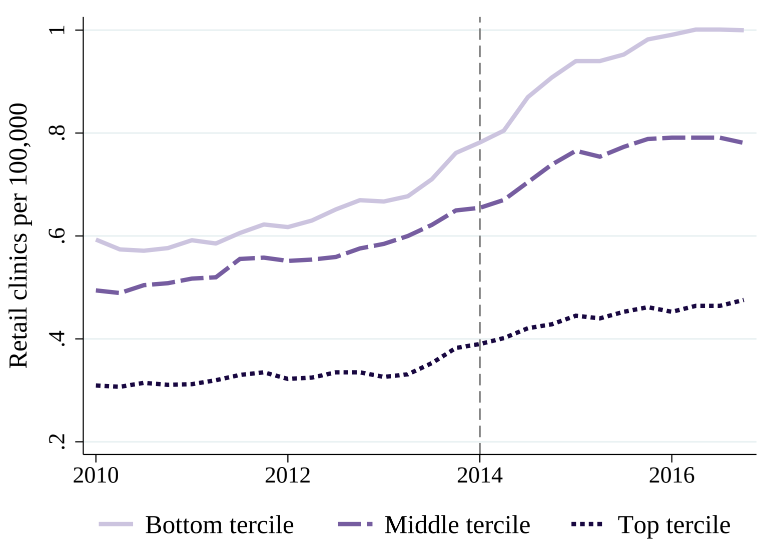

Notes: The above figures show how county-level changes from 2013 to 2015 in the share of the population with private insurance (subfigure (a)) and the share of the population covered by Medicaid (subfigure (b)) covary with county-level changes in the share of population with health insurance coverage 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 4(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 A6: Retail clinic penetration by changes in insurance: 2010–2016

(a) Change in share private (2013–2015)



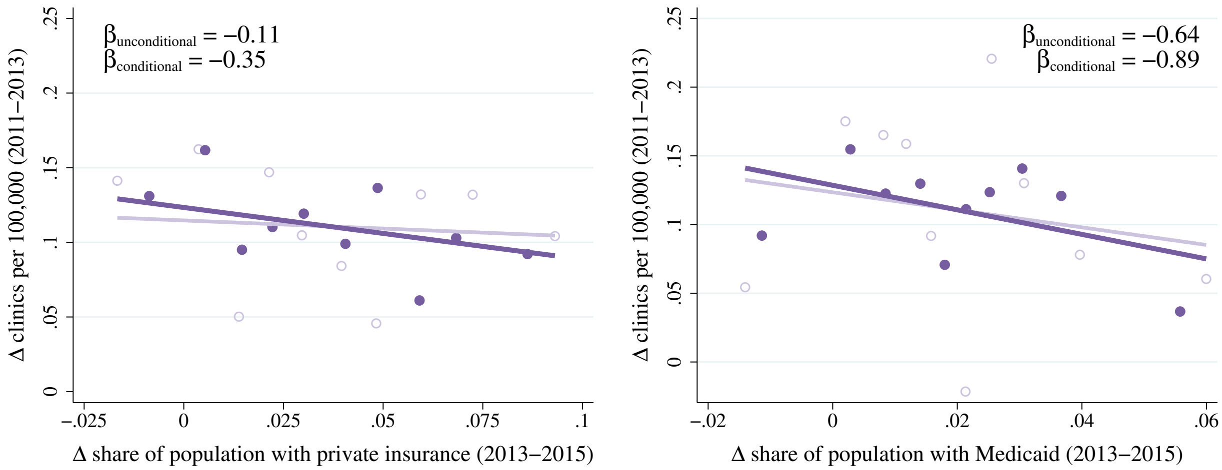

(b) Change in share Medicaid (2013–2015)



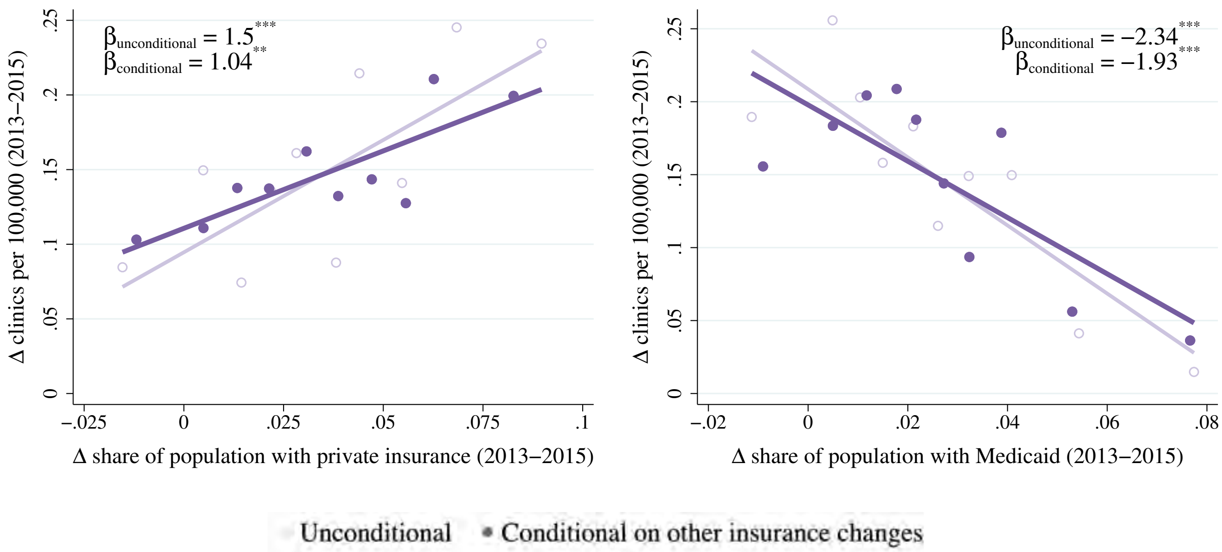

Notes: The above figures show the population-weighted average number of retail clinics per 100,000 people at the county-quarter level from 2010 to 2016 by changes in the share of the population covered by private insurance (subfigure (a)) and Medicaid (subfigure (b)) from 2013 to 2015. In subfigure (a), 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 A7: Changes in retail clinic presence versus changes in health insurance

(a) Change in retail clinic presence: 2011–2013

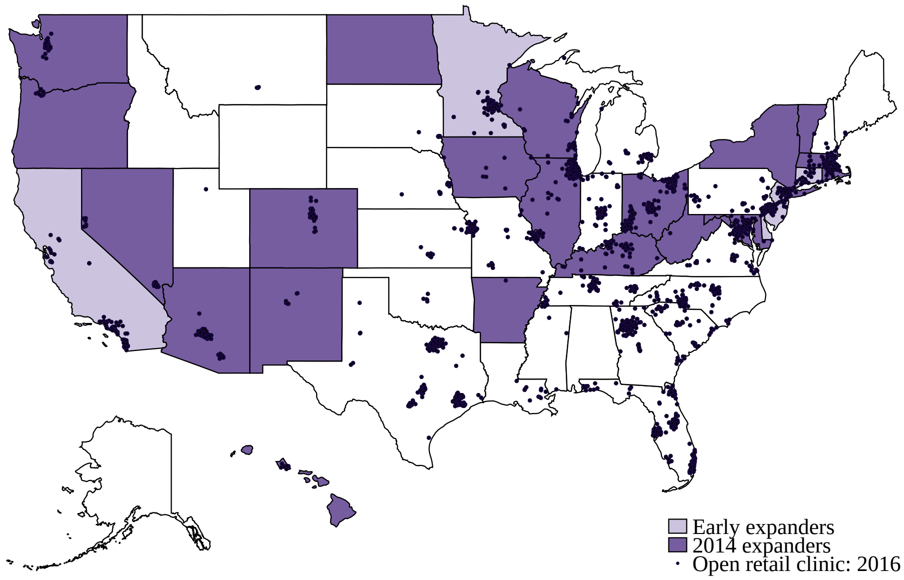



(b) Change in retail clinic presence: 2013–2015




Notes: The above figures show how county-level changes in retail clinics per 100,000 people from 2011 to 2013 (subfigure (a)) and from 2013 to 2015 (subfigure (b)) covary with county-level changes in the share of the population with private insurance coverage (left subplots) and Medicaid coverage (right subplots) from 2013 to 2015. 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. Subfigure (b) reproduces figures first shown in Figure 5(b) and (c). 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.

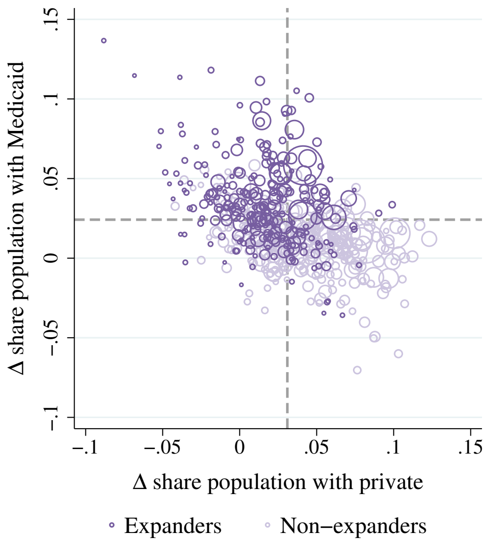
Figure A8: Medicaid expansion status and retail clinic locations



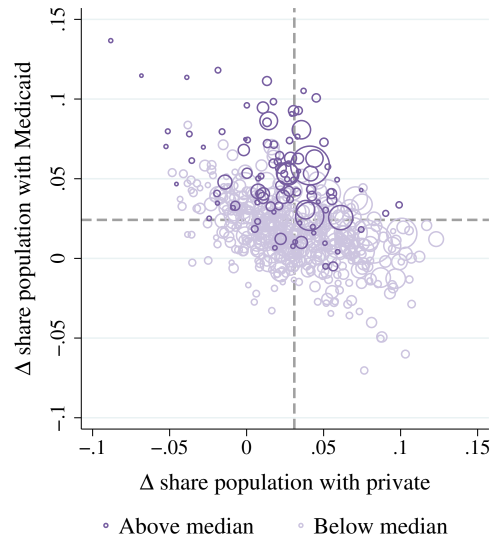

Notes: The above figure shows the locations of open retail clinics in 2016 (geo-coded dots) and state-level Medicaid expansions by 2014 (shaded states) across the United States. Data on retail clinics come from Merchant Medicine; data on Medicaid expansions come from the Kaiser Family Foundation.

Figure A9: Changes in insurance types (2013–2015) by instrument components

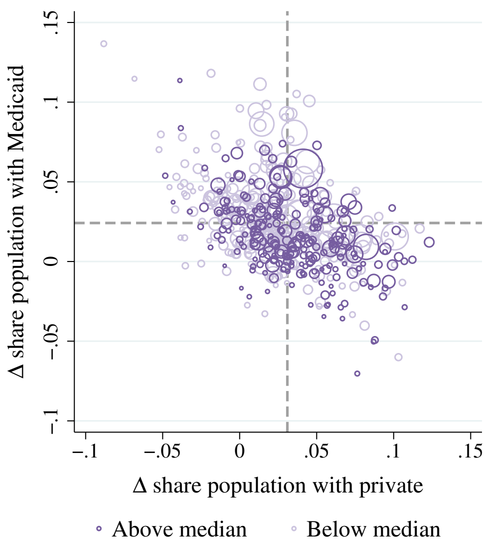
(a) Medicaid expansions



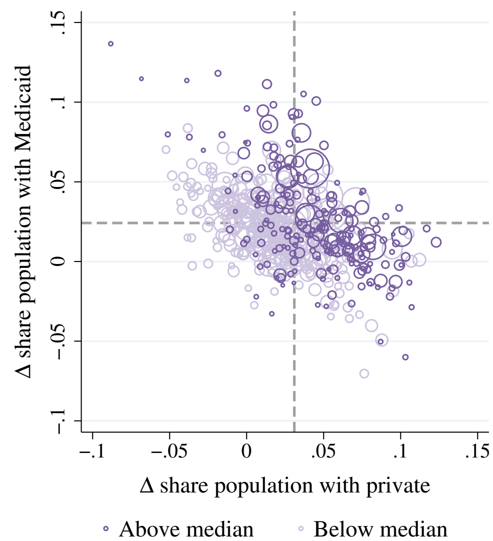

(b) Medicaid expansions + share ≤138% FPL




(c) Share employed

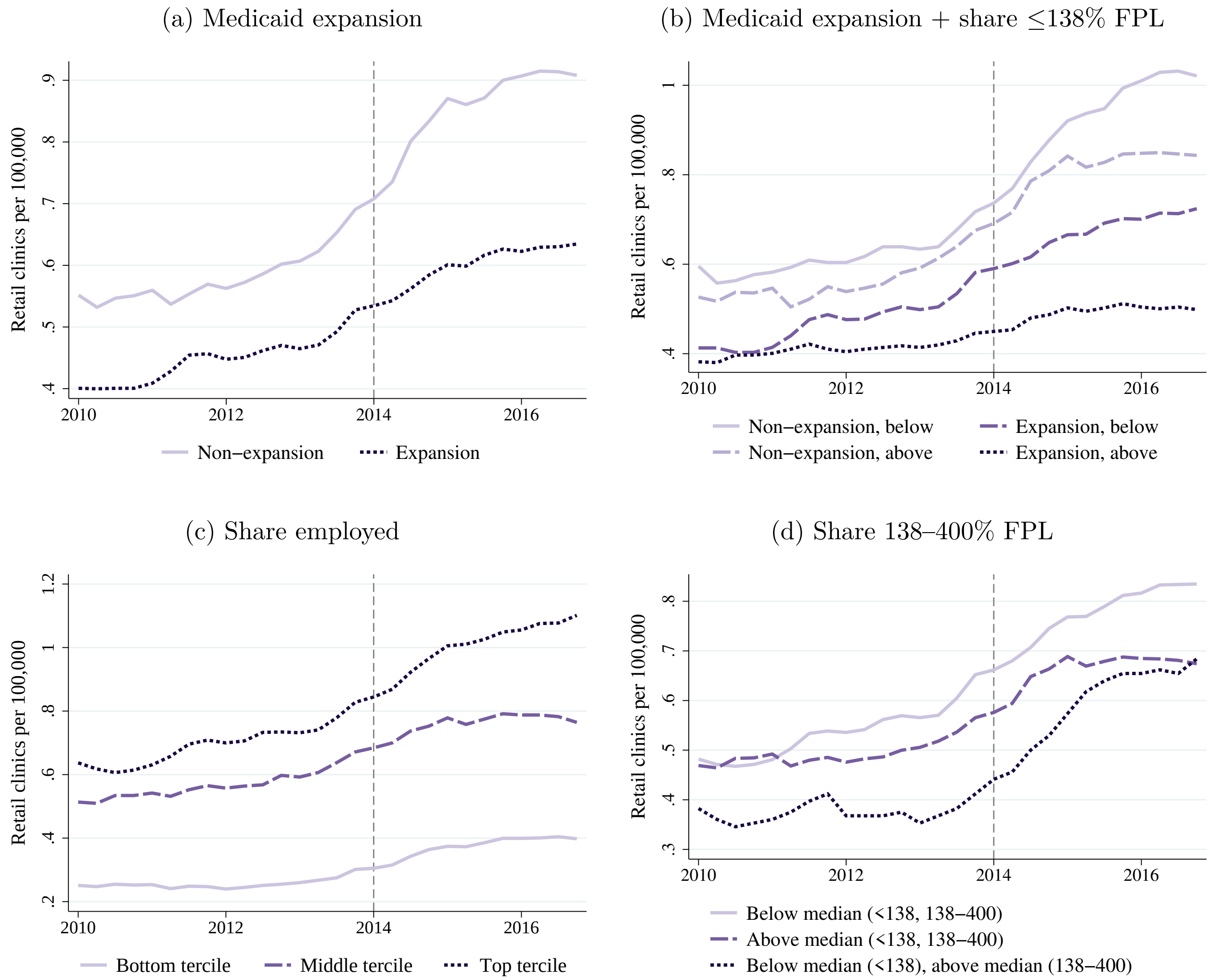



(d) Share 138–400% FPL




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.

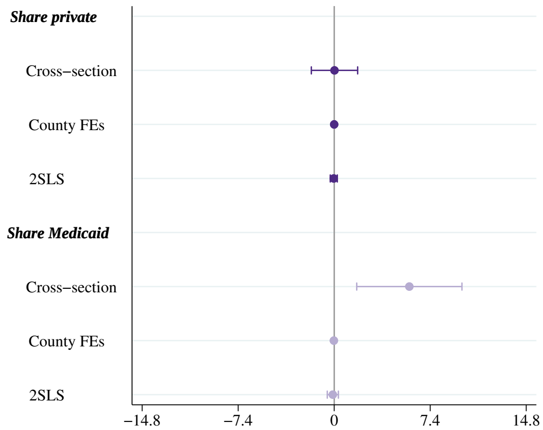
Figure A10: Retail clinic penetration by instrument components: 2010–2016



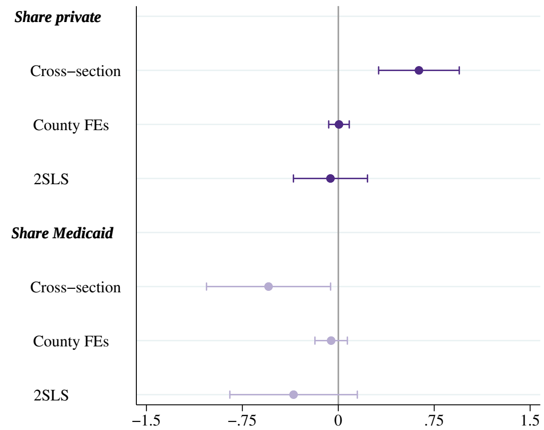

Notes: The above figures show the population-weighted average number of retail clinics per 100,000 people at the county-quarter level from 2010 to 2016 by different instrument components. In subfigure (a), the dark, dotted (light, solid) line considers counties in states that expanded (did not expand) 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 firms with 50–99 employees 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 A11: Balancing regressions on select potential confounders

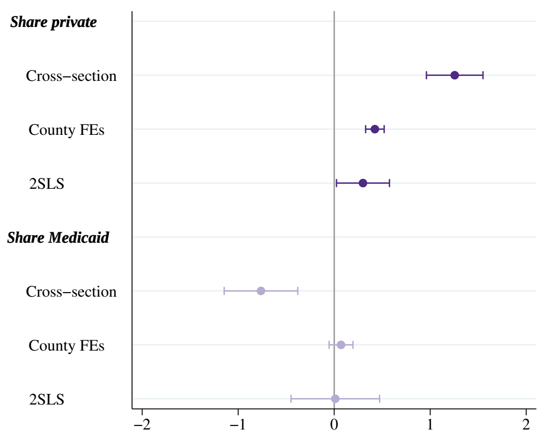
(a) Population density



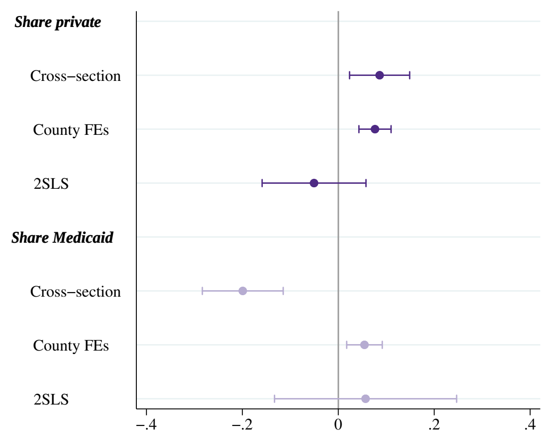

(b) Share White




(c) ln(Median income)



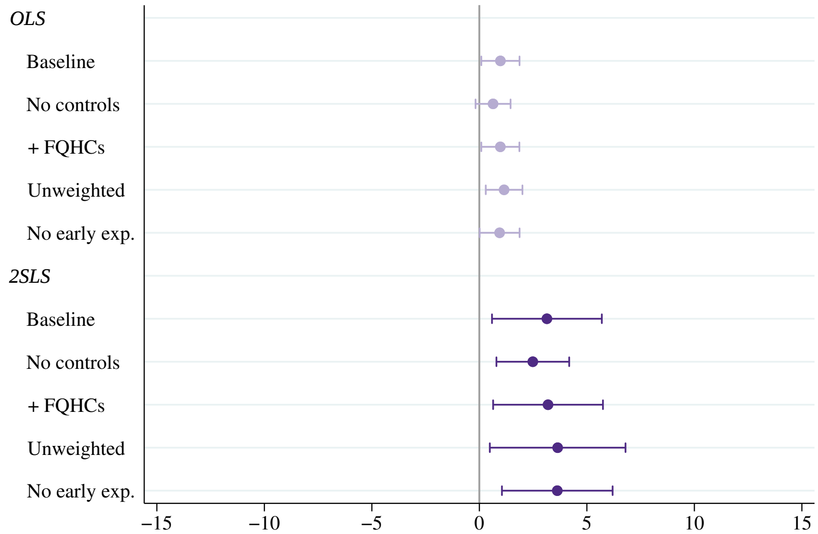

(d) Share employed



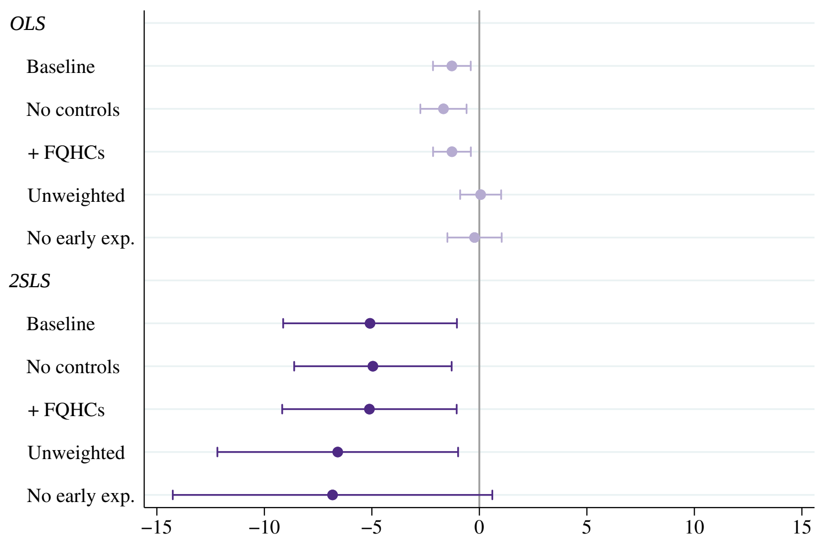

Notes: The above figures show output from estimation of the specifications denoted on the y-axis with different potential confounders as the dependent variable. "Cross-section" refers to estimation of equation (2) without county fixed effects, "county FEs" refers to estimation of equation (2), and "2SLS" 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 A12: Effects of insurance on retail clinic penetration: robustness

(a) Share private



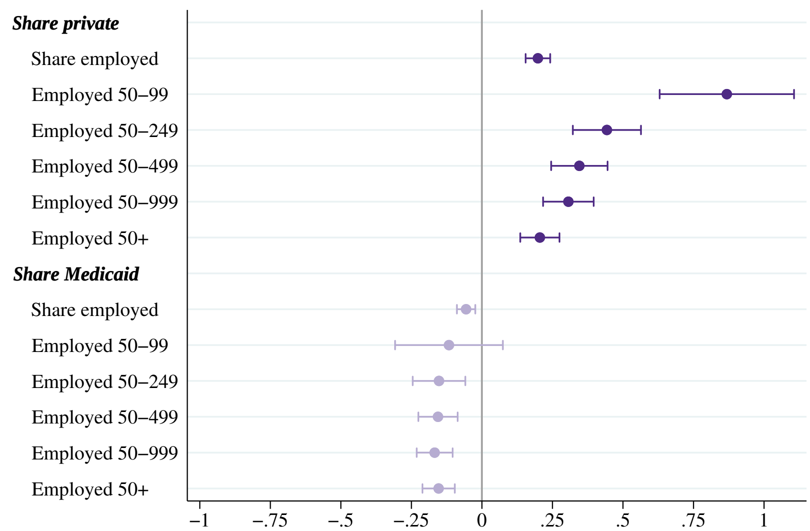

(b) Share Medicaid



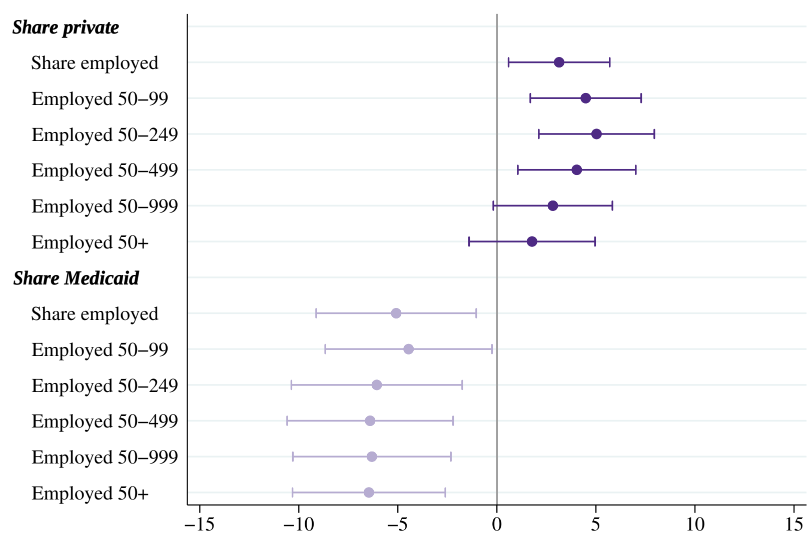

Notes: The above figures show the sensitivity of our baseline estimates to alternative empirical specifications. The top panel of each subfigure ("OLS") shows output from estimation of equation (2), and the bottom panel ("2SLS") shows output from estimation of equation (4). As outlined in these equations, 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 A13: Effects of insurance on retail clinic penetration: alternative employment shares

(a) First stage



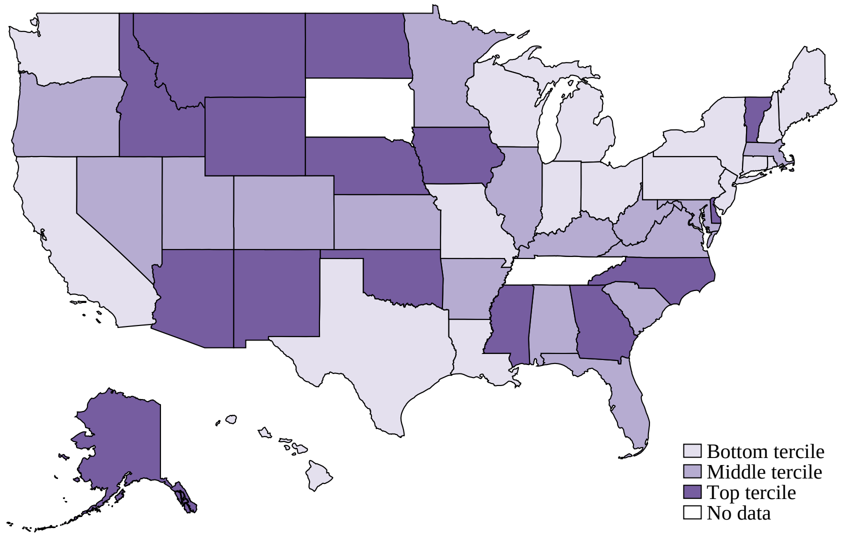

(b) Two-stage least squares



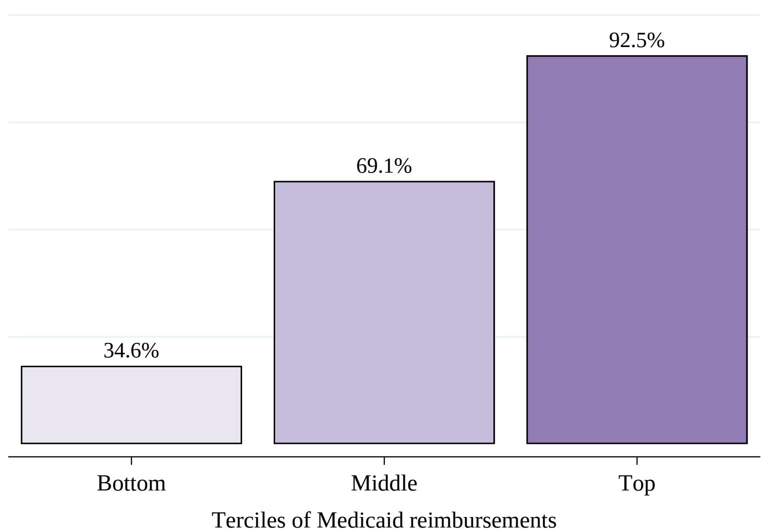

Notes: The above figures show the sensitivity of our baseline estimates to using alternative employment shares when constructing the employment instrument. Subfigure (a) shows output from estimation of equation (3); the outcome variable is either the share of the population with private insurance (top 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 A14: Medicaid reimbursement rates and coverage acceptance

(a) Medicaid reimbursement rates: 2010



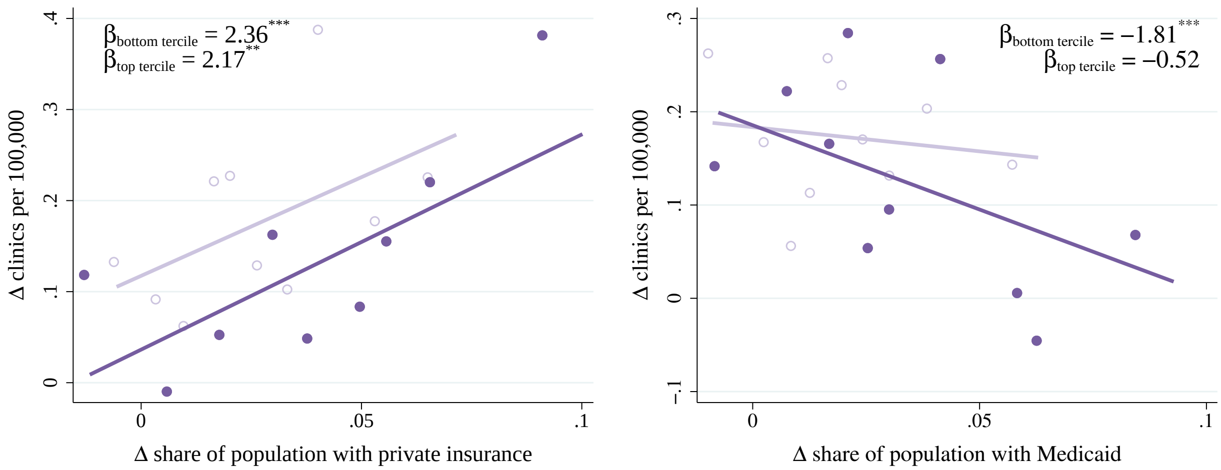

(b) Medicaid acceptance (2020) versus reimbursements (2015)



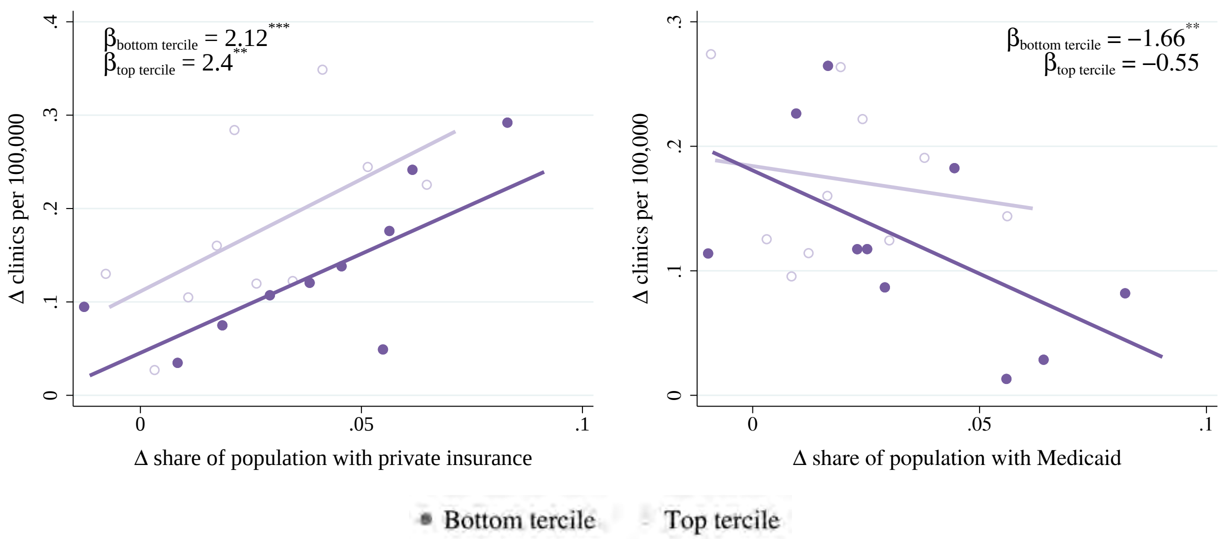

Notes: The above figures show Medicaid reimbursement rates and coverage acceptance by retail clinics across the United States. Subfigure (a) shows state-level Medicaid reimbursement rates for an office visit of low complexity (CPT 99201) in 2010. Subfigure (b) shows the share of states in which CVS MinuteClinics 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 (2024), and data on Medicaid acceptance by CVS MinuteClinics was collected by the authors as outlined in footnote 31.

Figure A15: Effects by primary care providers per capita at baseline

(a) Unconditional on baseline reimbursement rates



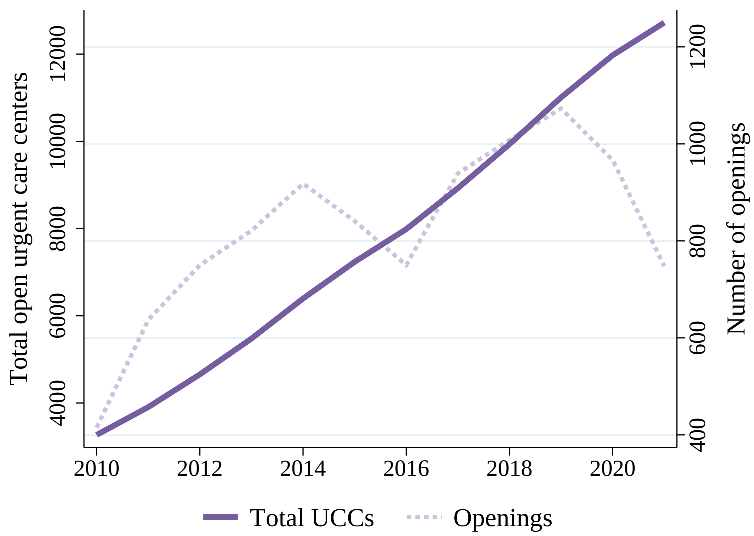

(b) Conditional on baseline reimbursement rates



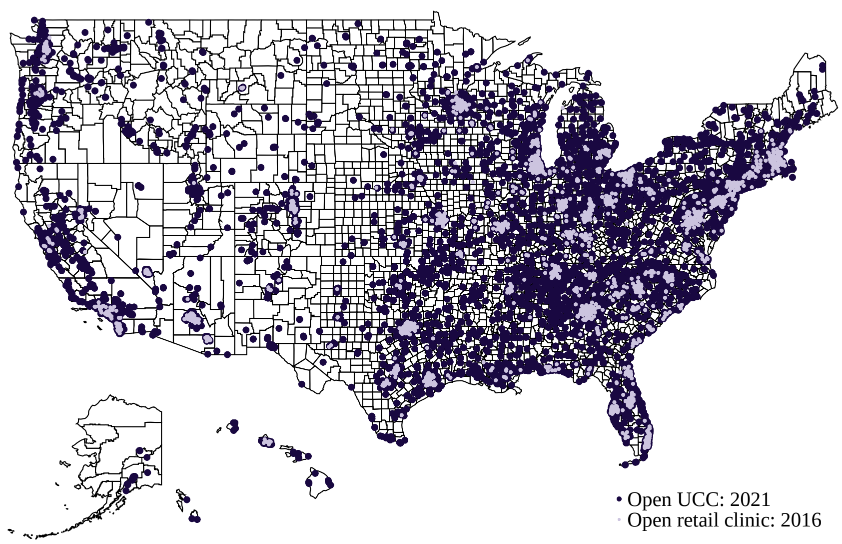

Notes: The above figures show how county-level changes in retail clinics per 100,000 people from 2013 to 2015 covary with county-level changes in the share of the population with private insurance coverage (left subfigures) and Medicaid coverage (right subfigures) over the same period. These relationships 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 (2024).

Figure A16: Urgent care centers across the United States

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



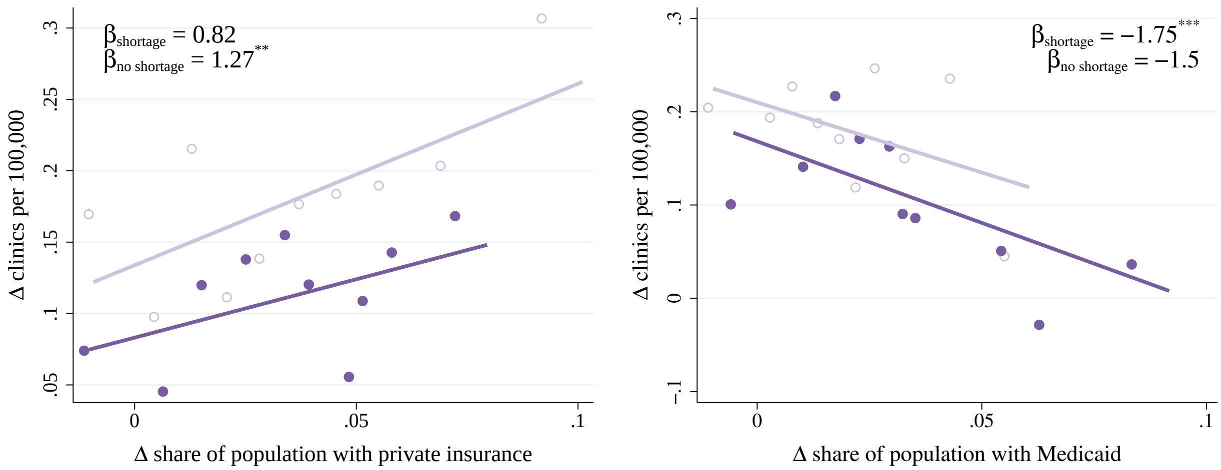

(b) Locations: 2021



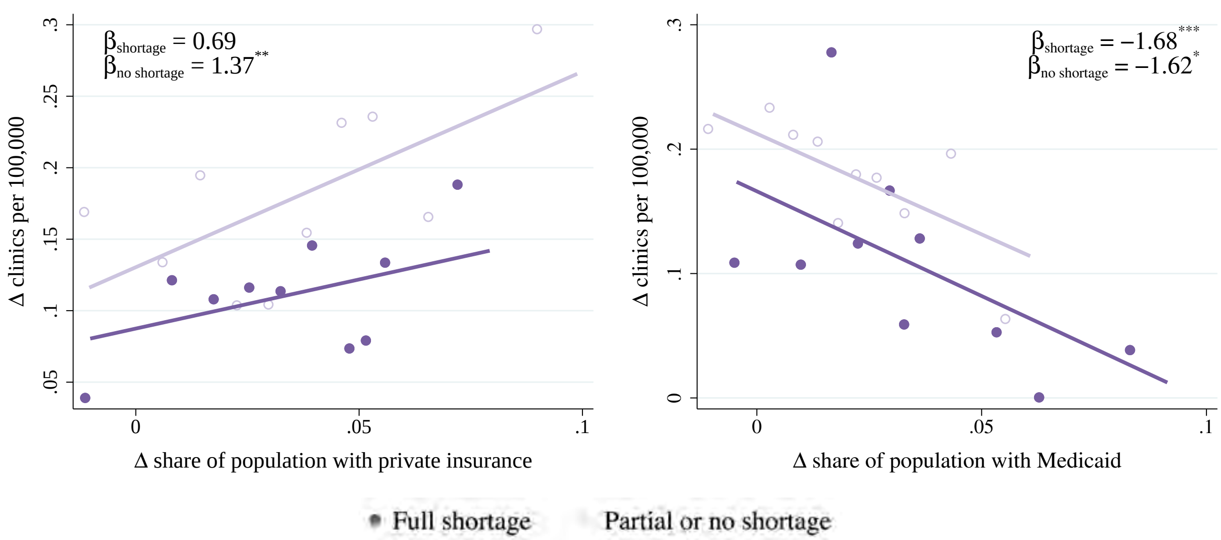

Notes: The above figures show the number and locations of urgent care centers across the United States. Subfigure (a) shows the total number of open urgent care centers (dark, solid line) and annual entry conditional on survival to 2021 (light, dashed line) from 2010 to 2021. Subfigure (b) shows the 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 A17: Effects by baseline primary care shortage designations

(a) Unconditional on baseline reimbursement rates




(b) Conditional on baseline reimbursement 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 (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 (2024).

# B Supplementary tables

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

| | In one-year ACS | | Not in one-year ACS | |
|---|---|---|---|---|
| Retail clinics in 2016: | One+ (1) | None (2) | One+ (3) | None (4) |
| **a. Retail clinics** | | | | |
| *2016* | | | | |
| Open clinics | 5.51 | 0 | 1.60 | 0 |
| Clinics per 100,000 | 1.13 | 0 | 2.39 | 0 |
| *2010–2016* | | | | |
| 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. County characteristics (2012–2016)** | | | | |
| *Basic demographics* | | | | |
| 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 |
| *Income and education* | | | | |
| 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 |
| *Health insurance* | | | | |
| 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: Changes in insurance and retail clinic penetration: OLS by expansion status

| | All counties | | Medicaid non-expanders | | Medicaid expanders | |
|---|---|---|---|---|---|---|
| Dependent variable: retail clinics per 100,000 | (1) | (2) | (3) | (4) | (5) | (6) |
| Share insurance | 0.242 | | 1.436** | | -1.525*** | |
| | (0.412) | | (0.557) | | (0.406) | |
| Share private | | 0.979** | | 1.736*** | | -1.370** |
| | | (0.453) | | (0.583) | | (0.577) |
| Share Medicaid | | -1.279*** | | 0.870 | | -1.465*** |
| | | (0.448) | | (0.785) | | (0.467) |
| 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 |
| $R^2$ | 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 A3: First-stage results from first-difference specification

| Dependent variable: | Share insured (1) | Share private (2) | Share Medicaid (3) |
|---|---|---|---|
| $Post_t \times Employed_c^{2013}$ | 0.066*** | 0.071*** | -0.002 |
| | (0.016) | (0.016) | (0.012) |
| $Post_t \times [138 - 400\%\ FPL]_c^{2013}$ | 0.277*** | 0.243*** | 0.021 |
| | (0.053) | (0.050) | (0.038) |
| $Post_t \times Expansion_s \times [< 138\%\ FPL]_c^{2013}$ | 0.148** | -0.019 | 0.170*** |
| | (0.065) | (0.052) | (0.037) |
| Year fixed effects | X | X | X |
| Demographic controls | X | X | X |
| Observations | 3,313 | 3,313 | 3,313 |
| $R^2$ | 0.516 | 0.422 | 0.295 |
| Mean dependent variable | 0.02 | 0.009 | 0.007 |

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 A4: County-level summary statistics by urgent care center presence

| | Number of UCCs in 2021 | | |
|---|---|---|---|
| | One or more (1) | None (2) | P-value (3) |
| **a. On-demand health care clinics** | | | |
| *Urgent care centers (2021)* | | | |
| Open clinics | 6.81 | 0 | |
| Clinics per 100,000 | 5.54 | 0 | |
| *Retail clinics (2016)* | | | |
| Share any clinic | 0.25 | 0.01 | |
| Open clinics | 1.08 | 0.02 | |
| Clinics per 100,000 | 0.36 | 0.05 | |
| **b. County characteristics (2016–2020)** | | | |
| *Basic demographics* | | | |
| 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 |
| *Income and education* | | | |
| 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 |
| *Health insurance* | | | |
| 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 A5: Effects of insurance on retail clinic penetration: ESI versus direct purchase

| | OLS | 2SLS | | | |
|---|---|---|---|---|---|
| Dependent variable: | Retail clinics per 100,000 (1) | Share ESI (2) | Share direct (3) | Share Medicaid (4) | Retail clinics per 100,000 (5) |
| **a. First stage** | | | | | |
| $Post_t \times Employed_c^{2013}$ | | 0.173*** | 0.025** | -0.056*** | |
| | | (0.019) | (0.011) | (0.017) | |
| $Post_t \times [138 - 400\%\ FPL]_c^{2013}$ | | 0.127** | 0.242*** | 0.089 | |
| | | (0.059) | (0.061) | (0.071) | |
| $Post_t \times Expansion_s \times [< 138\%\ FPL]_c^{2013}$ | | 0.106* | -0.142*** | 0.321*** | |
| | | (0.064) | (0.050) | (0.050) | |
| **b. OLS/ 2SLS** | | | | | |
| Share ESI | 0.904* | | | | 7.517*** |
| | (0.486) | | | | (2.275) |
| Share direct purchase | 1.235* | | | | -7.514* |
| | (0.697) | | | | (4.501) |
| Share Medicaid | -1.267*** | | | | -7.045*** |
| | (0.444) | | | | (2.479) |
| County fixed effects | X | X | X | X | X |
| Year fixed effects | X | X | X | X | X |
| Demographic controls | X | X | X | X | X |
| Observations | 3,869 | 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 | 0.6120 |
| 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 analogs 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 analogs 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

| | "Private" instrument 1 | | | "Private" instrument 2 | | |
|---|---|---|---|---|---|---|
| Dependent variable: | Share private (1) | Share Medicaid (2) | Retail clinics per 100,000 (3) | Share private (4) | Share Medicaid (5) | Retail clinics per 100,000 (6) |
| **a. First stage** | | | | | | |
| $Post_t \times Employed_c^{2013}$ | | | | 0.211*** | -0.053*** | |
| | | | | (0.024) | (0.017) | |
| $Post_t \times [138 - 400\%\ FPL]_c^{2013}$ | 0.431*** | 0.072 | | | | |
| | (0.064) | (0.072) | | | | |
| $Post_t \times Expansion_s \times [< 138\%\ FPL]_c^{2013}$ | -0.123 | 0.346*** | | 0.021 | 0.335*** | |
| | (0.080) | (0.052) | | (0.085) | (0.049) | |
| **b. Two-stage least squares** | | | | | | |
| Share private | | | -1.726 | | | 6.086*** |
| | | | (1.997) | | | (1.791) |
| Share Medicaid | | | -3.676 | | | -2.215 |
| | | | (2.241) | | | (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 II, 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 1 and 2 for the one-price case.[43] We therefore instead present

[43]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:

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

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 $\left(\frac{\tilde{q}}{\tilde{q}+N_M}\right)$ and Medicaid patients $\left(\frac{N_M}{\tilde{q}+N_M}\right)$, 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.[44] Mirroring the predictions in Section II and the findings in Figure 7, we show

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

[44]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 $\Pi'(0) = -p'(\tilde{q}^*)\frac{\tilde{q}^*}{\tilde{q}^* + N_M} q^*$.[45] 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

[45] 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.[46]

Now consider the effects of a Medicaid expansion. As shown in Figure 1(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.[47] Mirroring the predictions from the one-price case, we therefore have that Medicaid

[46]When the firm accepts Medicaid, this follows from the firm's first-order condition for $\tilde{q}$ $\left(\text{given by } p'(\tilde{q}^*)\tilde{q}^* + (p(\tilde{q}^*) - p_M)\frac{N_M}{\tilde{q}^*+N_M} = 0\right)$. 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.

[47]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 13 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 IV.C, 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 analogs of the two-way fixed effects specifications introduced in Section III.C 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:

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

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 analogs of the instrumental variables specifications introduced in Section IV.A 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:

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

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

$$
\begin{aligned}
\{Entries_{ct},\, Exits_{ct}\} = \; & \left\{ \beta \cdot \Delta \widehat{Insured}_{ct},\; \beta_1 \cdot \Delta \widehat{Private}_{ct} + \beta_2 \cdot \Delta \widehat{Medicaid}_{ct} \right\} \\
& + \alpha_4' \cdot \Delta \left\{ Post_t \cdot [< 138\%\ FPL]_c^{2013} \right\} + \alpha_5' \cdot \Delta \left\{ Post_t \cdot Expansion_s \right\} \\
& + \delta' \cdot \Delta X_{ct} + \eta' \cdot Clinics_{ct-1} + \gamma_t' + \epsilon_{ct}', \qquad \text{(A3)}
\end{aligned}
$$

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).