Competition and Selection in Health Insurance Markets

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Abstract

Competition in US health insurance markets is low and has declined in recent years. Insufficient competition is often assumed to increase plan premiums or decrease benefit quality, but the latter has been difficult to establish empirically. Moreover, why health insurance competition is so low is poorly understood. As recent health insurance expansions rely on private insurers to provide coverage, understanding why health insurance competition is low and how this affects consumers is important for policy.

Paper 1 tests for a relationship between insurer competition and health plan benefit generosity. I examine the impact of a regulatory change that led to the cancellation of 40% of the private plans participating in the Medicare program. I isolate the causal effect of cancellation using variation induced by insurers who removed all plans nationally. Insurers in markets affected by cancellation responded by reducing benefit generosity. In the average market, out-of-pocket costs for a representative beneficiary increased by about $130 per year. Tests of possible mechanisms suggest that insurers primarily responded to changes in competition, rather than the policy's direct costs or anticipated changes in enrollees' health risks. In the least competitive markets, out-of-pocket costs increased by more than $200 a year, while in markets with the most substitutes for cancelled plans, benefit generosity barely changed. These findings have crucial implications for markets such as health insurance exchanges, as they suggest health plan quality is degraded when competition is insufficient.

Paper 2 explores why health insurance markets are so concentrated. This paper tests how insurer and provider market power affects insurer exit using a policy change in Medicare Advantage. Under the policy, a group of indemnity insurers were forced to form provider networks *de novo.* Insurers cancelled two-thirds of the affected plans following passage of this mandate. Comparison across markets where insurers selectively withdrew plans suggests that greater provider market power led to increased exit while greater insurer market power protected against it. Insurers in markets at the top decile of physician and hospital concentration were respectively 17 and 15% more likely to exit than those in the bottom decile, while insurers in the top decile of insurer market share were 68% less likely to exit than those in the bottom decile. Additionally, insurer bargaining power is found to be most protective in the most concentrated hospital markets. Findings suggest that policies to foster insurer market participation must consider both insurer and provider market structure.

Paper 3 examines trends in Medicare Advantage enrollment. Medicare Advantage enrollment grew to its highest point in program history in 2014, despite five years of payment cuts and declining plan availability. This paper investigates whether recent enrollment growth can be expected to continue by examining trends in 65-year-olds' Medicare Advantage enrollment. As 65-year-olds are choosing among supplemental Medicare options for the first time, they may be more responsive to market conditions than other beneficiaries. Findings show that 65-year-olds' enrollment patterns differ from older cohorts, in that they increased between 2006-2009 and then leveled off between 2009-2011. Among a range of market and plan characteristics, changes in Medicare Advantage plan premiums and benefit generosity most plausibly explain slowing enrollment growth. The data also suggest that, absent the recession, enrollment might have further declined.

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Chapter 1

Paying more for less? Insurer competition and health plan generosity in the Medicare Advantage program.

Many health insurance markets, such as those formed under the Affordable Care Act (ACA) and Medicare Part D, are structured to foster insurer competition. Regulators set parameters on the number and type of plans insurers can offer, and insurers offer plans based on these parameters. Enrollees choose from a menu of plans, which, in theory, encourages insurers to offer efficient levels of premiums and benefits. This idea of "managed competition" drove market design in recent health insurance expansions; it is also the model used by many large employers for years.(1)

Policymakers regulate insurance markets to achieve specific policy goals. Some constraints – such as benefit mandates – specify precisely what services insurers must cover and are intended to insure access to specific services or for specific populations. Other regulations encourage minimum or maximum levels of benefit generosity. For instance, the ACA includes both a "Cadillac tax" on employer-sponsored plans, which is intended to discourage excessively generous policies(2), and mandatory out-of-pocket spending limits for exchange plans, to guarantee enrollees' financial protection. These restrictions advance different objectives, but all limit insurers' flexibility in offering plans.

Regulations are generally motivated by sound policy, but may also have unintended effects. In equilibrium, insurers may respond to regulations affecting only a subset of plans by adjusting the premiums or benefits of all plans. The expected direction of these equilibrium effects is not clear *ex ante*. For instance, when policy changes cause insurers to remove plans from the market, remaining plans gain market power. Added market power may give insurers leeway to reduce benefit generosity or increase premiums. Alternately, when plans are removed from a market, a subset of individuals will need to search for new plans. Enrollees making active choices are often more price- or benefit-elastic(3, 4), so insurers may lower premiums or increase benefit generosity to attract new members.

In the policy experiment analyzed here, Congress passed legislation intended to reduce overpayment of a particular type of private health insurance plan in Medicare – private-fee-for-service (PFFS) plans. PFFS plans are offered through the Medicare Advantage (MA) program, in which Medicare beneficiaries purchase health plans that provide equivalent coverage to traditional Medicare. Prior to the policy change, PFFS plans were not required to create provider networks. Rather, beneficiaries could visit any Medicare provider at no extra cost. While other MA plan types had to negotiate payment rates with healthcare providers, PFFS plans could pay providers lower, administratively set Medicare rates. As all MA plan types were compensated for accepting beneficiaries at the same rates, there was potential for PFFS plans to have higher margins.

Congress responded to reports suggesting PFFS plans were overpaid(5–7) by passing a law requiring insurers to establish provider networks for PFFS plans that were distinct from traditional Medicare. The law also stipulated that insurers must negotiate payment rates for providers in those networks, rather than paying administratively set Medicare rates.(8) These requirements removed PFFS plans' cost advantage and eliminated the characteristic that differentiated them. Insurers responded by canceling roughly two-thirds of their PFFS plans, forming networks for the remainder.

I explore how remaining plans' characteristics changed in response to PFFS cancellation using cross-county variation in cancelled plans' market shares. To address the fact that PFFS plans were not randomly distributed at baseline, I use a difference-in-differences specification with county and year fixed effects, where cancelled plans' market shares are used as a continuous "treatment" variable. To avoid endogeneity due to selective cancellation, I estimate cancellation's impact using baseline PFFS market shares for insurers who cancelled all PFFS plans nationally. National cancellations are plausibly unrelated to unobserved confounding variables, because an insurers' decision to cancel all plans is unlikely to be driven by unobserved changes in profitability in local markets.

I relate nationally cancelled plans' market shares to two measures of plan generosity: expected out-of-pocket costs for a representative enrollee, and plan premiums. I find clear evidence that out-of-pocket costs increased in markets with more cancellation. In the average county, beneficiaries paid about $10.80 more a month out-of-pocket due to plan cancellation; premiums were largely unaffected. This effect was not limited to plans directly affected by the policy. Preferred Provider Organization (PPO) plans, which were likely the closest substitutes to PFFS plans, also increased out-of-pocket costs by about $7.40 a month in the average county. These estimates suggests that cancellation reduced MA plans' generosity advantage over traditional Medicare by about $15 - 20\%$.[1]

Several possible mechanisms might account for cancellation's effect on benefit generosity. As noted above, insurers may have responded to increased market power. Alternately, cancelled plans' enrollees were in worse health on average than other plans' enrollees,

[1]On average, MA plans have more generous coverage than Traditional Medicare. Average generosity of MA relative to traditional Medicare is calculated by comparing out-of-pocket costs in traditional Medicare to out-of-pocket costs in all MA plans. During the study period, MA plans save beneficiaries an average of about $50 a month, relative to traditional Medicare.

so many plans staying in the market stood to enroll sicker (and potentially costlier) beneficiaries. Differences in enrollee health might have caused plans to adjust benefits in one of two ways: 1) after enrolling sicker beneficiaries or 2) in advance, to deter sicker beneficiaries from enrolling. I test each of these hypotheses using data on plans' risk scores and competition between PFFS plans' substitutes.

Results suggest that changes in enrollee health did not drive changes in benefits. Plans remaining in the market did not systematically enroll sicker beneficiaries following cancellation. Evidence also fails to support the hypothesis that changes in benefits were caused by insurers trying to drive away sicker enrollees. Specifically, benefit reductions are not explained by controlling for plans' anticipated risk – the risk a plan could expect if they enrolled cancelled plans' beneficiaries. In contrast, there is strong evidence that decreased competition drove benefit reductions. Benefits decreased most in markets with the least competition between PFFS plans' substitutes – the markets where insurers likely gained the most market power. In these markets, out-of-pocket cost for a representative beneficiary rose $\$15-20$ a month for plan types not directly affected by the policy.

This analysis focuses on a particular policy change, but the results are relevant to a broader literature on health insurance competition and plan characteristics. First, many studies have investigated the impact of plan competition on premiums. These studies generally find that premiums are higher in less competitive markets(9–15) and that consolidation increases premiums.(16, 17) However, there are exceptions; for instance, Feldman, et. al. (1996) find that HMO mergers only increase premiums in the most competitive markets and that mergers' effects dissipate quickly.(18)[2]

One challenge in assessing competition's effect on plan characteristics is that exogenous variation in competition is rare and omitted variables may cause both high premiums (or stingy benefits) and concentrated insurance markets. Researchers ad-

[2]They note two potential explanations for their finding. First, their data are from an era when the HMO market was more competitive (1985-1993). Second, most mergers they observe are between small market players, rather than dominant firms.

dress this issue in a variety of ways. One approach is to flexibly control for a broad range of variables that might affect both competition and plan characteristics.(15, 18) Another is to specify a structural model of firm competition and simulate competition's effects.(11, 13, 14, 17) Lastly, a few studies have identified novel sources of variation in insurance market competition.(9, 10, 12, 16) For instance, Dafny, et. al. (2012) examine the effect of competition on premiums, using local variation in insurance market structure caused by the merger of two national firms. This approach avoids many potential sources of endogeneity, as the two merging firms operated in all markets prior to the merger. Hence, insurers' decision to merge is unlikely to be related to omitted variables affecting premiums at the local level.

Though many papers have studied competition's effect on premiums, few have examined the effects of competition on other plan benefits.(14, 19–21) Moreover, their conclusions conflict. Town and Liu (2003) and Pizer and Frakt (2002) find that reducing competition in Medicare Advantage decreases benefits (drug coverage and cost sharing), while Chorniy, et. al. (2013) find that consolidation in Part D leads to better benefits (more generous formularies).(20) Finally, Duggan, Starc, and Vabson (2014) find that greater competition in Medicare Advantage has no effect on beneficiary out-of-pocket costs.

Lastly, several studies find that the relationship between competition and plan characteristics is complicated by adverse selection.(11, 13, 15, 21) For instance, using a structural model, Starc (2010) finds that higher market concentration increases premiums for supplemental Medicare plans (Medigap). She also finds that selection limits the degree to which firms can increase premiums, because insurers who charge high premiums risk drawing the sickest enrollees. Similarly, Lustig (2011) finds that the welfare losses from selection are greatest in the most competitive markets, because selection pressures are exacerbated when there are more firms.

This paper contributes to the literature in three ways. First, as noted above, isolating the causal effect of competition on plan characteristics is a challenge. This study shows, in a well-identified setting, that competition can have a major impact on insurance plan generosity. Second, few studies explore the relationship between competition and plan benefits (besides premiums), and these studies present a confusing picture. This study adds support to papers suggesting that decreased competition reduces benefits. Lastly, this study adds to a limited literature testing whether enrollee health mediates the relationship between competition and plan characteristics. I do not find that selection drove benefit reductions, but do find limited evidence that insurers may distort benefits to attract healthier enrollees.

This research has several policy implications. First, regulators and researchers often focus on premiums when assessing the effects of changing competition.(22) This research suggests this focus is insufficient, as insurers may modify benefits while leaving premiums unchanged. Second, results show that regulators should be cautious when limiting the plans insurers may offer – legislators did not intend to reduce competition when changing PFFS requirements, but this was one of the law's chief effects. Finally, benefits in the Medicare Advantage market are tightly regulated, relative to other insurance markets. This study shows that, even when policymakers regulate how benefits are set, plan generosity suffers when competition falls.

This paper is structured as follows. Section 1.1 is a conceptual discussion of the impact of insurance competition on benefit generosity. Section 1.2 summarizes the policy change. Section 1.3 describes the data and methods used to estimate cancellation's causal effect on remaining plans' characteristics. Section 1.4 describes the policy's effects on markets, and Section 1.5 explores cancellation's effect on benefit generosity. Section 1.6 shows that changes in generosity are greatest in the least competitive markets, while Section 1.7 presents evidence that changes in generosity are not driven by actual or anticipated changes in enrollee health. The last section concludes.

1.1 Benefit generosity and competition

Two findings from the theoretical literature on competition and quality are relevant to this study. The first is that changes in competition may increase or decrease benefit generosity. The second is that, under certain market conditions, insurers may prefer to adjust benefits rather than premiums.

To see that competition may increase or decrease health insurance benefit generosity, consider a simple model of competition and quality, similar to those used by Gaynor (2006) and Dorfman and Steiner (1953).(23, 24) In this model, health insurance benefit generosity, determined by plan financial characteristics such as copays, deductibles, and covered benefits, can be thought of as plan quality.[3] Profit maximizing firms set both price and quality, based on the following profit function:

$$\pi = s(p, q)(p - c(q))$$

where premiums are p, quality is q, consumer demand is $s(p, q)$, and cost is $c(q)$. Consumers always prefer lower prices and higher quality, so $s(p, q)$ is decreasing in p and increasing in q. Improving quality is costly, so firms' costs, $c(q)$, increase with q.

Gaynor (2006) shows that equilibrium price and quality are related based on the following condition:

$$q = \frac{\varepsilon_q}{\varepsilon_p} * \frac{p}{c'(q)} \tag{1.1}$$

where $\varepsilon_p = \frac{\partial s(p,q)}{\partial p} \frac{p}{s}$ is the price elasticity of demand, $\varepsilon_q = \frac{\partial s(p,q)}{\partial q} \frac{q}{s}$ is the quality elasticity of demand, and $c'(q)$ is the marginal cost of additional quality. Changes in competition change the slope of the demand curve faced by the individual firm, thereby changing ε_q and/or ε_p; for example, when more firms enter the market, consumers have a wider array

[3]Plan financial characteristics are only one measures of plan quality. Quality may also be defined by generosity of networks, customer service, etc.

of options and are therefore more price- or benefit-sensitive (ε_q and/or ε_p increase).[4]

It is a standard result that increases in market power increase mark-ups ($p - c(q)$). However, with this profit function, insurers may raise mark-ups either by increasing premiums or degrading quality (thereby reducing costs) or both. Equation (1.1) reveals why competition's effect on quality is ambiguous. When quality and price are both allowed to vary, increased market power could increase p or decrease q. As quality and price are determined by the ratio of ε_q to ε_p, any change that increases ε_p more than ε_q will result in lower quality for a given price.

Competition unambiguously increases quality in the special case where premiums are fixed – as in a market with regulated prices. With an exogenously fixed p, firm entry will only change ε_q, increasing optimal q. This reflects the general result that quality in markets with regulated prices may be "too high".(24)

The model defined above greatly simplifies the firm's problem, and it is reasonable to believe that many other factors mediate the relationship between competition and quality.[5] Given the ambiguity of theoretical predictions in a simplified model, cancellation's effect on generosity is an empirical question.

1.2 Background: policy change and Medicare Advantage

The Medicare program provides health insurance to the aged and disabled in the US. Beneficiaries can choose to receive coverage through traditional fee-for-service Medicare or enroll in a private insurance plan through the Medicare Advantage (MA) program. In 2014, almost one in three Medicare beneficiaries were enrolled in MA.(26)

[4] This presumes entry is exogenous.

[5] For instance, with adverse selection, costlier patients may be more willing to pay for health insurance and changing p will also affect c. When benefits are fixed, adverse selection is often thought to "restrain" mark-ups by limiting insurers' ability to raise premiums.(13) However, when benefits vary, competitive insurers may select patients by offering policies with low premiums and skimpy benefits. This suggests that quality might be higher in monopoly markets.(11, 25)

MA plans are paid by the federal government for accepting enrollees and are required to provide benefits actuarially equivalent to those covered in traditional Medicare.[6] However, conditional on this restriction, plans' benefit structure can differ substantially from traditional Medicare. For instance, MA plans often charge flat copays for office visits, while traditional Medicare charges a 20% coinsurance. MA plans also often set beneficiary out-of-pocket spending limits, even though there is no spending limit in traditional Medicare.[7]

The amount the government pays a plan for accepting enrollees is set based on county-level payment "benchmarks".[8] If government benchmark payments exceed insurers' expected costs of covering beneficiaries in a county, insurers are required to use excess payments to provide additional benefits. Plans can reduce cost sharing for standard benefits or provide extra benefits such as drug or dental care. Though MA plans' generosity varies, MA benefits are generally more generous than traditional Medicare.(28)

If insurers estimate that the costs of providing benefits exceed county-level benchmarks, they may charge enrollees an additional premium beyond the standard Part B premium paid by all Medicare enrollees. Plans cannot price discriminate; all enrollees must be charged the same premium and offered the same benefits. To reduce insurers' incentives to select healthier enrollees, the government risk-adjusts payments based on each enrollee's health.

Though insurers can charge enrollees an additional premium, most do not. Insurers

[6]Traditional Medicare benefits include hospital (Part A) and outpatient medical (Part B) services, but not drug (Part D) coverage.

[7]CMS encouraged plans to have out-of-pocket limits starting in 2010,(27) and required them to have out-of-pocket limits starting in 2012. Though there is no spending limit in traditional Medicare, beneficiaries often buy Medigap plans for additional financial protection.

[8]Between 2006-2011, plans submitted 'bids' reflecting their estimated cost of covering beneficiaries in a county. If plans bid above the county-level benchmark, then they were required to charge enrollees a premium. If a plans' bid was below county-level benchmarks, then 75% of the difference between the bid and the benchmark was rebated to the plan. Rebates must be used to provide additional benefits. This system was modified in 2012, so that the percent rebated to a plan is adjusted by its quality rating.

may also charge a "negative premium" by reducing the Part B premium, but few plans do this either. As a result, over half of all MA plans in 2007-2009 charged exactly \$0. This unusual pricing structure is likely driven by two institutional features of Medicare. The first is that enrollees must write checks for MA premiums, while Part B premiums are deducted directly from enrollees' social security. The second is that, while plan benefits and premiums are clearly summarized in Medicare promotional materials, Part B premium reductions are not.[9] These regulatory quirks may make MA premiums more salient to beneficiaries than reductions in Part B premiums, resulting in plans "bunching" at the \$0 price point.(29)

The most commonly offered types of MA plans are health maintenance organizations (HMOs), preferred provider organizations (PPOs), and private fee-for-service (PFFS) plans. MA HMOs and PPOs are structured similarly to their private market counterparts; HMOs contract with a network of doctors, and out-of-network care is generally only covered in emergencies. PPOs have broader networks, higher premiums, and allow patients to seek out-of-network care at higher levels of cost-sharing.

PFFS plans are unique to MA and most closely resemble indemnity plans. Before Congress passed a law changing how PFFS plans contract with providers, PFFS plans did not have to form networks or negotiate payment rates with providers. Instead, enrollees could visit any doctor accepting Medicare, and plans could pay providers based on the Medicare fee-for-service schedule, which is generally believed to pay doctors less than commercial rates.(30, 31) The ability to pay Medicare rates may have helped PFFS plans operate profitably in counties where other plans struggled to negotiate with doctors. Not having to form networks also reduced PFFS plans' costs of entry, allowing insurers to differentially enter counties where benchmark payments were relatively high.(5) PFFS plans were appealing to many beneficiaries, because enrollees could both access the

[9] For instance, there is nowhere to display Part B premium buy-downs on Medicare's plan finder website.

network of traditional Medicare and enjoy the more generous benefits of an MA plan.

Many policy experts claimed that favorable requirements led to Medicare paying more for PFFS enrollees than it would if enrollees chose other plan types.(5, 6, 30, 32, 33) For instance, PFFS enrollment disproportionately came from counties with higher benchmark payments. This resulted in an average enrollment-weighted payment rate for PFFS plans that was 122% of traditional Medicare costs, compared to 116% for other plan types.(5)

In July 2008, Congress responded to these concerns by passing a law changing MA policy. The Medicare Improvements for Patients and Providers Act required PFFS plans to form provider networks, changing their cost structure and profitability. PFFS plans were allowed to charge higher cost sharing for beneficiaries seeking care out of network, but were held to the same standards in building a network as HMOs and PPOs.(8)

This policy substantially changed the menu of options available in MA. Although the network requirement was not effective until 2011, insurers' response to the policy was immediate. In 2009, Wellcare, Coventry, and Healthnet, who collectively enrolled 20% of all beneficiaries in PFFS plans, announced the termination of all PFFS contracts starting in 2010.(34) A further 25 of the 62 remaining PFFS contracts were cancelled between 2010 and 2011, and total enrollment in PFFS plans fell over 75%.(35) Remaining plans complied with the policy change and formed networks.

Figure 1.1 illustrates the scope and geographic variation in PFFS plans' loss of market share, by showing the share of MA enrollees in PFFS plans each year between 2009-2012. The policy change affected most counties in the US, but its impact varied across counties based on PFFS plans' pre-policy market shares. As more PFFS plans were offered in counties where benchmarks were high relative to costs, some counties had greater exposure to the policy. The impact of the policy also differed across insurers, as firms that offered more networked plans (HMOs/PPOs) could more easily comply with the network requirement.

Plan cancellation had two immediate effects on plans that stayed in the market. First,

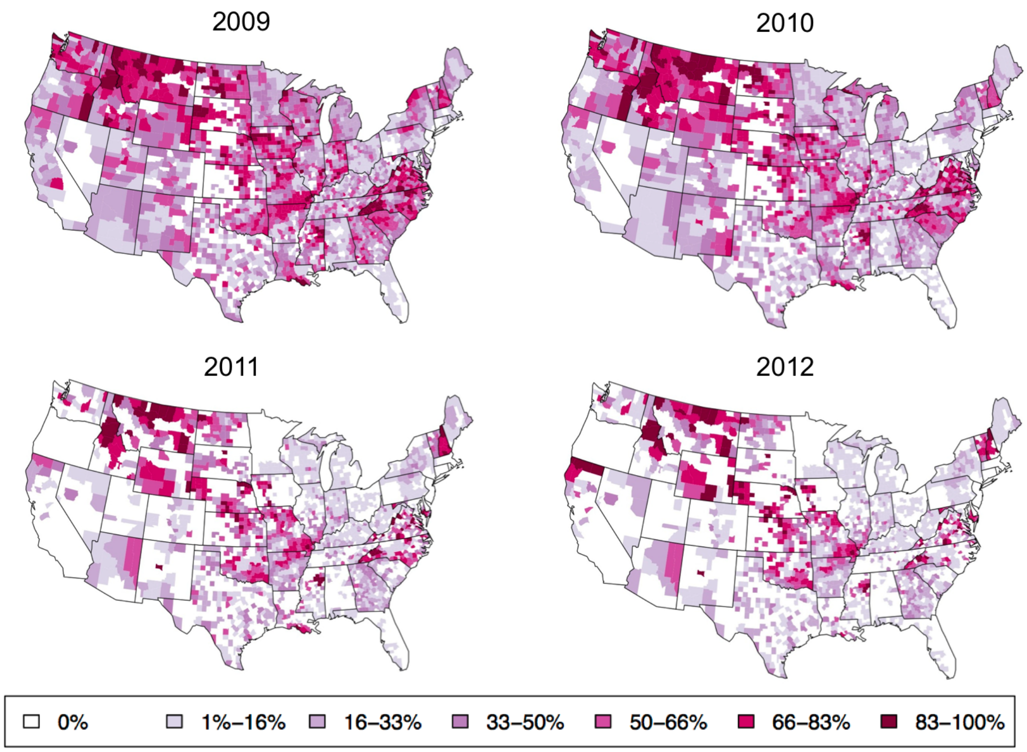

Figure shows the percent of the MA market enrolled in PFFS on the county level, immediately following the policy change, 2009 to 2012. Darker colors represent greater enrollment.

Figure 1.1: Private fee-for-service market shares (2009-2012)

cancellation changed the competitive environment. Many insurers offered PFFS plans in counties where they did not offer an HMO/PPO, so cancellation caused them to exit these markets entirely. Second, cancellation redistributed enrollees within Medicare. Insurers could not automatically reenroll beneficiaries into HMOs or PPOs they offered in the area,(36) so cancelled plans' enrollees actively chose a new plan or defaulted into traditional Medicare. The resorting of enrollees across plans changed the distribution of market shares, and may have changed the distribution of patient health risks between plans and between MA and traditional Medicare.

1.3 Methods and data

Cancellation's effects on premiums and benefits is estimated using a difference-in-differences specification with a continuous treatment variable. The treatment variable captures cancellation's impact on a market and is measured using cancelled plans' share of MA enrollment. Counties with greater enrollment in cancelled plans are more heavily "treated."

Difference-in-differences specifications control for baseline differences across markets. However, regressing plan characteristics on cancelled plans' shares will yield biased estimates if insurers selectively exited counties that were becoming less profitable. To avoid endogeneity due to selective cancellation, shares of PFFS plans cancelled in all markets are used to estimate causal effects. As the insurers offering these plans chose to cancel all PFFS plans nationally, the decision to cancel will be unrelated to changes in any single market.

Estimates may also be biased if insurers withdrew from less profitable markets first or responded to competitors' cancellations by temporarily expanding.[10] To reduce bias from endogenous timing of cancellation, shares of nationally-cancelled insurers are fixed at 2009 levels, prior to any insurer withdrawing from the PFFS market.[11]

The baseline difference-in-differences specification is:

$$Y_{jmt} = \beta_0 + \beta_1 Post_{mt} * S_{m(2009)} + \eta M_{mt} + \theta X_j + \gamma_m + \tau_t + \varepsilon_{jmt} \tag{1.2}$$

where Y_{jmt} measures premiums or plan generosity for plan j in market m at time t. $S_{m(2009)}$ is the local market share of nationally cancelled plans, fixed at 2009 levels, and

[10] Inspection of the data suggests both occurred. Wellpoint and Aetna reduced the geographic spread of their plans before ultimately canceling all PFFS plans, while CIGNA briefly increased plan offerings.

[11] It is possible that insurers anticipated the law's passage and responded preemptively to the expected decline in PFFS plans' profitability. This would a threat to validity if these endogenous responses were reflected in insurers' 2009 shares. To test this, I also fix shares at 2008 levels and test for effects on benefits. Results are similar (Appendix Table A1, Column 1), but this specification is not preferred, as it limits tests of pre-period parallel trends.

$Post_{mt} = 1$ after 2009, the year before the first PFFS plans were cancelled. Although many counties were not directly affected by cancellation until 2011 or 2012, $Post_{mt} = 1$ is fixed at 2010 to avoid potential endogeneity from unprofitable markets experiencing cancellation first.[12]

Baseline differences across counties are controlled for using a full set of county fixed effects (γ_m). Year fixed effects (τ_t) absorb overall shifts in benefit generosity common to all plans. A range of time-varying, county-level characteristics (M_{mt}) control for changes in the economic and health market conditions that might affect plan benefits (described below). Lastly, a vector of indicators for plan type (X_j) allows average generosity to vary across types. Regressions are weighted by plan enrollment, averaged across all years a plan is offered, and the unit of observation is the county-plan-year. Data span 2007-2012, three years before and three years after the first plans were cancelled in response to the policy.

Enrollment data from the Center for Medicare and Medicaid Services (CMS) administrative records are used to calculate cancelled plans' share of MA enrollment in a market.[13] Counties are treated as markets, both because firms offer plans on a county-by-county basis and because beneficiaries can only choose plans within their county. Insurers set benefits and premiums on the plan level, and plans can span multiple counties. To account for correlation across counties in plan benefits, standard errors in regressions are clustered at the plan level, and descriptive statistics are collapsed to the plan level. Clustering on the plan level also adjusts for autocorrelation in benefits and premiums.

The sample for analysis includes the three main plan types in MA: HMOs, PPOs, and PFFS plans. Plans that serve a limited set of enrollees or are subject to a different set of

[12] Allowing $Post_{mt}$ to equal 1 in the year where the first plans are cancelled in each county m yields similar results (Column 2, Table A1).

[13] Enrollment data comes from contract/plan/state/county files.

requirements, such as employer-sponsored plans, plans for institutionalized beneficiaries, and regional PPOs, are excluded.[14] Plans enrolling fewer then 11 people in a county are censored in CMS data, so they are omitted.

In some markets, only PFFS plans operated and were all cancelled following the policy change. To compare across a consistent set of markets, the sample is restricted to a balanced panel of counties (i.e., all counties with at least one HMO, PPO, or PFFS plan in all years). Imposing balance excludes 9% of counties, but does not substantially change results. The final sample includes 168,911 county-plan-year observations from 2,592 counties.[15]

Cancellation is identified using CMS Crosswalk files, which list the disposition of plans at the end of a contract year. Insurers can withdraw from a market by terminating a plan or reducing its service area. A county-plan combination is treated as cancelled if the plan is terminated or has a service area reduction, with no enrollment in the county in the following year. Descriptive statistics use all cancelled plans' county-level MA market shares – both terminated plans and plans with service area reductions – to characterize cancellation's impact on markets. Causal regressions such as those in Equation (1.2) only use shares from the subset of plans terminated nationally.

MA plans are frequently created, consolidated, and terminated. To accurately identify all PFFS plans that will eventually be cancelled, plans are linked across time using CMS crosswalks. Linking plans across time results in multiple observations for consolidated plans. Repeated observations are down-weighted by dividing plan enrollment equally across them.

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[14]Employer and special needs plans do not compete directly with other plan types for enrollees and often have special constraints on how benefits are set. For instance, employer-sponsored plans, offered by private insurers for Medicare-eligible retirees, set benefits through negotiation with employers. Regional preferred provider organizations (RPPOs) are also excluded, as RPPOs are paid differently and subject to different requirements than other plan types.

[15]These 2,592 counties contain 90% of the US population. Main results are replicated in unbalanced panel (Appendix Table A1, Column 2).

Two measures of plan characteristics are used as outcomes: expected out-of-pocket costs for the representative beneficiary and MA plan premiums (defined below). The first, out-of-pocket costs, is a summary measure of benefit generosity published by CMS for beneficiaries' use when selecting plans. It reflects spending for a representative individual in each plan after the plan covers costs. It is calculated using healthcare consumption data for a representative cohort of traditional Medicare beneficiaries. Holding consumption fixed, beneficiaries' expected spending is determined given each plan's copays, deductibles, spending limits, and covered benefits. As higher numbers reflect lower generosity, out-of-pocket costs can be thought of as measuring the "inverse" of generosity.

The main advantages of the out-of-pocket cost measure are that it is standardized across plans and captures a broad range of benefits, including spending on lab tests and diagnostics, prescription drugs, and vision and hearing services. Additionally, as it is calculated using a representative cohort, variation in out-of-pocket costs are not driven by the health or preferences of beneficiaries who endogenously opted to enroll in a particular plan.

One drawback to the out-of-pocket cost measure is that it is noisy. It is calculated using a different cohort of beneficiaries each year, and thus, varies idiosyncratically based on variation in their consumption patterns. To control for this, all regressions have year fixed effects. Another limitation is that out-of-pocket costs are calculated assuming all care is received in network. Insurers may have responded to PFFS cancellation by increasing out-of-network cost-sharing, as expansive networks were PFFS plans' main competitive advantage. Unfortunately, this measure will not capture changes in networks, potentially resulting in underestimates of insurers' response.

The second plan characteristic examined is the MA plan premium. As discussed in Section 1.2, MA insurers charge an additional premium above the standard Medicare Part B premium if their estimated costs exceed county benchmark payments. They may also reduce the Part B premium if benchmarks exceed estimated costs. I calculate one premium variable to capture both the MA premium and any reduction in the standard Medicare Part B premium. Premiums and benefits are considered separately, because insurers may adjust either characteristic in response to changes in market structure.

Market-level competition is characterized using Herfindahl-Hirschman indices (HHI), or the sum of insurers' squared market shares.[16] For descriptive statistics, HHI is calculated using insurers' shares of all three plan types – HMOs, PPOs, and PFFS plans. When testing for heterogenous responses to cancellation based on competition among PFFS plans' substitutes – HMOs and PPOs – PFFS plans are excluded from the HHI calculation. This is intended to produce a measure of "counterfactual HHI" and proxy for the market power that firms would have absent PFFS plans. (More details provided in Section 1.6.)

CMS Hierarchical Condition Category (HCC) risk scores are used to characterize enrollee health in cancelled plans and to test whether insurers adjust benefits in response to changes in enrollee health risk.[17] Risk scores capture average differences in spending across individuals based on demographic and diagnostic information.(37) They are normalized to 1, and higher numbers indicate worse risk (higher expected spending).

Additional county-level variables are used to control for factors that might affect benefit generosity (M_{mt} in equation (1.2)). These variables capture changes in the health care market (total number of hospital beds per 1000 residents from the American Hospital Association Annual Survey), healthcare utilization (risk-standardized per capita fee-for-service Medicare costs from CMS), generosity of plan payments (Medicare benchmark rates from CMS), economic factors (unemployment rate from the Bureau of Labor Statistics, percent of residents below poverty and per capita income ($1000's) from the Census Small Area Income Poverty Estimates), and market size (logged number of

[16]Data on plan ownership is available from contract/plan/state/county files.

[17]County- and plan-level risk scores are available in CMS plan payment files.

residents older than 65). The shares of beneficiaries enrolled in employer plans or special needs plans are used to control for shifts in closely related markets.[18]

Modifications to the baseline specification are used to distinguish competing mechanisms. Though baseline results show clearly that insurers decreased benefits, a range of factors might have driven them to do so. First, changes in benefits might be compositional, if cancelled plans were more generous than plans remaining in the market. Second, insurers might pass on the policy's costs to consumers by reducing benefits, as the policy both forced PFFS plans to build networks and pay providers higher (negotiated) prices. Third, as discussed above, insurers might reduce benefits due to decreased competition. Lastly, descriptive statistics show that cancelled plans had sicker enrollees. This fact suggests two additional hypotheses: 1) that cancelled plans passed on higher costs after enrolling sicker beneficiaries or 2) that cancelled plans modified benefits in advance to avoid attracting sicker beneficiaries.

Each of these mechanisms are tested separately. Whether changes in benefits are mechanically driven by generous plans' cancellation is tested by excluding any plan ever cancelled from the sample. (This restriction is maintained for all later tests.) The cost pass-through hypothesis is tested first by allowing Equation (1.2) to vary by plan type and then by excluding plans introduced in the post-cancellation period. The competition hypothesis is tested directly using the measure of counterfactual market competition described above. Indicators capturing variation in this measure are interacted with $S_{m(2009)}$ and $Post$ to test whether cancellation's effects differ based on the amount of market power a plan stood to gain. Lastly, hypotheses regarding changes in enrollee health are tested using data on plans' risk scores. I first test whether plans passed on higher costs of sicker enrollees by testing whether their enrollees' average health changes. Then, I test whether plans modified benefits in advance by constructing a

[18]Plan benchmarks come from published CMS ratebooks. FFS costs from from plan payment worksheets. Shares of beneficiaries in employer-only and special needs plans are calculated from contract/plan/state/county files.

measure capturing the amount of risk they stood to gain if they enrolled cancelled plans' beneficiaries. This variable is added directly to Equation (1.2) as a control. Further details of each test and mechanism are detailed in sections that follow.

1.4 Markets' and cancelled plans' characteristics

Table 1.1 summarizes baseline (2009) plan characteristics.[19] Cancelled PFFS plans were newer, smaller, and had higher risk scores than PFFS plans that stayed in MA, consistent with the idea that firms cancelled marginal plans. Cancelled plans were also less likely to have an HMO/PPO offered by the same insurer in their county, indicating that insurers with established networks could adapt more easily to the policy change.

At baseline, cancelled plans had lower out-of-pocket costs (greater generosity) than other PFFS plans, but higher out-of-pocket costs than HMOs and PPOs. Cancelled PFFS plans also had higher risk scores (worse average enrollee health) than PPOs and other PFFS plans, but lower risk scores than HMOs.[20] Differences in risk scores suggest that PPOs and PFFS plans' risk might have increased if they absorbed cancelled plans' enrollees. This motivates analysis testing whether canceled plans' risks affect benefits (Section 1.7).

Cancellation was widespread and had a significant impact on markets. Table 1.2 summarizes cancelled plans' market shares, the total number of cancelled plans, and the number of counties affected each year. Between 2009-2012, most counties were affected

[19] 2009 is treated as "baseline" because of the timing of plans' annual contracting and insurers' decision to cancel. MA plans contract with CMS on an annual calendar. When the Medicare Improvements for Patients and Providers Act passed in July 2008, plan contracts, which define benefits and market participation, had already been set for the 2009 calendar year. The first wave of cancellations were announced in July 2009, after these benefits were already set, and became effective in January 2010.

[20] HMO risk scores could appear higher artificially if HMOs engage in more up-coding. This is discussed further in Section 1.7.

Table 1.1: Cancelled and incumbent plans' 2009 characteristics, by plan type

	Cancelled PFFS plans	Other PFFS plans	HMOs	PPOs
Out-of-pocket cost ($'s)	296.937	305.633	241.160	252.813
	(28.427)	(21.573)	(44.457)	(30.527)
Contract Age	3.779	5.217	15.536	3.765
	(2.124)	(1.325)	(6.757)	(1.276)
Premium ($'s)	24.265	13.174	20.130	38.355
	(38.632)	(24.466)	(36.654)	(43.486)
Risk Score	0.939	0.914	1.021	0.889
	(0.136)	(0.085)	(0.133)	(0.096)
HMO or PPO in county	0.242	0.296	.	.
	(0.428)	(0.457)	.	.
Number of enrollees (unweighted)	77.468	103.548	679.001	233.238
	(190.527)	(203.863)	(2474.901)	(517.067)
Observations	10553	7063	6772	3023

Cancelled plans include terminated plans and plans whose service areas were reduced. Observations on the plan-county-year level. Weighted by enrollment. Contract age reflects the first year the contract was established with CMS. Plans are established within contract and can be created or terminated with affecting plan age.

by cancellation, and cancelled plans enrolled a substantial portion of the average MA market. For instance, between 2009 and 2010, 19% of MA beneficiaries (Row 1) were enrolled in a cancelled plan, and 77% of counties were affected by cancellation (Row 3).

Cancellation had a significant effect on competition, as measured by HHI. Row 4 of Table 1.2 show results from regressing the annual change in HHI on all cancelled plans' MA market shares separately by year; the final column shows results of regressing the cumulative change in HHI (2010-2012) on the cumulative share of enrollees in cancelled plans. Results suggest that each additional percentage point of cancellation increased HHI by an average of 13-26 points (on an average base of 4,464). For instance, nineteen precent of enrollees were in cancelled plans in the average county in 2010, so cancellation increased HHI in these counties by 258.40 points. This increase is well above the threshold

Table 1.2: Cancellation's impact on markets

	2009-2010	2010-2011	2011-2012	Cumulative
Cancelled Plans' MA Market Share	0.19	0.11	0.05	
(SD)	(0.22)	(0.20)	(0.16)	
Number of cancelled plans	680	344	181	
Number of counties affected by cancellation (out of 2,592)	2004	1366	430	
Cancellation's Marginal Effect on HHI	13.60**	13.35**	25.20**	17.09**
	(1.13)	(1.79)	(2.90)	(1.16)
Cancellation's Marginal Effect on Enrollment	-0.33**	-0.43**	-0.33**	-0.34**
	(0.02)	(0.02)	(0.03)	(.02)

Cancelled plans include terminated plans and plans whose service areas were reduced. The top panel presents summary statistics at the county level. The bottom panel presents regressions of HHI and county-level enrollment on total cancelled plans' share. $^{**}p \leq 0.05$. *Last column captures cumulative changes in HHI and enrollment (2010-2012). Observations are on the county-level. Regressions are unweighted.*

at which the Department of Justice scrutinizes mergers and acquisitions for anti-trust effects.(38)

The data suggest that roughly a third of cancelled plans' enrollees left MA for traditional Medicare. Cancellation's effect on overall MA enrollment is evaluated by regressing county-level MA penetration – the percent of all Medicare beneficiaries enrolled in MA – on cancelled plans' market share (Row 5, Table 1.2). Each additional percentage point of cancelled plans' share was associated with a 0.3 percentage point decline in MA penetration, suggesting that only a third of cancelled plans' enrollees switched to traditional Medicare.[21]

The main independent variable of interest is the share of insurers who cancelled all PFFS plans between 2010-2012 ($S_{m(2009)}$ in Equation (1.2)). Table 1.3 lists the 13 insurers that cancelled all plans and summarizes the number of counties affected by their cancellation. Many of these insurers offered plans in hundreds of counties. For instance,

[21] This is consistent with results from Sinaiko, et. al. (2014), who also find most cancelled PFFS plans' enrollees re-enroll in another MA plan.(39)

Table 1.3: Number of counties in which insurers canceling all PFFS plans offered a PFFS plan, 2008-2012

Insurer	**2008**	**2009**	**2010**	**2011**	**2012**
Coventry*	1363	1457	0	0	0
HealthNet*	200	134	0	0	0
Wellcare*	710	694	0	0	0
Blue Cross Blue Shield of Massachusetts	8	11	0	0	0
Blue Cross Blue Shield of South Carolina	104	106	37	0	0
Blue Cross Blue Shield of Michigan	85	85	83	0	0
Aetna*	107	202	61	0	0
Harvard Pilgrim Health Plan	18	17	19	0	0
Cigna*	22	70	410	0	0
Geisinger	14	19	26	0	0
Blue Cross Blue Shield of Florida	31	34	31	0	0
Blue Cross Blue Shield of Idaho	18	23	29	4	0
Wellpoint*	762	739	692	163	0

*Table summarizes number of counties in sample in which each insurer operated, by year. * indicates an insurer included in specification limiting $S_{m(2009)}$ to insurers operating in ≥ 5 states.*

Coventry cancelled plans in half (n=1457) of all counties in the sample simultaneously. Others were only operating in a small number of counties (i.e., Blue Cross Blue Shield of Idaho, Blue Cross Blue Shield of Michigan.) It is unlikely that observed changes in benefits are driven by unobserved variation in these insurers' markets, as they collectively cover a large and geographically dispersed set of counties. However, in robustness checks, the set of cancelled plans is limited to insurers offering PFFS plans in 5+ states (Section 3.6).

Figure 1.2 provides evidence that $S_{m(2009)}$ varies substantially across markets, and that affected markets are geographically distributed. $S_{m(2009)} = 0.20$ in the average county (SD=0.23), and affected counties are distributed across the country, with substantial variation in $S_{m(2009)}$ within regions.

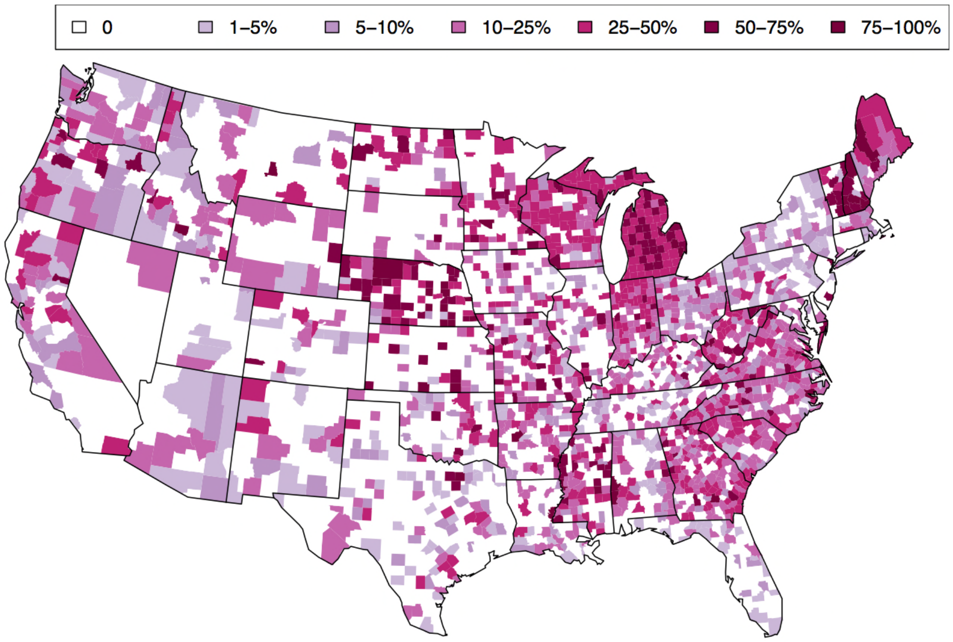


Figure shows the percent of the MA market enrolled in nationally cancelled PFFS on the county level.

Figure 1.2: Distribution of nationally cancelled plans' 2009 shares across counties

1.5 How cancellation affected premiums and benefit generosity

Table 1.4 shows results of regressing out-of-pocket cost and premiums on nationally cancelled plans' shares (S_{2009}). As described in Section 1.3, I first document the overall effect of cancellation on all plans' benefits. Next, to separate the direct effects of the policy from equilibrium effects, cancellation's impact is allowed to vary by plan type. I then offer a potential explanation for why insurers might change benefits rather than premiums and test several assumptions underlying main results. Mechanisms are discussed in Sections 1.6 and 1.7.

Among all plans in the sample, (Column 1, Table 1.4), cancellation led to consistent increases in out-of-pocket costs (decreased benefit generosity). The coefficient on $Post_{mt} * S_{m(2009)}$ suggests that, for each additional percentage point of the MA market enrolled in

Table 1.4: Effect of cancellation on out-of-pocket cost and premiums

VARIABLES	(1) Baseline OOPC	(2) No Cancelled Plans OOPC	(3) By Plan Type OOPC	(4) No New Plans OOPC	(5) No Cancelled Plans Premium	(6) By Plan Type Premium	(7) Zero-Premium OOPC	(8) National Plans OOPC
S_{2009}*Post	53.83**	42.26**	3.87	1.51	18.43**	17.67	28.76**	
	(8.60)	(9.20)	(10.48)	(10.63)	(5.87)	(9.87)	(9.87)	
S_{2009}*PPO*Post			36.97**	33.10**		-14.30		
			(12.28)	(15.55)		(15.42)		
S_{2009}*PFFS*Post			85.29**	94.27**		14.46		
			(15.34)	(17.57)		(10.41)		
S_{2009}*0 Premium*Post							45.52**	
							(14.74)	
$S_{Nat'nal}$*Post								0.39
								(16.17)
$S_{Nat'nal}$*Post*PPO								44.77**
								(16.61)
$S_{Nat'nal}$*Post*PFFS								105.59**
								(19.78)
Medicare benchmark	-0.19**	-0.18**	-0.18**	-0.22**	-0.10**	-0.10**	-0.20**	-0.18**
	(0.05)	(0.05)	(0.05)	(0.05)	(0.02)	(0.02)	(0.05)	(0.05)
PFFS plan	45.80**	49.93**	43.94**	33.96**	1.65	0.19	47.42**	43.56**
	(3.52)	(4.33)	(4.52)	(3.55)	(3.19)	(3.55)	(4.13)	(4.56)
PPO plan	4.48**	5.36**	3.64	0.48	10.44**	11.56**	4.23	4.05
	(2.25)	(2.30)	(2.53)	(2.65)	(3.73)	(4.33)	(2.41)	(2.56)
Lagged FFS costs	0.04**	0.04**	0.04**	0.04**	-0.00	-0.00	0.03**	0.04**
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Log 65+ population	-17.37	-23.44	-23.44	-25.06	-23.33**	-23.53**	-26.14	-23.21
	(12.40)	(13.26)	(13.09)	(13.96)	(8.02)	(7.99)	(13.61)	(12.97)
Percent below poverty	-1.63**	-1.88**	-1.77**	-1.27**	-0.77**	-0.76**	-2.03**	-1.70**
	(0.42)	(0.46)	(0.45)	(0.47)	(0.26)	(0.26)	(0.47)	(0.45)
Unemployment rate	-1.88**	-2.03**	-2.22**	-2.31**	-0.76**	-0.80**	-1.99**	-2.37**
	(0.78)	(0.86)	(0.84)	(0.87)	(0.35)	(0.35)	(0.89)	(0.84)
Employer-sponsored share	-4.22	0.99	-3.06	-8.49	-8.47	-9.52	5.38	-6.15
	(11.37)	(13.21)	(12.87)	(14.21)	(7.24)	(7.27)	(14.34)	(12.65)
Special needs plans' share	-9.99	-7.45	-7.53	-15.85	-3.19	-2.96	2.02	-6.30
	(27.25)	(30.98)	(31.05)	(33.75)	(8.50)	(8.43)	(31.55)	(31.12)
Log hospital beds per 1000	-12.36**	-13.47**	-13.30**	-11.36**	-4.67**	-4.62**	-13.96**	-13.15**
	(3.49)	(4.05)	(3.96)	(4.25)	(1.73)	(1.72)	(3.98)	(3.94)
Per capita income ($1000's)	0.66	0.60	0.59	0.70	0.42**	0.41**	0.51	0.58
	(0.40)	(0.43)	(0.43)	(0.43)	(0.16)	(0.16)	(0.45)	(0.43)
0 Premium							-3.53	
							(2.73)	
Observations	168,932	128,570	128,570	92,356	128,570	128,570	116,793	128,570
Excludes Exiters	No	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Excludes New Plans	Yes	Yes	No	Yes	No	No	No	No

OOPC=Out-of-pocket cost, less premiums. Higher numbers indicate less generosity. Premium is the MA plan premium, less any reductions in the Medicare Part B premium. FFS=Fee-for-service. S_{2009} is nationally cancelled plans' 2009 share. $S_{Nat'nal}$ excludes regional and local insurers. Observations on the plan-county-year level. Weighted by average enrollment over all years. All regressions include year and county fixed effects. SEs clustered on the plan-level. $^{**}p \leq .05$

cancelled plans, out-of-pocket costs increased by about 54 cents. As cancelled plans had 20% market share in the average county, this point estimate implies that beneficiaries in the average county paid an additional $10.80 out-of-pocket per month due to cancellation. This is a third of the total change in out-of-pocket costs post-2010, and constitutes a 20% reduction in average MA generosity, relative to traditional Medicare.[22]

Coefficients on time-varying controls generally have expected signs and reasonable magnitudes. Out-of-pocket costs are higher in counties with higher risk-normalized Medicare fee-for-services costs (i.e., counties where beneficiaries use more medical care). The coefficient on lagged fee-for-service costs indicates that each additional dollar in traditional Medicare spending decreases plans' overall generosity by about 4 cents. Higher benchmarks and increases in the number of hospital beds decrease out-of-pocket costs. The coefficient on logged hospital beds per 1000 people is -12.3, indicating that a percentage point increase in the supply of hospital beds in a county increases benefit generosity by 12 dollars.

The coefficient on benchmarks is $\approx -.20$, indicating that plans pass on 20 cents of each additional dollar of Medicare payments.[23] This estimate implies that the government would need to increase benchmarks by an additional $2.70 to offset the reduction in benefits caused by each additional percentage point of PFFS cancellation. As cancelled plans enrolled 20% of the average MA market, this translates into a $52 increase in average benchmark payments per-member-per-month. This is more than double the average benchmark cut implemented under the ACA.

As discussed in Section 1.3, several possible mechanisms might explain changes in

[22]Recall that average relative generosity of MA is calculated by comparing out-of-pocket costs in Traditional Medicare to out-of-pocket costs in all MA plans. Average relative generosity during the study period was about $50 a month.

[23]This finding is broadly consistent with Stockley, et al. (2014), and slightly lower than estimates in Song, et al. (2012), possibly due to payment freezes implemented during the study period.(29, 40) In contrast, Duggan, et. al. (2014) find no effect of increased benchmarks on benefits. This may be because they are using a regression discontinuity design and therefore have a more limited source of variation.(21)

benefits. First, as shown in Section 1.4, cancelled PFFS plans were more generous at baseline than other PFFS plans (but not HMOs or PPOs). If differences in generosity were large enough, average out-of-pocket costs might increase mechanically due to these plans' removal. To test this, I exclude any plan ever cancelled or withdrawn from the sample (Column 2). Excluding all cancelled plans is a conservative restriction as cancellation occurred in multiple waves. Hence, a plan cancelled in 2011 or 2012 might have been a competitor to a plan cancelled in 2010. However, the point-estimate of $\beta_1 \approx \$42.26$ in this reduced sample is smaller but statistically the same as estimates in Column 1, suggesting that observed changes are not driven by the removal of generous PFFS plans.

Changing costs are another mechanism that might decrease generosity. In addition to the fixed cost of building networks, the policy change may have permanently increased the prices PFFS plans paid providers, by removing their ability to pay traditional Medicare rates.[24] Insurers may have passed on these increased costs to beneficiaries by reducing benefits. To test this, I allow cancellation's effects to vary by plan type. Though changes in PFFS plans' benefits may be directly driven by changes in provider prices, generosity among HMOs/ PPOs should not be.

Effects are allowed to differ across plan types by interacting $Post_{mt} * S_{m(2009)}$ with indicators for type (Column 3 – HMOs are the omitted category). This tests whether cancellation's effects on out-of-pocket costs differ based on whether the plan is an HMO, PPO, or PFFS plan.[25] Results show that benefit changes were not limited to PFFS plans. Specifically, generosity decreased among both PFFS plans and PPOs (but not HMOs). PPOs, which were not directly affected by the policy, increased out-of-pocket costs by 37

[24] Recall that administratively-set rates paid to providers by Medicare are thought to be lower than commercially-negotiated prices paid by private insurers.

[25] In theory, this regression could also include an additional term interacting plan types with a *Post* variable, to distinguish benefit changes among all PFFS and PPO plans after 2010 from benefit changes among PPOs and PFFS plans in more heavily impacted counties. In practice, the data are not sufficient to do this, as there are relatively few PFFS and PPO plans in counties with no cancellation and plan types are unequally distributed across pre- and post- periods.

cents for each additional percentage point in cancelled PFFS plans' share.

The law clearly had more direct impact on PFFS plans, as it forced PFFS plans to build networks and increased the prices they paid to providers. However, changes in PPO benefits might reflect the costs of building networks if insurers replaced their PFFS plans with new PPOs.[26] To test this hypothesis, the sample of plans is restricted to those introduced prior to any cancellation (2010). This limitation cuts the sample by almost a third, but coefficients are similar to those in main results (Column 4). Most notably, the coefficient reflecting incumbent PPOs' response to cancellation is $\beta_{PPO} = 33.10$, statistically the same as the coefficient among all PPOs.

In contrast to benefits, premium changes are smaller and only significant for PFFS plans. Premiums increased significantly in counties with more cancellation (Column 5), but average increases were a third as large as increases in out-of-pocket costs. ($\beta_1 = 18.42$). In the specification where effects vary by plan type (Column 6), changes are only significant for PFFS plans ($\beta_{PFFS} = 17.67 + 14.46 = 32.13$, F-statistic=22.76).[27]

These results show that insurers adjusted benefits more than premiums, but do not reveal why. Discussion in Section 1.1 suggests that insurers may be more likely to modify benefits when constrained in adjusting premiums. This hypothesis is tested by allowing effects on out-of-pocket costs to differ between plans that charge $0 or negative MA premiums and plans that charge positive premiums. If Medicare beneficiaries find a $1 increase in premiums more salient when plans charge $0 at baseline, then insurers charging a $0 premium may prefer to modify benefits.(29) To test this, I generate an indicator equal 1 if a plan charges $p \leq \$0$ premium in time $t-1$.[28] This indicator is

[26] Some insurers, including those who continued to offer PFFS plans, seem to have done this. For instance, the percent of Humana enrollees in PPOs grew from 13% in 2009 to 29% in 2012, while the percent of enrollees in PFFS plans fell from 42% to 16%.

[27] Coefficients are also positive for HMOs ($\beta_{HMO} = 17.67$), but only significant at the $p < .1$ level.

[28] 72,036 plan-county-year observations, or 47% of the sample, had $0 premiums in time $t-1$.

added to Equation (1.2) as a control and interacted with $S_{m2009} * Post$.[29]

Results confirm the hypothesis that plans charging $0 premiums were more likely to adjust benefits (Column 7). Plans charging positive premiums modify benefits in response to cancellation, increasing out-of-pocket costs by 28.76 cents for each additional percentage point of cancellation. Plans that charge $0 premium modify benefits by substantially more, with out-of-pocket costs rising by (28.76+45.53)=74.29 cents for every additional percentage point of cancellation.[30]

Before discussing additional mechanisms driving observed changes, I test several underlying assumptions. First, the key identifying assumption of difference-in-differences is that pre-period trends in outcomes are parallel across markets differentially affected by cancellation. Figure 1.3 shows trends in unadjusted, unweighted average out-of-pocket costs, grouped by levels of $S_{m(2009)}$. Prior to 2010, there is variation in levels of out-of-pocket costs, but trends appear parallel.

This assumption is further explored by allowing treatment to vary by year in the post period and testing whether $S_{m(2009)}$ affects benefits in 2008. Neither premiums nor benefits were significantly affected by $S_{m(2009)}$ in 2008 (Columns 1 and 2 in Table 1.5). Additionally, coefficients on both out-of-pocket costs and premiums in the post-treatment years change as expected, increasing between 2010-2012. This is consistent with both a lagged response and the fact that cancellation affected some markets in later years (2011 or 2012).

To further explore whether changes in premiums and benefits are due to underlying trends, I interact a linear time trend with $S_{m(2009)}$ and test whether (linear) trends

[29]New plans are excluded from this regression, as they have no premium in time $t-1$. To include plans from 2007, I assume that plans with $0 premiums in 2007 also had $0 premiums in the prior period. This assumption is reasonable given that premiums are highly autocorrelated ($\rho = .91$), but is tested by omitting 2007 (Appendix Table A1). Results are similar.

[30]Appendix Table A1 Column 4 shows the same regression, using premiums as an outcome. Plans that charge positive premiums are not statistically more likely to adjust premiums, although this null result may be caused by the $0-premium indicator absorbing most of the variation in premiums.

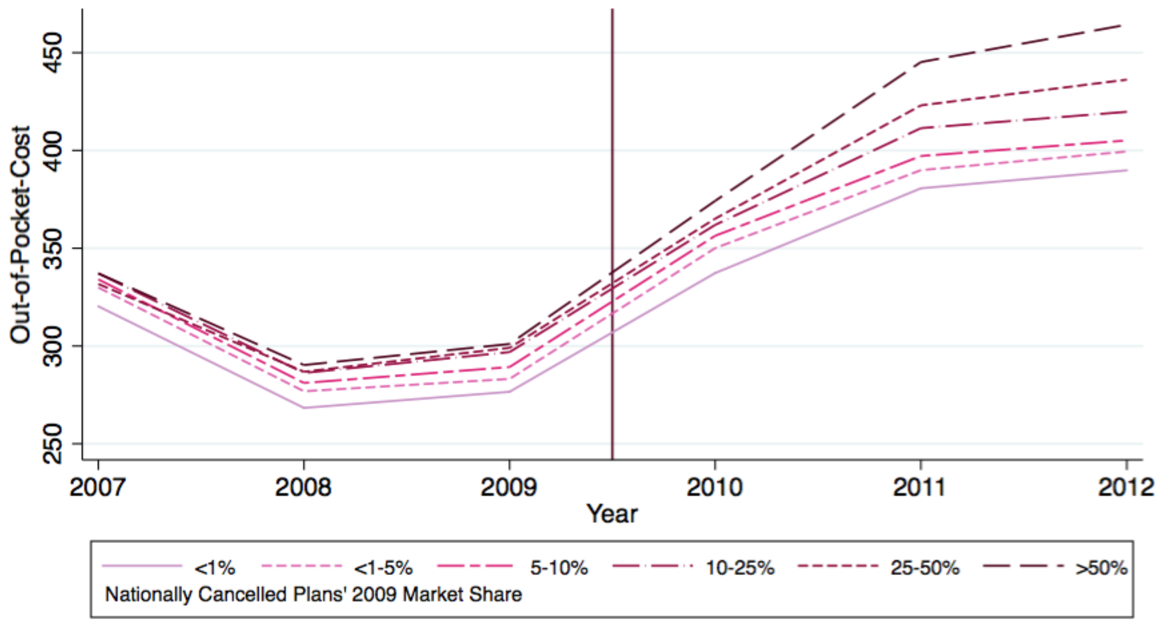


Figure 1.3: Out-of-pocket costs during the study period (2007-2012)

Table 1.5: Placebo tests

VARIABLES	(1) OOPC	(2) Premium	(3) OOPC	(4) Premium
$S_{2009} * (t-1)$	-6.68	1.41		
	(12.85)	(4.68)		
$S_{2009} * (t+1)$	35.13**	20.37**		
	(16.88)	(8.59)		
$S_{2009} * (t+2)$	47.66**	20.49**		
	(17.25)	(8.81)		
$S_{2009} * (t+3)$	77.75**	27.95**		
	(19.51)	(8.83)		
Linear pre-trend			-1.09	2.57
			(2.87)	(3.48)
Observations	128,570	128,570	79,921	79,921
County FE	Yes	Yes	No	No

OOPC=Out-of-pocket cost, less premiums. Higher numbers indicate less generosity. Premium is the Medicare Part C premium, less any reductions in the Medicare Part B premium. Observations on the plan-county-year level. Weighted by average enrollment over all years. Plans ever cancelled are excluded. SEs clustered on the plan-level. $^{**}p \leq .05$

differ across markets prior to cancellation. They do not (Table 1.5, Columns 3 and 4). Coefficients are small ($\beta_t = -1.09$ for out-of-pocket costs and $\beta_t = 2.572$ for premiums) and insignificant.

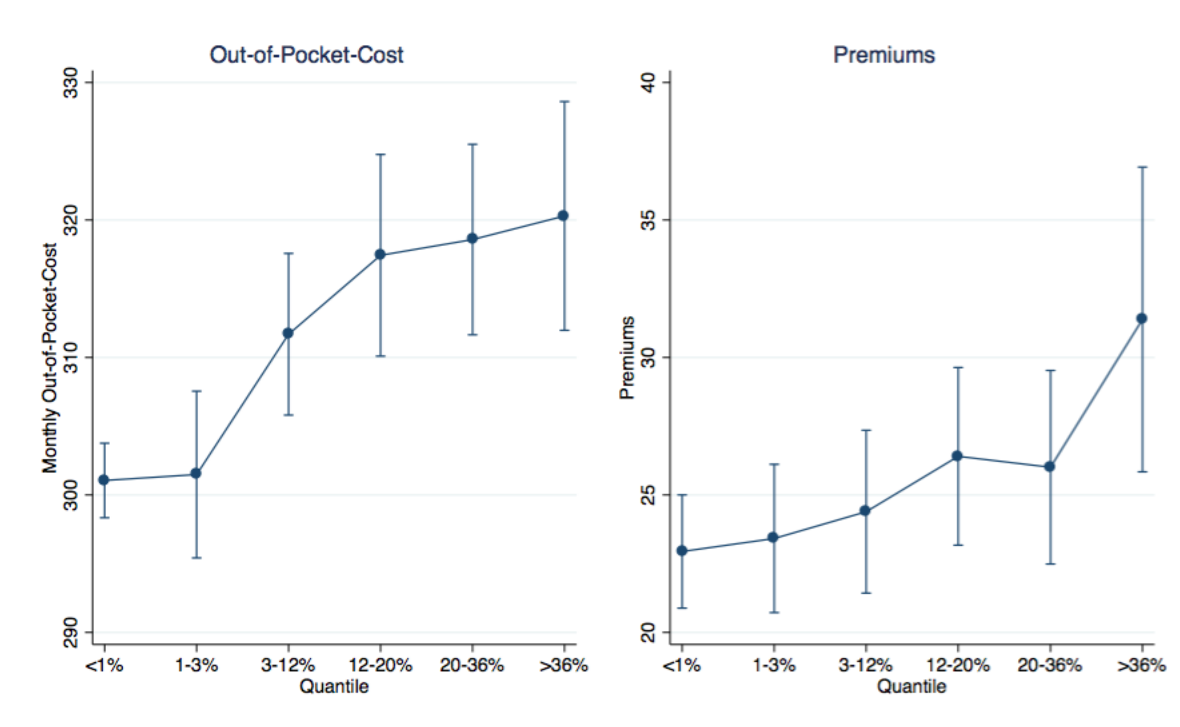
Another assumption made in Equation (1.2) is that cancellation's effects are linear in $S_{m(2009)}$. Imposing linearity may overstate effects if the most heavily impacted markets drive observed changes. To test this, responses are allowed to differ across levels of exit, by interacting indicators for quantile of cancelled share with *Post*.[31]

Results (Figure 1.4 and Appendix Table A2) show that changes in generosity are observed at most levels of cancellation and changes in generosity are not driven by plans in the top quantile of counties. Plans in counties with $< 1\%$ cancellation and $1-3\%$ cancellation do not significantly adjust benefits, while plans in all other counties do ($p < .001$). Premiums, in contrast, are noisy, and effects seem largely driven by the top quantile of cancellation.

A final assumption behind baseline specifications is that insurers who cancelled all PFFS plans were not selectively exiting less profitable markets. However, as discussed in Section 1.4, some of these insurers only operated on the region or state level. As their decision to cancel might be a response to changes in one large market, an alternate *Share* variable is constructed using only insurers who offered PFFS plans in more than 5 states.[32] Results show that selective exit by regional/state insurers do not drive benefit reductions (Column 8, Table 1.4). The coefficient on share, excluding regional plans, is 59.46 and significant. This magnitude is similar to coefficients from the baseline regression.

[31] Quantiles are generated by dividing plans into quintiles by level of 2009 cancellation and adding an additional category for plans in counties where cancelled plans had less than 1% market share.

[32] Coventry, HealthNet, Wellcare, CIGNA, Aetna, and Wellpoint.



Margins (predicted out-of-pocket costs and premiums) from regressing benefits on quantile of $S_{m(2009)} * Post$, 95% confidence intervals. Includes all controls plus year and state fixed effects. Quantiles defined based on dividing 2009 plans into 5 equal groups, plus one group containing plans in counties where $S_{m(2009)} < 1\%$.

Figure 1.4: Margins of out-of-pocket cost and premiums, by level of exit

1.6 Competitive effects

Results from Section 1.5 show that plans reduced benefit generosity in response to cancellation and reduced it by more in more heavily affected counties. Results also suggest that benefit changes were not driven primarily by insurers passing-through the policy's increased costs. Specifically, decreased benefits are observed among both PFFS plans and PPOs, even though PPOs were not directly impacted by the policy.

There remain several other mechanisms that might cause insurers to reduce benefit generosity. The first relates to market competition. PPOs were likely the closest substitutes to PFFS plans, as they had more expansive networks than HMOs and allowed beneficiaries to seek out-of-network care.[33] Insurers offering PPOs may have gained

[33]Most HMOs do not cover any care received out of network.

the most market power from PFFS plans' cancellation and may have been most able to increase mark-ups.

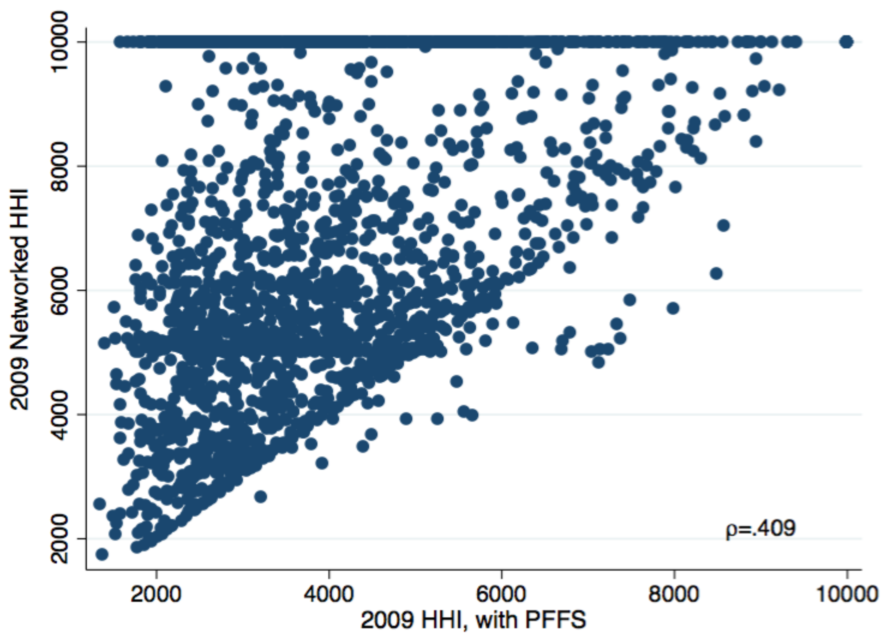
The fact that PPOs had significantly lower average risk scores than cancelled PFFS plans at baseline (see Table 1.1) suggests two alternate mechanisms that might drive changes in benefits. First, PPOs might have enrolled cancelled plans' beneficiaries and passed on higher costs of care by reducing benefits. Alternately, if insurers had information about cancelled plans' risks, they might have adjusted benefits in advance to deter sicker enrollees from choosing their plans.[34]

To distinguish between different mechanisms, I first test the competition theory directly. Specifically, I investigate whether cancellation's effects vary with counterfactual concentration – or the level of concentration in a market when all PFFS plans are removed. Then, I test remaining theories about changes in enrollee health risks using data on risk scores.

If changes in competition drove decreased generosity, then plans gaining more market power should have more latitude to reduce benefits. As the policy change essentially disallowed PFFS plans, changes in generosity should vary based on competition between PFFS substitutes. Moreover, theory and past empirical work suggest that changes will be greatest in markets where only one or two firms offered substitutes for PFFS.(41, 42)

To measure concentration among PFFS plans' substitutes, I construct an alternate HHI using insurers' market shares of networked plans (HMOs or PPOs). This measure is used to test whether counties with higher "networked HHI" in 2009 reduce their benefits by more in post-cancellation years. Figures 1.5-1.7 summarize key information about networked HHI. First, Figure 1.5 shows that HHI and networked HHI are distinct concepts. The measures are correlated ($\rho = .409$), but much of this correlation is driven by the fact that networked HHI is almost always higher than HHI when PFFS plans are

[34] Insurers who cancelled plans could not automatically reenroll their beneficiaries into other plans, but were allowed to market their other plans to cancelled plans' enrollees. Enrollees who didn't choose a new plan defaulted into traditional Medicare.



2009 values of networked HHI and HHI, including PFFS plans. Networked HHI is HHI calculated only using firms' HMO and PPO market shares. Observations are on the county level. Plans in markets with no HMOs or PPOs in 2009 are omitted (n=465 markets, 18% *of all markets).*

Figure 1.5: Correlation between HHI and networked HHI

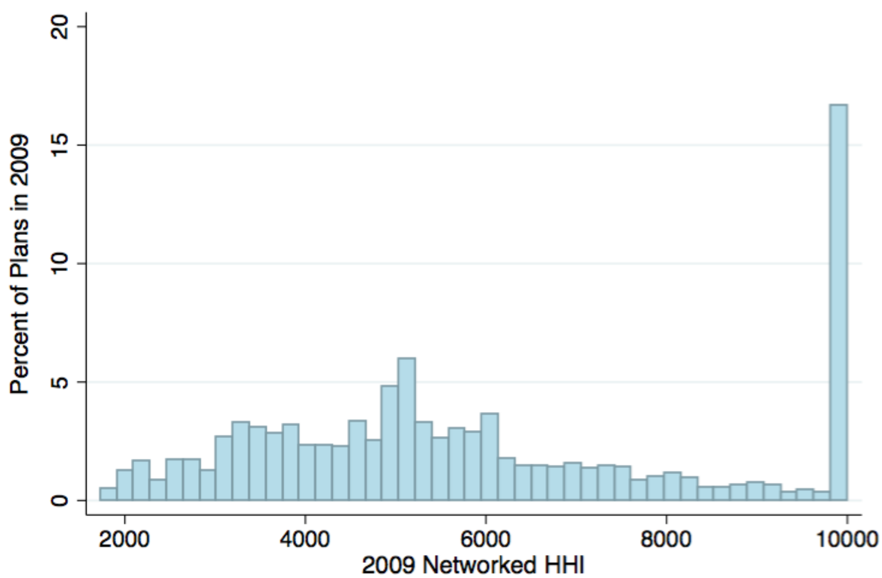
included.[35]

Second, though there is a substantial variation in networked HHI across plans (Figure 1.6), networked HHI is highly concentrated and its distribution is skewed. (Mean networked HHI=6951, SD=2477.) Just one firm controls all networked plans in 667 (25%) of all markets. A further 465 markets (18%) have no networked plans in 2009,[36] though HMOs or PPOs have entered all but 211 markets (8%) by 2012.

Lastly, networked HHI is not directly related to PFFS plans' shares in a county, so cancelled plans' shares vary within it. Figure 1.7 shows the relationship between 2009 networked HHI and nationally cancelled plans' 2009 market shares ($S_{m(2009)}$). Although

[35]Networked HHI can be lower than HHI with PFFS if a county is, for instance, dominated by one large firm providing a PFFS plan and several smaller firms providing HMOs/ PPOs. This is rare.

[36]Networked HHI for these counties is undefined and these counties do not appear in statistics in Figures 1.5-1.6. They are included in regressions testing for heterogenous effects across counties with different networked HHI by adding a indicator capturing when a plan is in a county with no networked plans at baseline.



2009 distribution of networked HHI. Networked HHI is HHI calculated using only firms' HMO and PPO market shares. Observations are on the plan-county level. Plans in markets with no HMOs or PPOs in 2009 are omitted (n=465 markets, 18% of all markets).

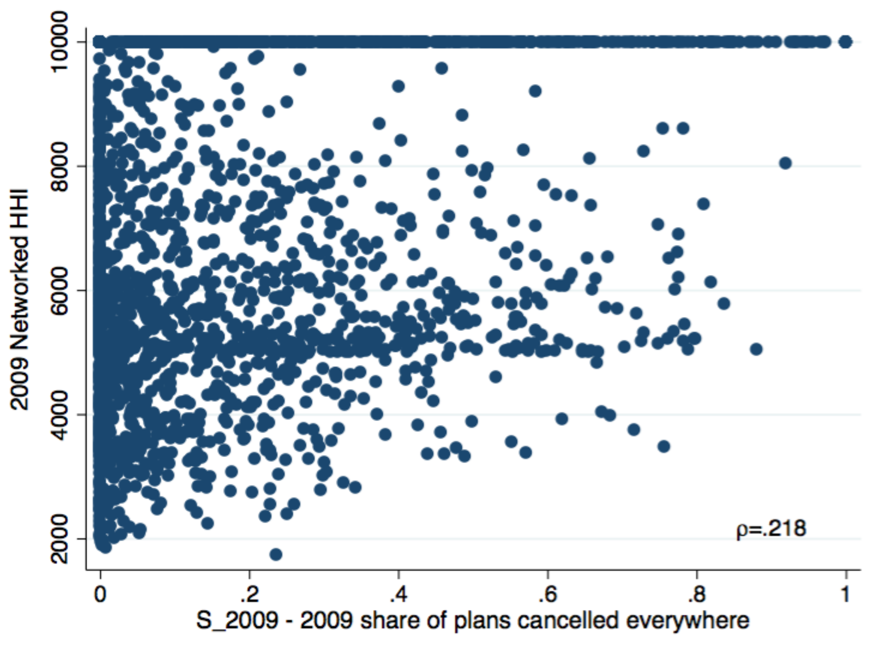
Figure 1.6: Distribution of networked HHI

networked HHI and $S_{m(2009)}$ are correlated ($\rho = .22$), there is substantial variation in $S_{m(2009)}$ for any fixed level of networked HHI, including within markets with only one networked firm. This variation in $S_{m(2009)}$ for a fixed level of competition is necessary to test whether responses to cancellation differ conditional on networked HHI.

To test whether cancellation's effects differ by counterfactual levels of competition, counties are divided into five categories based on 2009 values of networked HHI. Categories are defined by quartiles of networked HHI, with a fifth category for counties with no networked plans in 2009. Then, indicator variables capturing what type of county a plan is in are interacted with $S_{m(2009)} * Post$. The specification is:

$$Y_{jmt} = \beta_0 + \sum_{k=2}^{5} \beta_k Post_{mt} * S_{m(2009)} * D(m \in k) + \eta M_{mt} + \theta X_j + \gamma_m + \tau_t + \varepsilon_{jmt} \tag{1.3}$$

where $D(m \in k)$ is an indicator equal to 1 if a plan is offered in a county in group k (where the first category is omitted and baseline differences are absorbed by fixed effects). All other variables are as before, and, as in previous sections, plans that are eventually

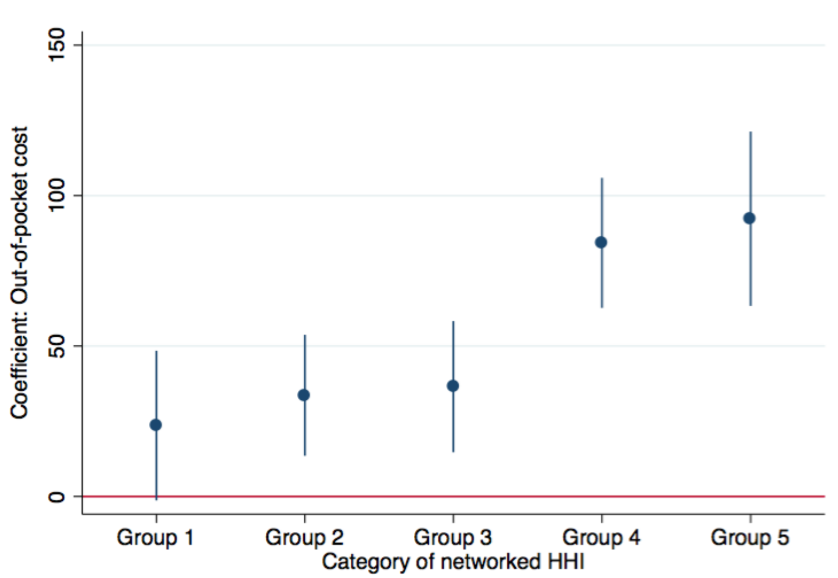


2009 values of networked HHI and shares of plans cancelled everywhere ($S_{m(2009)}$). Networked HHI is HHI calculated only using firms' HMO and PPO market shares. Observations are on the county level. Plans in markets with no HMOs or PPOs in 2009 are omitted (n=465 markets, 18% of all markets).

Figure 1.7: Relationship between 2009 networked HHI and share of nationally cancelled plans

cancelled are excluded. Identification in this regression comes from variation in $S_{m(2009)}$ within category of networked HHI, and coefficients β_k capture whether plans respond more to cancellation, conditional on counterfactual competition.

Firms responded more to cancellation in areas where plans stood to gain more market power (Figure 1.8 and Table A3). Benefit changes in competitive counties (Group 1, HHI$<$ 5031) were small and marginally significant ($\beta_1 = 23.5$, significant at the 10% level), while plans in less competitive counties (Group 2, 5031 $\leq$ HHI $\leq$ 6389 and Group 3, 6389 $\leq$ HHI $\leq$ 10000) increased out-of-pocket costs by an amount similar to the overall sample average ($\beta_2 = 33.62$ and $\beta_3 = 36.47$, significant at the $p < .001$ level). Consistent with theory, plans in counties where only one firm offered HMOs or PPOs (Group 4, HHI=10,000) modified benefits substantially ($\beta_4 = 84.27$, $p < .001$), as did plans in



Coefficients from interacting county group with share and plan type. 95% confidence intervals. Counties are divided by baseline HHI without PFFS. Group 1: HHI< 5031, Group 2: 5031 ≤HHI< 6389, Group 3: 6389 ≤ HHI< 10000, Group 4: HHI= 10000, Group 5: No networked plans at baseline.

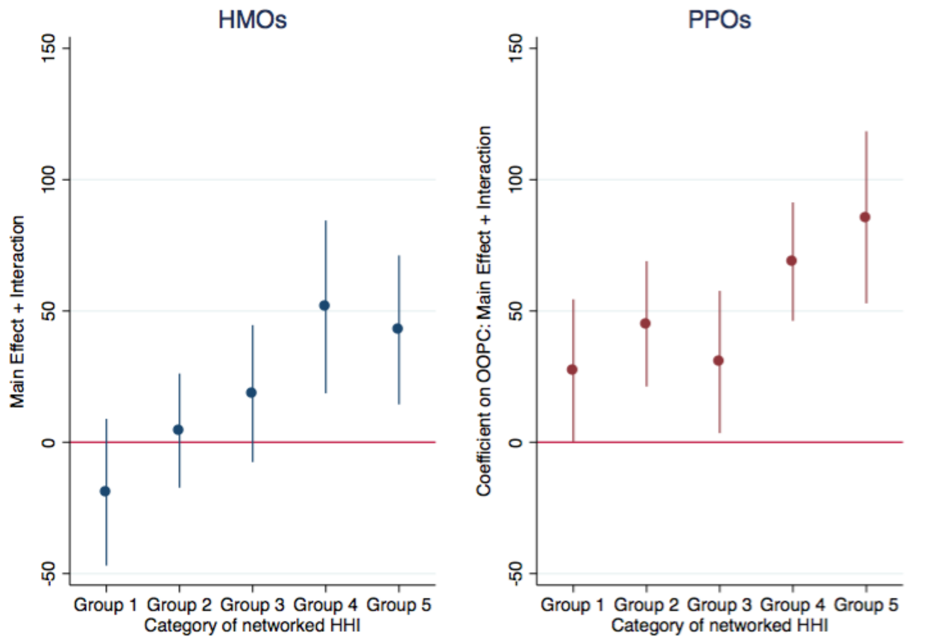
Figure 1.8: Response to cancellation, by category of HHI

counties with no networked plans at baseline (Group 5, $\beta_5 = 92.31$, $p < .001$).[37]

To test whether observed patterns are driven by remaining PFFS plans adapting to the policy, $S_{m(2009)}$ and HHI groups are also interacted with indicators for plan type. Results are noisy, but reflect the same pattern seen among all plan types (Figure 1.9 and Table A3). PPOs responded to PFFS plans' cancellation at all levels of networked HHI, but were most responsive in the least competitive markets (Groups 4 and 5). For instance, the total effect on PPOs in counties where only one firm offered an HMO or PPO (Group 4) was 68.78. As nationally cancelled plans enrolled a third of beneficiaries in this group ($S_{m(2009)} = .30$), this suggests that PPO out-of-pocket costs for a representative beneficiary in these counties increased by about $21 a month.

HMOs generally did not respond to cancellation, except in the least competitive markets (Groups 4 and 5). In these markets, HMO out-of-pocket costs increased by more

[37]The majority (87%) of observations in category 5 markets post-2009 are PFFS plans, but HMOs and PPOs do enter these markets between 2010-2012.



Coefficients from interacting county group with share and plan type. 95% confidence intervals. Counties are divided by baseline HHI without PFFS. Group 1: HHI$<$ 5031*, Group 2:* 5031 $\leq$*HHI*$<$ 6389*, Group 3:* 6389 $\leq$ *HHI*$<$ 10000*, Group 4: HHI*$=$ 10000*, Group 5: No networked plans at baseline.*

Figure 1.9: Response to cancellation among HMOs and PPOs, by category of HHI

than the average PPO (β_4=51.55 and β_5=42.79, $p < .05$).[38] These coefficients suggest that HMO out-of-pocket costs for a representative beneficiary increased by and average of $.3 * 51.55 = \$15.46$ and $.3 * 42.79 = \$12.83$ a month, respectively.

1.7 Changes in enrollee health and risk deterrence

The remaining mechanisms that might explain decreased benefit generosity relate to changes in PPO and HMOs' risks. The first possible explanation is that plans staying in

[38]Networked competition in groups 4 or 5 might be low at baseline because of high provider costs; if insurers replaced PFFS plans with new HMOs or PPOs in these counties, they might offer less generous benefits to offset costs of building networks. The approach used here should be robust to this, as Equation (1.3) is identified using variation in $S_{m(2009)}$ within category of HHI. However, I use the approach from Section 1.5 to explore whether new plans drive results, by restricting the sample to plans introduced before 2010. The sample size is much smaller ($n = 90,671$), but observed effects are very similar to those in the full sample.

the market enrolled cancelled plans' sicker beneficiaries and subsequently passed on the higher costs of treating these beneficiaries in the form of reduced benefits. The second explanation is that insurers anticipated enrolling sicker beneficiaries and reduced benefit generosity in advance to deter beneficiaries from enrolling.

To test the first hypothesis, that plans staying in the market enrolled riskier beneficiaries, I test whether cancellation is correlated with changes in risk among HMOs and PPOs. To do this, changes in aggregate, county-level HCC risk scores are regressed on the percent of all Medicare beneficiaries in a county in cancelled plans – both those cancelled everywhere and those selectively removed from markets. Regressions are of the form:

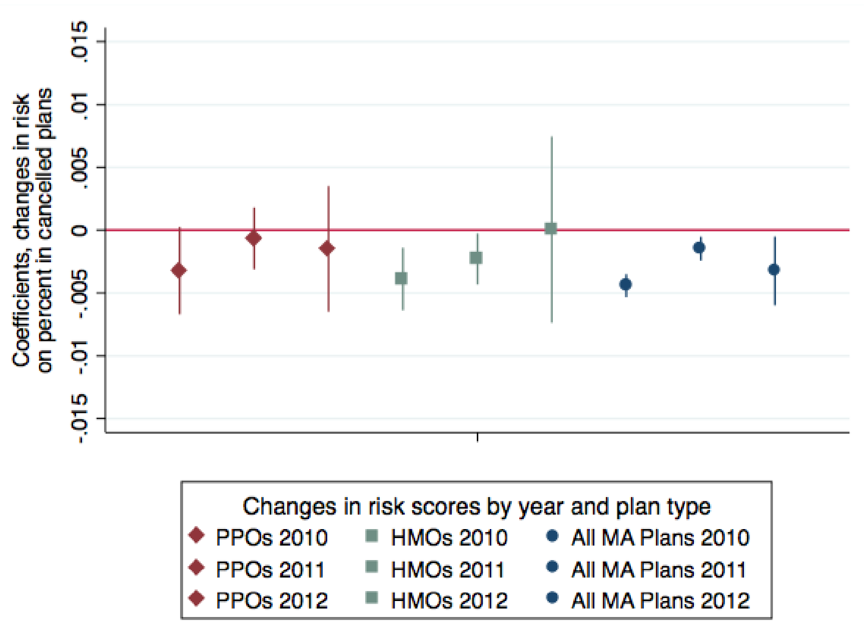
$$\Delta R_{jmt} = \beta_0 + \beta_1 Percent\_Cancelled_{m(t-1)} + \beta_2 \Delta Percent\_Employer_{mt} + \varepsilon_{mt} \tag{1.4}$$

where ΔR_{jmt} is the county-level change in average risk scores for plan type j in county m between time t and $t-1$. Changes in average, county-level risk scores among both HMOs and PPOs are used as outcomes. $Percent\_Cancelled_{m(t-1)}$ is the percent of all beneficiaries in a county in cancelled plans at time $t-1$.[39] β_1 reflects the correlation between cancellation and variation in HMO or PPO risk scores. $\beta_1 > 0$ would suggest that cancellation led to HMOs and PPOs enrolling worse risks, consistent with the hypothesis outlined above.

$Percent\_Employer_{mt(t-1)}$ is the percent of beneficiaries in a county in MA employer-sponsored plans. It is added as a control, as CMS does not publish county-level risk scores for employer- and non-employer plans separately.[40] Regressions are run separately by year, and only counties where $Percent\_Cancelled_{m(t-1)} > 0$ are included. Regressions are unweighted because the distribution of both the MA population and the total Medicare

[39] This includes all cancelled plans' enrollees, not just the nationally cancelled plans.

[40] These plans are only offered through direct contracts with employers, and therefore are not part of the overall MA market.



Coefficients from regressing changes in aggregate county-level risk scores on all cancelled plans' shares. Higher numbers indicate greater risk (less health.) All regressions control for changes in employer plans' shares as aggregate risk scores are not published separately for employer and non-employer plans.

Figure 1.10: Effect of cancellation on risk scores

population are skewed across counties.[41]

Results do not support the hypothesis that HMOs/PPOs drew sicker enrollees. (Coefficients by year and plan type are displayed in Figure 1.10 and Appendix Table A4.) Cancellation had no statistically significant effect on PPOs' risk in any year ($p < .05$).[42] Cancellation is correlated with negative shifts in HMOs' risk in 2010 or 2011 ($\beta_1 = -.003$ and $-.002$), implying that HMOs drew slightly *healthier* enrollees in these years. This is consistent with the fact that HMOs had higher risk scores than cancelled PFFS plans at baseline.

To test whether riskier beneficiaries are leaving the MA program altogether, changes

[41]The median county in 2009 had 6,195 Medicare eligibles, but the 95th percentile had over 70,000. Weighting by the number of Medicare eligibles in a county does not qualitatively change results. (See Appendix Figure A1 and Table A5.)

[42]Cancellation had a marginally negative effect on PPO risk in 2010, but this result was not significant at conventional levels. $p = .07$. Moreover, $\beta_1 < 0$ suggests that PPOs were drawing healthier enrollees.

in average, county-level risk among all plans are used as an outcome in Equation (1.4). Results from these regressions (Figure 1.10 and Appendix Table A3) suggest that cancelled plans' riskiest enrollees left MA for traditional Medicare. Cancellation is correlated with significantly lower overall MA plan risk in all post-treatment years (β_1=-.004, -.001, and -.003, $p < .05$). This implies that the enrollees who left MA for traditional Medicare after cancellation were sicker than the average.

Taken together, these results do not support the hypothesis that benefit changes among HMOs/PPOs are the result of plans passing on higher costs of care. However, using risk scores to test for changes in plans' risk has several limitations. First, risk scores only explain about 13% of the variation in total Medicare spending.(37) Though it is unclear how much of residual spending is systematic, it is still possible that plans' risk could systematically increase on unmeasured dimensions without any changes in risk scores.

Second, these tests are limited by the fact that MA plans can "upcode" or increase risk scores by encouraging providers to diligently record diagnoses.(43, 44) Most upcoding concerns focus on differences between MA and traditional Medicare. However, if different plan types systematically code with more intensity, then risk scores across types are hard to compare.[43] This concern is somewhat secondary in this context, as regressions use changes in risk scores rather than levels. If plans have baseline differences in coding, the movement of enrollees across plan types should still yield significant changes in risk scores. Such shifts are not observed among PPOs, which are the plans driving benefit reductions.

To test the second hypothesis, that plans adjusted benefits in advance to avoid attracting sicker enrollees, a measure of "counterfactual" or "expected" risk is constructed using risk scores. This measure is intended to capture each plan's incentives to distort

[43] For instance, PFFS plans might be less able to upcode than HMOs because they do not contract directly with doctors. If this is the case, than HMOs' risk might decrease, as it did in 2010 and 2011, even if HMO plans were enrolling less healthy beneficiaries.

benefits by measuring how much the plan's average enrollee health would change if it drew cancelled plans' enrollees.

Expected risk for each plan is measured using a weighted average of the plan's risk scores and the risk scores of all plans cancelled in its markets.[44] Expected risk for plan j is:

$$E(r_{j(t-1)}) = \frac{r_{j(t-1)} * N_{j(t-1)} + \sum_{i \in M} r_{i(t-1)} N_{i(t-1)}}{N_{j(t-1)} + \sum_{i \in M} N_{i(t-1)}}$$

where $r_{j(t-1)}$ is plan j's pre-cancellation risk score, and $r_{i(t-1)}$ is the risk score of plan i, cancelled at time t. M is the set of plans cancelled in j's markets, and $N_{j(t-1)}$ and $N_{i(t-1)}$ are the number of enrollees in plan j and cancelled plan i, respectively.

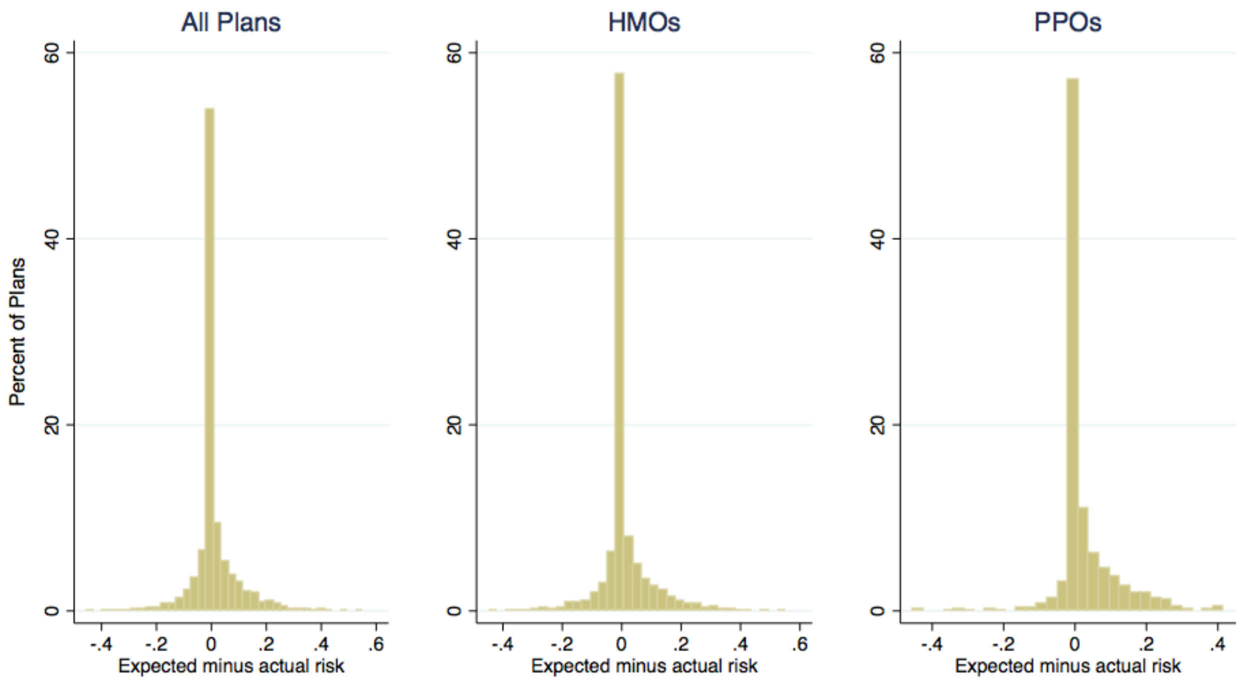
To capture how much risk a plan could expect to gain (or lose), expected risk is compared to actual risk, prior to cancellation ($\Delta R = E(r_{j(t-1)}) - r_{j(t-1)}$). One limitation in using risk scores to measure plan expectations is that Medicare adjusts plan payments based on risk scores. If plans are adequately compensated for measured risk, then they should be indifferent to changes in risk score. However, recent evidence suggests that Medicare's risk adjustment may undercompensate plans for risky enrollees.(45, 46) If this is the case, then plans still have an incentive to avoid high-risk enrollees.[45]

Figure 1.11 shows the distribution of ΔR for plans in counties with any cancellation; for most plans, ΔR is nearly 0.[46] However, gains in risk are potentially large for a subset of plans. Five percent of plans would gain more than a standard deviation of risk

[44] This assumes that plans draw all cancelled enrollees in a county and no enrollees default back to traditional Medicare. Making alternate assumptions that some enrollees default to traditional Medicare or that plans draw enrollees proportionally to their market shares generates smaller but similar estimates of expected risk.

[45] Another limitation of using risk scores to test whether plans distorted benefits is that they do not capture all predictable spending in Medicare. However, there is substantial debate about the degree to which MA plans can distort benefits to select favorable risks conditional on risk adjustment.(47–51) Specifically, plans do not seem to cream-skim within disease categories,(52) suggesting that plans still have incentives to avoid beneficiaries in overall worse health.

[46] 35% of all plans and 44% of HMOs /PPOs would gain less than a hundredth of a standard deviation of risk ($\sigma = .156$). 56% of all plans and 60% of HMOs/PPOs would gain less than a tenth of a standard deviation.



Plan-level weighted average of ΔR, assuming each plan draws all cancelled plans' enrollees. Includes only plans in counties with some cancellation. Excludes new plans and cancelled plans by construction. Standard deviation of ΔR=.08. Standard deviation of plan level risk scores=.156.

Figure 1.11: Potential gain in risk for plans staying in the market

($\sigma = .156$), and one percent stood to gain more than two standard deviations.

To test whether plans with potentially large changes in risk reduced benefit generosity, ΔR is added as a control to Equation (1.2). Coefficients on ΔR are identified by variation across plans within counties. ΔR also varies across time, as a different set of plans cancelled in each county in each year. By construction, ΔR does not exist for new or cancelled plans,[47] so Table 1.6, Column 1 shows baseline results excluding these plans for comparison.

Table 1.6 Column 2 shows results including ΔR as a control. The coefficient on ΔR is large and positive but insignificant, and its inclusion does not substantially change the relationship between $S_{m(2009)}$ and out-of-pocket costs.

[47]New plans have no baseline risk ($r_{j(t-1)}$) and cancelled plans cannot draw worse risks.

Table 1.6: Effect of the gap in risk on benefit generosity, without premiums

VARIABLES	(1) Baseline OOPC	(2) Controlling for Risk Gap OOPC	(3) Controlling for Risk Indicator OOPC	(4) Controlling for Risk Indicator and Own Plan Indicator OOPC
$S_{2009} * Post$	18.20	16.38	10.57	9.99
	(11.53)	(11.67)	(11.48)	(11.36)
$S_{2009} * Post * PPO$	33.70**	34.03**	32.99**	32.00**
	(12.92)	(12.91)	(12.30)	(12.03)
$S_{2009} * Post * PFFS$	63.08**	64.56**	66.41**	67.30**
	(17.29)	(17.34)	(17.58)	(17.52)
ΔR		21.71		
		(13.39)		
$\Delta R > 0$			11.95**	11.40**
			(2.53)	(2.61)
$\Delta R > 0 * Ownplan$				7.10**
				(2.91)
Observations	83,494	83,494	83,494	83,494

OOPC=Out-of-pocket cost, less premiums. Higher numbers indicate less generosity. Regressions exclude new plans and cancelled plans by construction. Observations on the plan-county-year level. Weighted by average enrollment over all years. All regressions include year and county fixed-effects and all baseline controls. SEs clustered on the plan-level. $^{**}p \leq .05$

ΔR is likely a noisy measure of plans' expectations, which might bias coefficients towards 0. To address this, expected changes in risk are also measured using an indicator equal to 1 if plans stood to gain risk. When this measure is used as a control (Column 3), results indicate that plans that might attract sicker enrollees increased out-of-pocket costs by about \$12, relative to plans that wouldn't gain risk. This result is significant, but again, does not substantially change cancellation's effects.

To explore whether ΔR is a reasonable measure of plan expectations, I test whether plans with more information respond more to potential increases in risk. Specifically, insurers who offered another plan in a market where they cancelled a PFFS plan plausibly had more information about potential enrollees' health. To capture variation in information, $1(\Delta R > 0)$ is interacted with an indicator for when plan j is offered by an insurer canceling their own PFFS plans in market m ($Ownplan_{jmt}$). The coefficient on this interaction (Column 4) shows that plans with more information were more responsive to potential gains in risk; plans with more information increased out-of-pocket costs by about $18 on average, more than the $11 increase observed among insurers who did not have information on cancelled plans' health risks. This helps validate ΔR as a measure of plan expectations. However, controlling for ΔR among plans with added information still does not explain main results.

1.8 Discussion and conclusion

Regulated private markets play a large role in US health insurance coverage. Over 50% of Americans access insurance through markets managed by employers or state governments.(53) Even more receive public benefits through private markets in Medicare, Medicaid Managed Care, and Medicare Part D. These markets are designed so that enrollee choice and insurer competition play a large role in promoting efficient levels of premiums and benefits. Given the prevalence of these "managed competition" markets, understanding how regulation affects competition and benefit generosity is important for consumer welfare.

This study uses a policy change that induced the cancellation of a large number of plans. Analysis of this policy change shows that it reduced competition, leading to decreased plan benefit generosity (increased out-of-pocket costs). Out-of-pocket costs increased by about $130 annually for a representative beneficiary in counties with an average level of cancellation. In counties where HMOs and PPOs stood to gain the least market power, there were negligible changes in benefits, while in counties where HMOs/PPOs stood to gain the most market power, out-of-pocket costs increased by more than $200 annually. Premiums were unchanged for all plans except those directly affected by the policy, potentially due to a quirk in MA regulation that makes them difficult to adjust. Plans may also distort benefits to avoid bad risks, but these changes do not appear to drive observed effects.

Implications of these findings are three-fold. First, results support prior studies finding that lower competition leads to higher cost-sharing and less generous benefits.(14, 19) Given the lack of competition in many health insurance markets, these findings should be a concern.(54) Second, findings show that insurers may modify benefits even if premiums stay unchanged. This suggests that regulators and researchers must consider a range of outcomes when assessing how changes in competition will impact health insurance markets. Lastly, this study indicates that regulators must be cautious when limiting the options available to enrollees. Though plan characteristics are often restricted for sound policy reasons, limiting flexibility in plan design may have unintended consequences.

There are several limitations of this analysis. First, out-of-pocket costs are only one measure of benefit generosity and do not capture some dimensions of plan quality that may be particularly important. For instance, PFFS plans differentiated themselves by offering generous networks. Restricting their networks may have led HMOs or PPOs to reduce network breadth or increase out-of-network cost-sharing. This will not be captured by changes in the out-of-pocket cost measure.[48]

Second, some characteristics of the MA market may limit generalizability. Specifically, features such as PFFS plans and $0-premium plans are idiosyncratic to MA. However, market characteristics such as minimum mandatory benefits, bans on premium discrimination, and risk-adjusted plan subsidies generalize particularly well to markets such as the Affordable Care Act exchanges and Medicare Part D.

Lastly, conclusions about cancellation's effect on overall welfare are limited for several reasons. Though decreased benefit generosity reduced consumer welfare, it may have also reduced unnecessary care if benefits were too high. Estimating this offsetting effect is not possible without data on MA enrollees' consumption.[49] Consumer welfare may also have been reduced by limiting PFFS plans' generous networks, as enrollees likely

[48]Recall that out-of-pocket costs are constructed assuming all care is received in network.

[49]CMS does not make such data available to researchers.

value the option of visiting any provider in Medicare.[50] Calculating these losses also requires information on MA enrollees' consumption of care. Additionally, though both reductions in benefits and reduction in network generosity suggest a negative overall effect on consumer welfare, these decreases may be offset by reductions in government overpayments to plans. Because of this, the change in overall welfare is not only difficult to calculate but also hard to sign.

In summary, results suggest that greater attention must be paid to how benefits respond to changes in competition and market regulation. Moreover, benefit reductions were observed in a highly regulated market where beneficiaries had relatively generous alternatives; this suggests that effects may be stronger in other, less regulated markets, and that further research on how insurers choose plan characteristics is warranted.

[50] Both Ho (2006) and Dunn (2010) show that beneficiaries place significant value on plan network size.(55, 56)

Chapter 2

Dropped out or pushed out? Insurance market exit and provider market power in Medicare Advantage[1]

US health insurance market concentration is high, with two insurers controlling over half the commercial market in 45 states.(54) Health insurance market concentration may be high in part because insurers find it difficult to enter new markets.(22, 57) Many factors potentially limit insurer entry, including state laws and regulations, the existence of scale economies, and the importance of insurer reputation.(22, 41, 58, 59) However, the difficulty of negotiating with providers and forming networks is also frequently cited as an entry deterrent. New entrants face a "chicken-and-egg problem"; they need large enough networks to attract customers, and a large enough volume of customers to add providers to their networks at favorable payment rates.(58) Insurers may find it particularly hard to secure low provider payment rates in markets where hospital- or physician market concentration is high.

[1]Data on physician competition and vertical integration constructed by Hannah Neprash and Michael McWilliams.

Though provider market concentration theoretically restrains entry, the empirical literature on its effects are limited. Due to data constraints, the literature on insurer market participation has focused primarily on the effects of provider supply, rather than provider market structure.[2] Though provider supply and market structure are related, the policy implications differ. If low provider supply suppresses insurer entry, policies should focus on increasing physician supply and hospital beds. If high provider market concentration suppresses entry, this supports expanded anti-trust enforcement.

To test the hypothesis that provider market concentration suppresses insurer market participation, this paper studies insurer exit following a policy experiment in which insurers were forced to form networks *de novo*. Historically, a group of plans in the private Medicare insurance market (Medicare Advantage) were not required to form provider networks. Instead, enrollees in these plans – called "private fee-for-service" or PFFS plans – could visit any provider who accepted Medicare.[3] Instead of negotiating service prices with providers, PFFS plans could pay providers administratively-set Medicare prices for services.[4]

Evidence suggests that insurers capitalized on the lack of network requirement by differentially entering markets where Medicare plan payments were high, relative to costs.(5, 30) In 2008, Congress responded to reports suggesting PFFS insurers were exploiting favorable requirements by passing a law requiring them to form provider networks in most counties.[5] Passage of this legislation led to the cancellation of over two-thirds of this type of plan over the next four years.(61)

This policy change provides an opportunity to study barriers to insurance market

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[2]Exceptions are Ho (2009) and Dranove, et.al. (2003).(41, 58)

[3]91% of doctors were accepting new Medicare patients as of 2012.(60)

[4]These favorable requirements likely reduced insurers' fixed costs of entry and lowered their variable costs, as Medicare prices are believed to be lower than rates paid by commercial insurance plans.

[5]58% of counties, containing 90% of the population, were subject to the network requirement in 2011.

participation by examining insurers' selective exit decisions. Specifically, a subset of insurers complied with the network requirement and continued to offer PFFS plans. However, these insurers also selectively exited 44% of markets where they offered PFFS plans, prior to the policy.

To test which characteristics encouraged insurer exit, provider- and insurer-market structure is compared across plans and markets where insurers were forced to build networks from scratch. Analysis focuses on plans in counties affected by the network requirement where the insurers did not offer an HMO or PPO at baseline. Exit indicators are regressed on a rich set of provider and insurer market variables in logistic models. Based on reports that Medicare Advantage insurers use Traditional Medicare's size and low service prices to gain leverage with providers(62), insurers' bargaining power with providers is measured using the share of all beneficiaries covered by Medicare.

Results suggest that provider market structure and an insurer's "Medicare share" are both important predictors of staying in a market. Moving from the lowest to highest decile of insurer's Medicare share reduces the probability that an insurer exited from 71% to 5%, while moving from the highest to lowest decile of physician HHI and hospital HHI reduces the probability of exit by 18% and 12% respectively. Moreover, interactions between hospital market concentration and insurer market power suggests that insurer market share is most important in the most concentrated hospital markets. In the most concentrated hospital markets, each additional percent point of Medicare share reduced the probability of exit by 4 percentage points. In the most competitive hospital markets, Medicare share had no protective effect.

This paper contributes to two related sets of literature. The first examines the determinants of health insurer market participation(58, 59, 63–68),[6] and the second

[6]Most of these studies examine market participation in Medicare Advantage in earlier eras of the program, before the plan types studied here had a significant presence. The main finding of this literature is that market size (as measured in population) and Medicare payments to plans are significant and robust determinants of entry.

explores how insurer-provider bargaining determines provider prices. Of the studies on insurer market participation, only two focus on provider market structure.[7] Dranove, et. al. (2003) find that greater hospital competition encourages entry among commercial HMOs(41),[8] while Ho (2009) finds that insurers face a circular problem: they must attract enough doctors to their networks to draw enrollees and draw enough enrollees to profitably negotiate with doctors.[9]

This study also relates to the insurer-provider bargaining literature. A large literature suggests that hospital consolidation increases the prices insurers pay for services.[10] More recent literature models prices as the result of a two-sided bargaining problem and finds that insurers exercise "countervailing" pressure on provider prices. These papers find that higher insurance market concentration reduces provider market prices, leads to higher service utilization, or reduces health care workers' wages.(16, 72–75) The only paper that models bargaining and finds that insurer market concentration does not lower provider market prices is Ho and Lee (2014). They find that prices for most hospitals are *lower* in more competitive insurer markets, perhaps because insurers in these markets are less able to pass higher prices on to consumers.(76)

This paper contributes to the literature in several ways. First, it is one of only a few studies to model how provider market structure affects insurer market participation. Additionally, it is the first paper to test how physician market structure and vertical inte-

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[7]Most studies test whether provider supply affects insurer market participation. Even when comparing across studies that use the same measures of supply, results differ widely. Of the five studies that use physicians per capita, two find that greater physician supply fosters entry(59, 66), one finds a negative correlation with entry(63), and one finds no effect.(64, 67) Results using hospital beds per capita as a measure of supply are similarly mixed. One study finds that insurers enter markets with more hospital beds(66), another finds that beds discourage entry(59), and two find no effect.(63, 67)

[8]Competition is measured using "excess hospitals," or the number of hospitals in a market beyond what would be expected given market size.

[9]Ho (2009) models entry among vertically-integrated insurers (i.e., insurers who own provider networks). These insurers may face higher entry barriers, as they simultaneously enter provider and insurance markets, but this circular problem likely affects other insurers' entry as well.(58)

[10]Gaynor and coauthors provide several excellent and recent reviews.(69–71)

gration affect insurer market participation. Second, the use of a natural experiment helps separately identify provider markets' effects on insurer market participation. As insurers in these markets have already incurred many of the costs of entry, the policy change reduces the set of factors that plausibly affect insurer market participation. Lastly, this paper is able to model insurer heterogeneity in market participation decisions. To avoid a multiple equilibrium problem, most previous studies treat insurers as interchangeable or differentiated on very limited dimensions. As this study examines insurers who are already operating in a market, a greater amount of insurer heterogeneity can be modeled.

Section 2.1 describes the policy change, Section 2.2 describes methods and study sample, and Section 2.3 summarizes the data. Section 2.4 describes insurer behavior, and Section 2.5 models differences across counties where insurers selectively exited or kept their plans. Section 2.6 concludes.

2.1 Policy change

Medicare is the government-sponsored program that insures the elderly and disabled. Beneficiaries can enroll in Traditional Medicare, in which the government pays doctors and hospitals directly, or Medicare Advantage. In Medicare Advantage, private insurers are paid a per-enrollee fee for accepting beneficiaries. Insurers are then responsible for paying providers for enrollees' care.

Within Medicare Advantage, insurers set up different types of contractual arrangements with providers. During the time period studied (2007-2012), most beneficiaries enrolled in health maintenance organizations (HMOs) or preferred provider organizations (PPOs). PPOs and HMOs establish provider networks, negotiate payment rates with in-network providers, and define rules on where and how enrollees seek care.[11] Insurers

[11]HMOs generally require patients to obtain referrals to see specialists and do not allow enrollees to seek care out-of-network. PPOs generally do not require referrals and allow enrollees to seek out-of-network care, albeit at higher levels of cost-sharing. PPOs are statutorily forbidden from requiring referral

offering HMOs and PPOs are bound by Medicare's "network adequacy" standards, intended to guarantee that beneficiaries have access to a sufficient number and range of providers.[12]

Prior to July 2008, insurers could offer another plan type – the private fee-for-service (PFFS) plan – without forming a network or contracting with providers. Rather, enrollees in PFFS plans could visit any provider in Medicare, and PFFS insurers paid providers traditional Medicare payment rates for services. The ability to pay providers Medicare rates and offer plans without forming networks may have given PFFS plans a competitive advantage. Not having to form networks may have reduced fixed costs. Being able to pay providers Traditional Medicare rates may have reduced variable costs, as Medicare payment rates are often thought to be below commercial insurers' negotiated rates.[13]

Evidence suggests that PFFS plans differentially entered counties where Medicare payments were high, relative to the costs of providing beneficiaries' care. These differential entry patterns resulted in PFFS plans being paid more per beneficiary than Medicare paid for most beneficiaries in HMO/PPOs or Traditional Medicare.(5, 30) In July 2008, Congress responded to reports that plan were "overpaid" by passing the Medicare Improvements for Patients and Providers Act. This law required PFFS insurers to form provider networks in the majority of counties and removed their ability to pay in-network providers Traditional Medicare rates.(8) The policy applied to all counties where at least two "networked plans" (HMOs or PPOs) already operated. By 2011, when the law became effective, this meant that insurers were forced to build networks and pay

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out-of-network, but are permitted to charge substantially higher copays.

[12]Medicare defines network adequacy standards based on the number of providers in a network and the distance/time that patients must travel to seek care. Standards vary by both specialty type (i.e., primary care vs. cardiac care) and geographic area in which the plan is operating.

[13]Commercially negotiated provider payment rates are thought to be as much of 30% higher than Traditional Medicare rates.(31) Anecdotally, Medicare Advantage payment rates are thought to be somewhere between commercial and Traditional Medicare rates, but no published data on MA payments exist.

negotiated prices in 58% of counties, containing 90% of the US population.[14]

The network requirement and the loss of price advantage, combined with the fact that PFFS plans had fewer ways to manage care, appear to have significantly increased PFFS costs.[15] As documented elsewhere, some insurers responded to the law by canceling their plans *en masse*.(61) Though the network requirement was not effective until 2011, cancellation among PFFS plans at the end of 2009 increased fourfold over the previous year (Figure 2.1).[16] Over the next three years, two-thirds of PFFS plans were cancelled, and PFFS enrollment fell from a little over two million in 2008 to around 500,000 by 2012.

To better understand the role of provider and insurer market power in plan exit decisions, this analysis focuses on insurers who selectively cancelled their PFFS plans.[17] Inspection of the data (Section 2.4) suggests that most PFFS insurers made cancellation decisions on the product level, choosing to cancel all plans rather than build networks. In the most extreme example, one insurer exited more than 1500 counties at once. However, a subset of insurers chose to comply with the network requirement, build networks in selected counties, and continue to offer their PFFS plans. To cleanly identify market-level determinants of exit, analysis focuses on these insurers – the subset who continued to offer PFFS plans in 2012.

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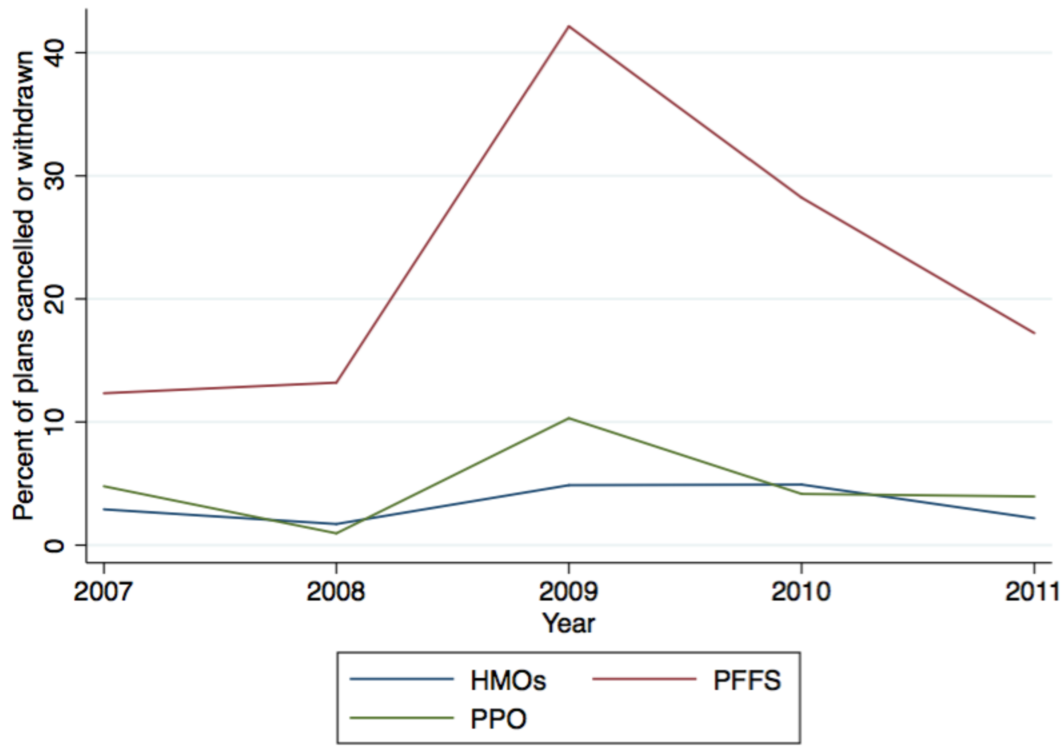
[14]The Center for Medicare and Medicaid Services defined 2011 networked areas based on 2009 contracts. This reduces the likelihood that insurers could manipulate the set of counties considered "networked", at least in the short term.

[15]Like PPOs, PFFS plans could not prohibit patients from seeking out-of-network care, but were allowed to charge higher cost-sharing out-of-network. Unlike PPOs, PFFS plans could not penalize providers for providing more care and were statutorily forbidden from using "gatekeepers" (i.e., primary care doctors from which patients must obtain a referral before seeking specialty care).

[16]Insurers were locked into annual contracts for 2009 at the time the law passed. The network requirement in the Medicare Improvements for Patients and Providers Act was coupled with payment cuts, which may explain the small increase in PPO cancellations in 2009.

[17]Plan service areas are defined on the county level, though an insurer can offer one plan across multiple counties. Insurers form contracts with Medicare and can offer multiple plans within contracts, where benefits can vary across plans. If an insurer decides that a plan is not profitable in a county, it can either cancel the plan, cancel the contract under which the plan is offered, or request a "service area reduction", in which the plan is removed only from specified counties.



Percent of plan-county observations cancelled by year and type. Cancellation includes terminations and service area reductions (when an insurer removes a plan from one county but not another.)

Figure 2.1: Percent of county-plan observations cancelled by year, by type

2.2 Methods and variables

To test which market characteristics encouraged insurers' exit, we compare the set of counties where insurers operated in 2009, before implementation of the policy, to the set of counties where they operated in 2012, the year after implementation.[18] For reasons discussed above, the sample is limited to insurers who continued to offer PFFS plans in 2012.

Analysis focuses on plans in counties affected by the policy and plan-county observations where the sponsoring insurer only offered a PFFS plan at baseline. Counties where an insurer already offered an HMO/PPO were in principle affected by the policy, as

[18] 2009 is treated as baseline, because insurers were locked into 2009 contracts before the policy passed. 2012 is treated as the "post" period, as it is the first year after the network requirement became effective.

insurers had to meet new administrative requirements and pay providers higher prices. However, rather than building a network, insurers could use their existing HMO/PPO network to comply with the law and secure more favorable prices. As a result, insurers who offered PFFS plans exited fewer than 4% of the counties where they already had an HMO/PPO. Focusing on affected counties where insurers only offered PFFS plans at baseline results in a sample of 5,836 plan-county observations, covering 1,592 counties and nine insurers. These 1,592 counties span 87% of the counties where the network requirement held.

Each county-plan observation is coded with a binary indicator for 'exit' if the insurer offered a PFFS plan in that county in 2009 and cancelled all plans in that county by 2012.[19] As will be discussed further in Section 2.4, many insurers replaced their PFFS plans with HMOs/PPOs in some counties. However, the choice of product appears idiosyncratic to the insurer, rather than driven by market characteristics. The focus of this research is insurer market participation, not product choice, both continuing to offer a PFFS plan and swapping a PFFS plan with an HMO/PPO necessitated building a network in this subset of counties. Hence, both decisions are coded as "staying" (exit=0).[20]

Baseline (2009) characteristics of counties and plans are used to explain exit using a logit model. Models are of the form:

$$\begin{aligned} Pr(Exit_{ijm} = 1) = \beta_0 + \gamma M_{jm(2009)} + \theta P_{m(2009)} + \beta C_{ijm(2009)} \\ + \delta[M_{jm(2009)} * P_{m(2009)}] + [\eta_{j2009} + \nu_{m2009}] + \varepsilon_{ijm(2009)} \end{aligned} \tag{2.1}$$

where $Exit_{ijm}$ is an indicator = 1 if insurer j offering plan i cancelled all plans in market m

[19]Consistent with prior studies(59, 66) ,"operating" is defined based on enrollment rather than contracting. A plan is treated as operating in a county if it has more than 11 enrollees, the threshold at which Medicare censors enrollment data.

[20]Replacing a PFFS plan with an HMO/PPO might be advantageous if it helped insurers manage costs through use of capitation or referral. Swapping a HMO/PPO for a PFFS plan may also have been driven by patient health. Insurers were not permitted to automatically reenroll PFFS enrollees into HMOs or PPOs, and inertial enrollees might be difficult to re-enroll. Based on this, the decision to swap plans might be driven by whether insurers wanted to keep their existing enrollees.

by 2012. $M_{jm(2009)}$ is a vector of insurer-market characteristics, $P_{m(2009)}$ captures provider-market characteristics, and $C_{ijm(2009)}$ is a vector of plan- and/or market-specific controls (discussed below). As insurers make county-by-county decisions regarding where to offer plans, the unit of observation is the plan-county level. Variables are fixed at 2009 levels, unless otherwise noted.

The preferred specification includes insurer fixed effects (η_{j2009}) to capture unobserved variation in insurer strategy and state fixed effects (ν_{m2009}) to capture unobserved effects of state regulatory environments.[21] As insurers showed a surprising amount of consensus regarding which counties to exit, standard errors are clustered on the county level.

Before discussing the variables included in (2.1), insurer-provider bargaining in commercial markets and Medicare Advantage is briefly discussed. In the commercial (under-65) market, service prices (p) are often modeled as the equilibrium result of bargaining between insurers and providers.(76–78) In this model, each insurer is competing for enrollees. Enrollees value broad networks, and insurers lose market share if they exclude providers. An insurer will contract with any provider where the profit they earn from including the provider and paying service price p exceeds the profit they earn if they exclude the provider and lose market share.

Providers are also profit maximizing. Their profit increases as a function of patient volume and service prices p and decreases in administrative costs and the costs of providing care. A provider will contract with any insurer for which p exceeds administrative costs and costs of treatment.[22] Each insurer-provider pair agrees on an equilibrium price p that jointly maximizes their profits.

[21] For instance, any-willing-provider laws, or laws that require insurers to contract with any interested provider, may blunt insurers' ability to negotiate prices.(69)

[22] Conversations with industry experts suggest there are administrative costs from contracting with Medicare Advantage insurers. Private insurers do not pay as promptly, are more likely to deny a claim, and demand greater documentation of medical necessity for care than Medicare. In hospitals, Medicare Advantage beneficiaries may be sicker than Traditional Medicare patients admitted for the same condition, because many Medicare Advantage plans require referrals.

Anecdotal reports suggest this bargaining process is slightly different in Medicare Advantage.(62, 79) Specifically, any provider accepting Medicare can take Traditional Medicare patients at Traditional Medicare rates or take Medicare Advantage insurers' patients at some mark-up over Medicare prices ($p = (1 + c)$, where c is the mark-up). The fact that Traditional Medicare is so large and prices are comparatively low helps insurers negotiate lower values of c. Insurer bargaining power may be augmented by the fact that providers cannot refuse emergency services to a Medicare beneficiary, even if the provider is not in the patient's network.

For Medicare Advantage insurers, this implies that their market power may be primarily a function of their size, relative to Medicare. From the provider perspective, if an insurer covers a large portion of the Medicare market, failing to be in that insurer's network is a bigger loss of profit than if the insurer is small, relative to Medicare. Based on the idea that insurers' bargaining power is defined relative to Medicare, individual insurer bargaining power in Medicare Advantage is measured using each insurer's total share of all Medicare beneficiaries ("Medicare share").

In competitive Medicare Advantage markets, other insurers' share of Medicare may also affect bargaining power. If other insurers enroll a large share of the Medicare market, than providers may be able to more credibly threaten to exclude a smaller insurer. This second component of bargaining – the effect of other insurer's Medicare share on an insurer's bargaining power – is captured by the gap between each insurer's share and the share of the largest Medicare Advantage insurer in the market ("Medicare distance").

These two variables are the primary measures of insurer market power used in Equation (2.1). To separately identify their effects, Medicare share is included first, before adding Medicare distance. Though an insurer's Medicare share plausibly reflects bargaining power with providers, it is also closely related to other parameters that affect insurer profit, such as enrollment (insurer scale), Medicare Advantage penetration (which may reflect the appeal of Medicare Advantage), and insurers' market power with consumers (share within Medicare Advantage.) To test whether these variables better explain insurer exit, they are added later as additional controls.

The existence of Medicare may change the bargaining problem and tilt bargaining power in favor of insurers. Still, other, more standard determinants of bargaining power still likely play a role. Equilibrium prices may hinge partly on whether a provider can credibly threaten to exclude an insurer's customers. Provider market power in this data is capturing using two measures of horizontal competition (physician and hospital Hirschmann-Herfindahl Indices (HHI))[23] and a measure of physician hospital vertical integration. HHI in hospital markets is constructed using hospital systems' shares of admissions in a county, while physician HHI is calculated using practices' shares of office, outpatient, and facility spending. The degree of vertical integration is captured using the percent of physicians billing in hospital outpatient departments. These variables do not capture all determinants of market power, but may accurately reflect providers' ability to exclude insurers' patients. (Data are discussed further in Section 2.3 and Appendix B.2.

An insurer's bargaining power may likewise be enhanced by share in other markets. As many Medicare Advantage insurers also have a commercial insurance business (covering individuals under 65), hospitals and doctors may be less willing to bargain aggressively when insurers control large portions of the commercial market. To test this, all models include insurers' share of fully- and self-insured markets, constructed from Interstudy data.

The baseline model starts with the most parsimonious specification:

$$\begin{aligned} Pr(Exit_{ijm} = 1) = \beta_0 + \gamma_1 Medicare\_Share_{jm} + \gamma_2 Commercial\_Share_{jm} + \\ \theta_1 Physician\_HHI_m + \theta_2 Hospital\_HHI_m + \\ \theta_3 Vertical\_Integration_m + \varepsilon_{ijm(2009)} \end{aligned} \quad (2.2)$$

[23]HHI is the sum of each firm's squared market shares. It is widely used because it captures variation in both the number of firms in the market and the distribution of shares across firms.

where $Medicare\_Share_{jm}$ is insurer j's share of all Medicare beneficiaries in market m, $Commercial\_Share_{jm}$ is the insurer's share of commercial enrollees in the same market, and $Physician\_HHI_m$, $Hospital\_HHI_m$ and $Vertical\_Integration_m$ are hospital HHI, physician HHI, and vertical integration in county m.

Theory predicts that insurer market variables should reduce exit ($\gamma_1, \gamma_2 < 0$). If greater provider market power primarily increases provider prices, than these variables will predict exit ($\theta_1, \theta_2 > 0$). However, dealing with larger, more integrated practices may reduce insurer administrative burden or coordination costs, which may attenuate effects.

Subsequent models control for plan- and market-level variables that affect insurer profit. On the market level, provider supply, Medicare payments to plans, Medicare costs, market size, and economic conditions may all affect insurer profitability. Provider supply may affect insurer costs, as insurers require an adequate number of providers with whom to contract. Variation in provider supply is measured using per capita number of hospital beds and physicians in a county. Medicare payments to plans have been found to be a strong predictor of insurer market participation(59, 63–67) and were being reduced over this time period. To control for county-level variation in payments, regressions include both 2009 benchmark levels and 2009-2012 benchmark cuts.

County-level Medicare costs may reflect local variation in practice patterns, which likely affect plan profitability. Specifically, insurers may be less able to manage care in areas where physicians and patients prefer more intensive care. To reflect this, regressions include average, county-level per-patient costs in Traditional Medicare. Costs are standardized by risk scores, so as to reflect variation in the amount of care provided, not population health. As county-level population and economic conditions have also been found to encourage insurer participation, the number of individuals in a county over 65, per capita income, the percent of the population below poverty, and county unemployment are all included as controls.

On the plan level, controls are added for premiums, benefit generosity, plan age, and the health of a plan's enrollees. Insurers' profits are increasing in premiums and decreasing in benefit generosity, so both likely affect exit decisions. Plan age may also affect consumer demand and/or insurer operating costs.[24] Though payments to plans are risk-adjusted, enrollee health may also affect plan profitability. Measures for all these controls are discussed further in the Data section.

Lastly, interactions between provider market variables and the main variable of interest, $Medicare\_Share_{jm}$, are added to test whether insurer bargaining power is more important in markets with greater provider concentration.

Exit decisions among the same set of insurers in counties exempt from the policy are used as placebo tests. Theoretically, provider- and insurer-market variables should be less important in markets where insurers were not compelled to form networks. However, in many ways, these counties make poor controls for networked counties. Most importantly, by definition, their insurer market structure differs substantially, as counties were exempt from the network requirement when they only had one HMO/PPO. Additionally, spatial correlation in exit may drive insurers to leave many non-networked counties.

2.3 Data

Data on plan exit and market characteristics are constructed from several sources. Data on Medicare Advantage plans come from unrestricted datasets published by Medicare. Provider market variables are constructed using data from the American Hospital Association, the Medicare Carrier File, and Medicare outpatient claims. These data are supplemented by information on commercial insurance markets from Health Leaders Interstudy and control variables from the Census Bureau, the Area Resource File, the

[24]Plan age affects demand in that more established plans may have better reputations with consumers.(56, 67) In markets with significant consumer inertia, older plans may also be larger. On the cost side, less costly/better managed plans may survive longer, so plan age may indicated greater plan efficiency.

Bureau of Labor Statistics.

Medicare Advantage data capture insurer market participation and bargaining power. Center for Medicare and Medicaid Services (CMS) data on plan-county level enrollment identify markets where insurers offered PFFS plans prior to the network requirement and markets they exited after its implementation.

As discussed in Section 2.2, Medicare Advantage insurers' bargaining power with providers may hinge on two variables – Medicare share and Medicare distance. The first variable is constructed using the share of Medicare eligibles in a county enrolled in an insurer's Medicare Advantage plans (HMOs, PPOs, and PFFS).[25] The second is measured using the difference between insurer j's Medicare share and the Medicare share of the largest insurer in the county. For the largest insurer, this number is 0.

Insurers' commercial market share is constructed using HealthLeaders Interstudy data. Commercial market shares reflect an insurer's share of all lives covered in fully- and self-insured policies in a county.[26] Commercial market share data is matched to Medicare data based on insurer name and county. Details of the matching process are described in the Data Appendix.

Provider market concentration measures are constructed using Medicare claims and American Hospital Association data. (Means and standard deviations for these variables are summarized in Table 2.1.)

Competition within both physician and hospital markets (horizontal market structure) is captured using HHI. Physician HHI is calculated using a practice's share of office, outpatient, and facility spending in Medicare claims in a county, where practices are identified using tax identification numbers. Using tax identification number to identify

[25]Special needs plans, cost plans, and demonstration plans are excluded from this calculation, because enrollees in these plans likely only interact with a specialized subset of providers.

[26]Medicaid managed care and Federal Employee Health Benefit Program (FEHBP) enrollment is excluded from calculations of shares. The Medicaid market is excluded because Medicaid payment rates to providers are so low that providers are unlikely to treat Medicaid patients as substitutes for Medicare patients. FEHBP plans are excluded due to data issues.

practices likely underestimates physician concentration, as physicians can bill under multiple tax IDs. However, this method of identifying physician practices has to be consistent with alternate methods.(80)

Hospital HHI is measured on the hospital system level, using a system's shares of total admissions.[27] Shares exclude federal, psychiatric, long-term, and children's hospitals. Though insurers make county-by-county decisions about where to operate, hospitals likely operate across broader geographic areas. For this reason, hospital HHI is calculated on the hospital service area (HSA) level. HSAs are matched to counties, and analysis is performed on the county-plan level.

An additional dimension of provider market power, vertical integration, is measured using the percent of physician Medicare charges that are billed in a hospital outpatient department. This measure plausibly captures vertical integration, as physicians are only allowed to bill under hospital outpatient codes when they are employed by a hospital or when their practice is owned by that hospital. Due to data limitations, this variable is constructed on the core based statistical area (CBSA) level and matched to counties. Further details for all these variables are provided in the data appendix.

For all three variables, there is insufficient data to calculate HHI for some observations. For physician HHI and vertical integration, this occurs where there are fewer than 10 claims in a market. For hospital HHI, this occurs when there are no non-federal and non-specialty hospital beds in an HSA. For all variables, an indicator is constructed to reflect that the data are missing and values of the variables are set to 0. Analysis that varies treatment of missing data is presented in the Technical Appendix.

As discussed in Section 2.2, a range of plan- and market- characteristics that affect plan profitability are included as controls. On the plan level, data include plan premiums, benefit generosity, plan age, and risk scores, calculated using publicly-available Medicare data and Medicare out-of-pocket cost files.

[27] In the Appendix, an alternate measure is constructed using hospital beds.

Plan premiums reflect the additional premium a plan charges beyond the standard Medicare Part B premium. Premiums are also adjusted to reflect any amount by which the plan reduces the standard Part B premium.[28] Plan age is captured using the age of the contract under which a plan is offered.

Benefit generosity is measured using expected out-of-pocket cost, a standardized measure of generosity constructed by Medicare for beneficiaries' use in choosing plans. It is calculated using spending data for a representative cohort of enrollees and reflects expected out-of-pocket spending for the average beneficiary based on each plan's copays, deductibles, and covered benefits. It captures spending for a variety of services, including inpatient and outpatient hospital services, primary care and specialist services, lab tests, diagnostics, durable medical equipment, and prescription drugs.

Enrollee health is controlled for using average, plan-level CMS Hierarchical Condition Category (HCC) risk scores. Risk scores capture expected spending based on demographic and diagnostic categories and are used to compensate plans for enrolling sicker beneficiaries.(37) In theory, risk scores should not affect profitability, as plans are compensated based on them. However, recent evidence suggests that plans are under-compensated for higher levels of risk,(45, 46) and univariate regression of plan risk on exit suggests that plans with higher risk are indeed more likely to exit.[29]

On the market level, controls are added to reflect market size, county economic conditions, Medicare benchmark payment levels, future payment cuts, the supply of healthcare providers, healthcare spending for Traditional Medicare beneficiaries in a county, the health of enrollees in the Medicare market, and Medicare Advantage penetration. Market size is captured using the number of enrollees in a county who are over age 65 (from the Area Resource file), while economic factors are captured using

[28] Insurers reduced Part B premiums for only 3% of plans in 2009.

[29] A standard deviation increase in a plan's risk score increases the probability of exit by 3 percentage points, significant at the $p < .05$ level.

county-level per-capita income, the percent of the county population below poverty (from the Census Small Area Income Poverty Estimates), and the unemployment rate (from the Bureau of Labor Statistics).

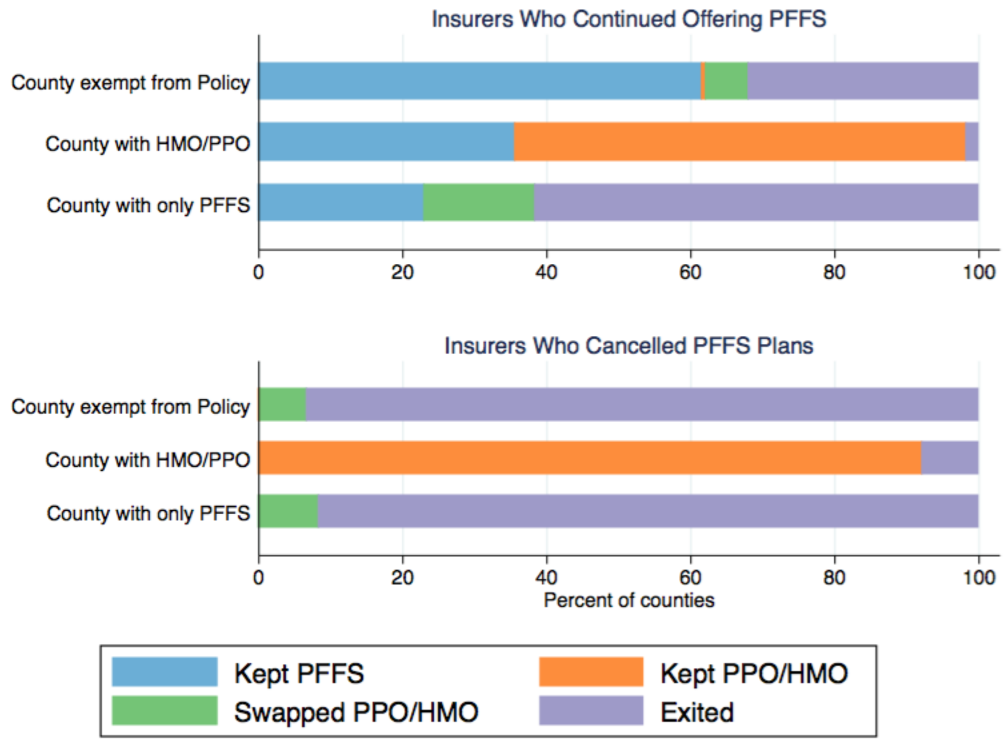
Though provider market structure and supply are closely related, provider supply may capture additional information about input costs. Variation in supply is captured by the number of hospital beds and the number of doctors per 10,000 people from the American Hospital Association data and the Area Resource file. Spending in Traditional Medicare is captured using risk-standardized average cost for beneficiaries in Medicare Parts A and B in a county.[30] Average, county-level risk scores among Traditional Medicare beneficiaries and all Medicare Advantage plans in a county are added to reflect the health of insurers' potential enrollees.

Values for all variables are fixed at baseline levels, so they are unrelated to the level of exit in a county. However, many variables, particularly premiums and benefits, may still be endogenously related to insurer and provider market structure variables. For instance, 2009 premiums may be higher in concentrated hospital markets because insurers pass on higher provider prices to enrollees. To avoid biasing coefficients by including endogenous covariates, main results are presented with and without controls.

2.4 Insurer strategies

Figures 2.2-2.4 describe insurer behavior and characteristics. Figure 2.2 examines insurers' actions divided across three groups of counties: 1) counties exempt from the policy change, 2) affected counties where the insurer offered an HMO/PPO at baseline, and 3) affected counties where the insurer only offered a PFFS plan. The top panel shows the distribution of actions taken by insurers who continued to offer PFFS plans, while the

[30]Traditional Medicare costs are standardized by dividing by the average risk score of beneficiaries in Parts A and B. In theory, standardized costs reflect only variation in spending, rather than population health.



Distribution of insurer actions (keep PFFS, swap HMO/PPO for PFFS, keep HMO/PPO, or exit entirely) across types of counties (exempt from the policy, where insurer already had an HMO/PPO, and where they only had a PFFS plan.) Top panel includes all insurers who continued to offer PFFS plans in 2012, and bottom panel includes all insurers who cancelled all PFFS plans.

Figure 2.2: Distribution of insurer actions across types of counties

bottom panel shows the distribution for insurers who cancelled them.

Behavior differed between insurers who continued to offer PFFS and those who did not, particularly in counties exempt from the policy. Insurers who cancelled all PFFS plans exited 93% of counties exempted from the policy, while insurers who continued to offer PFFS exited only 32%. As these counties were theoretically unaffected by changes in network costs, this suggests that most insurers' decisions were driven by product-level rather than county-level concerns and supports the choice of focusing on insurers who continued to offer PFFS plans.

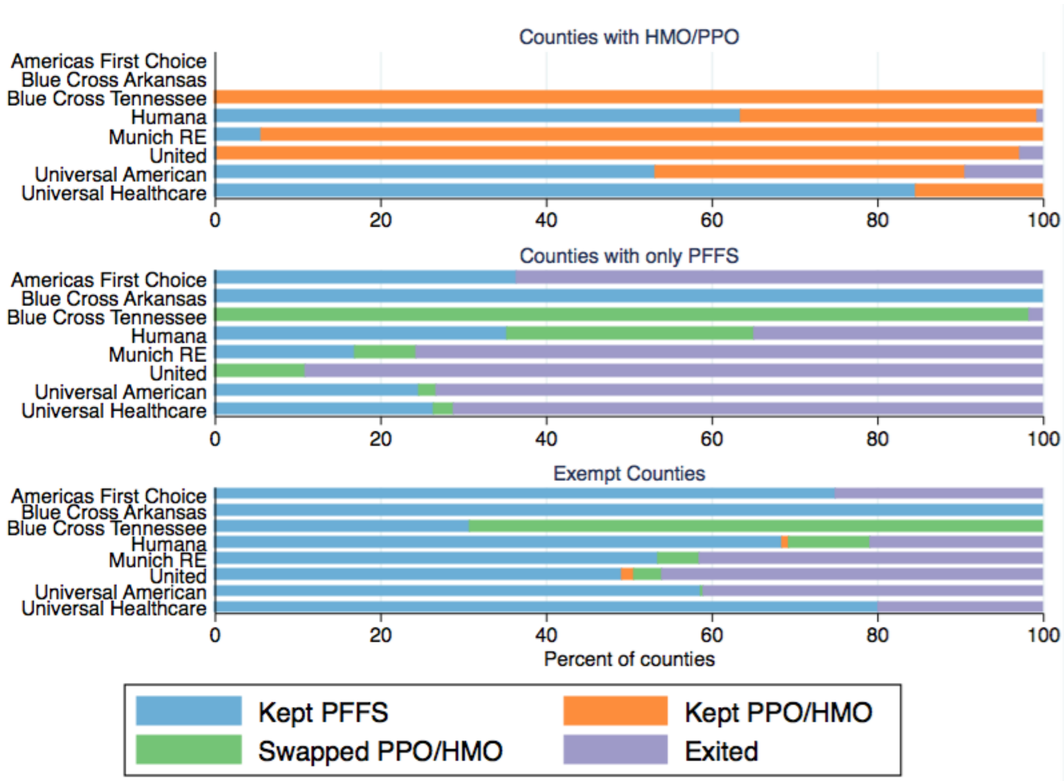
Figure 2.2 also shows that exit was extremely rare when an insurer offered an HMO/PPO in a county at baseline. Insurers who continued to offer PFFS plans exited only 2% of these counties. This is likely due to the fact that insurers rarely cancel HMO/ PPOs. However, insurers were also less likely to cancel PFFS plans in these counties than in counties where they only offered a PFFS plan (35 vs. 23% in the top panel). This suggests that having an HMO/PPO in a county substantially lessened the costs of complying with the policy and that the focus of analysis should be on county-plan observations where an insurer had no HMO/PPO at baseline.

Insurers who continued to offer private-fee-for service plans pursued a variety of strategies (Figure 2.3). Some insurers clearly continued to focus on their PFFS business. For instance, Blue Cross Blue Shield of Arkansas[31] continued to offer PFFS plans in every county where they had offered them in 2009. Other insurers (i.e., America's First Choice) cancelled in most counties, but still continued to only offer PFFS plans. Other insurers generally replaced PFFS plans with HMOs/PPOs. The most dramatic example is Blue Cross Blue Shield of Tennessee, which removed its PFFS plans from 61 counties affected by the policy and replaced them with a PPO in all but one. As product choice seems to vary more on the insurer than county level, "exit" is defined as the decision to cancel all plans in a county, not just PFFS plans.

The top panel of Figure 2.4 shows actions for insurers who continued to offer PFFS in the subset of counties where they had no HMO/PPO. On average, insurers exited 62% of these counties, but exit rates varied across insurers. Rates ranged from 0% for Blue Cross Blue Shield of Arkansas to 100% for Medica of Minnesota. There was also significant heterogeneity among these insurer's characteristics. For instance, the bottom panel of Figure 2.4 shows the distribution of commercial market presence across insurers. With the exception of Humana, insurers had commercial plans in all or none of their counties. This suggests that controlling for insurer identity is important for analysis.

Table 2.1 summarizes characteristics for plan-county observations in the sample of interest (PFFS plans offered by insurers in counties where they had no HMO/PPO at

[31]Offering plans under the name USAble Mutual Insurance Co.

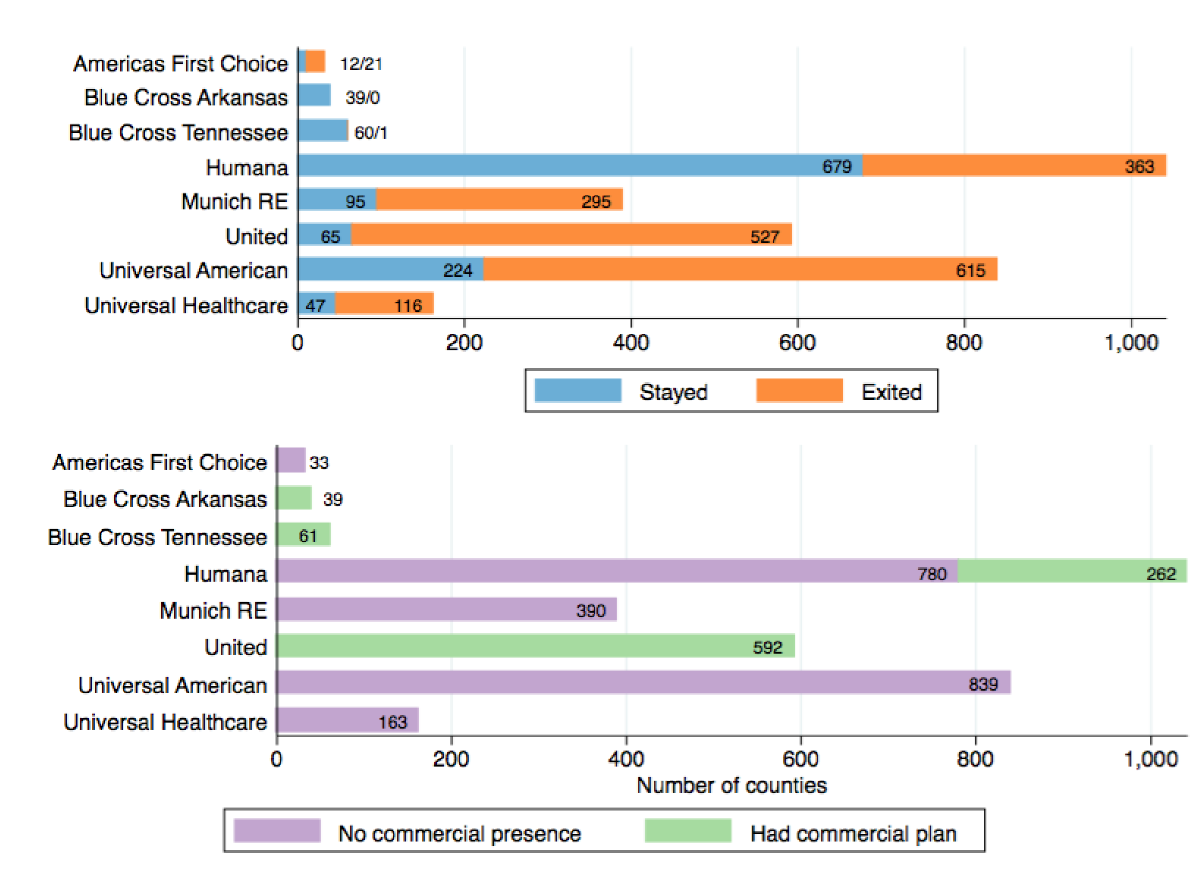


Percent of each type of strategy pursued by insurers who continued to offer PFFS. Counties are divided into those unaffected by the network requirement, counties where an insurer already had an HMO/PPO, and counties where they only had a PFFS plan. Not shown is Medica of Minnesota, which cancelled all PFFS plans it offered in 2009 and entered new counties exempt from the policy. America's First Choice and Blue Cross Blue Shield of Arkansas are not shown in the top panel, as they did not offer HMO/PPOs.

Figure 2.3: Actions taken by insurers continuing to offer PFFS plans

baseline). Differences in insurance markets can easily be observed. Insurers had larger Medicare shares in the counties where they kept their plans (3% vs. 1%). Medicare distance also shows that insurers stayed in counties where they were larger. Their shares were on average 4 percentage points smaller than the largest insurer in counties where they kept plans, vs. 5 percentage points smaller than the largest insurer in counties where they exited. In the commercial market, insurers were less likely to have a commercial plan, but had a larger share of the market when they did. Insurers also kept operating in counties where the Medicare Advantage markets were bigger, as measured by penetration, though differences were small.

Counties where insurers left their plans had a greater supply of providers than



Top panel shows percent of counties in the sample exited by insurers who continued to offer PFFS plans. Bottom panel shows the percent of counties in the sample where the insurer had a commercial market plan. The sample of counties includes counties affected by the policy where the insurer only offered PFFS plans at baseline. Blue Cross Blue Shield of Arkansas is sold under the trade name USAble Mutual Insurance in Medicare.

Figure 2.4: Insurer actions and characteristics in sample counties

counties where they cancelled plans. Counties where insurers left plans had an extra .13 doctors and .31 hospital beds per 1000 people. Though these differences are small, they are statistically different from each other at the $p < .05$ level.

Differences in provider markets are less clear. Physician HHI was higher by more than 100 points in counties where insurers kept plans, while 1% more doctors and hospitals were more vertically integrated in counties where insurers kept plans Hospital HHI was about 50 points lower in markets insurers stayed in (implying more competition), but differences are not statistically significant.

Physician market data is missing for 2-3% of counties, hospital market data is missing for 10-11%, and the vertical integration measure is missing for almost of a quarter of counties. The amount of data missing also differs across markets, but differences in the amount of missing data are statistically significant only for vertical integration. Moreover, regressions run with and without missing data look qualitatively similar (see Table B.2).

Means and standard deviations for other control variables are discussed in Appendix B.1 and listed in Table B.1.

Table 2.1: Plan-county characteristics

	Stayers	Leavers
Insurer Market Characteristics		
Insurer Medicare Share (%)	0.03	0.01
	(0.03)	(0.02)
Medicare distance (%)	0.04	0.05
	(0.05)	(0.05)
Insurer Share of MA (%)	0.25	0.15
	(0.24)	(0.20)
Has Commercial Plan in market	0.24	0.32
	(0.42)	(0.47)
Commerical market share (%)	0.05	0.03
	(0.13)	(0.07)
MA penetration (%)	0.21	0.20
	(0.10)	(0.11)
Provider Market Characteristics		
Doctors per 1000	1.50	1.37
	(1.43)	(1.28)
County-level hospital beds per 1000	4.60	4.29
	(4.77)	(4.89)
Hospital HHI (admissions)	8364	8410
	(2456)	(2366)
Physician HHI	2544	2415
	(2524)	(2545)
Percent of docs vertically integrated	0.24	0.23
	(0.13)	(0.12)
Missing vertical integration	0.25	0.20
	(0.44)	(0.40)
Missing hospital HHI	0.11	0.10
	(0.31)	(0.31)
Missing physician HHI	0.02	0.03
	(0.15)	(0.16)
Missing Data		
Missing vertical integration	0.25	0.20
Missing hospital HHI	0.11	0.10
Missing physician HHI	0.02	0.03
Observations	2482	3354

Characteristics for plan-county observations where insurer exited (leavers) and did not exit (stayers).

2.5 Results

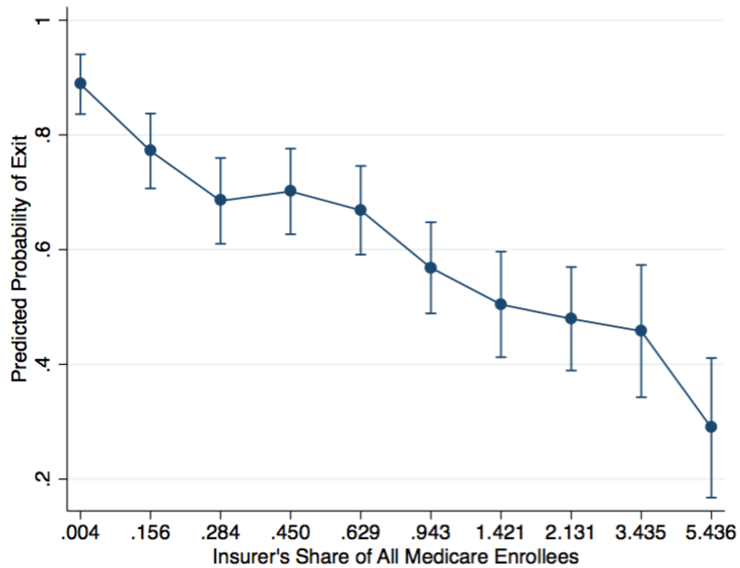
Table 2.2 shows the result of regressing exit indicators on county and plan characteristics. Column 1 shows results from regressing exit on provider market structure and insurer bargaining power, without controls or fixed effects.[32] Insurers who enrolled a greater share of Medicare were consistently less likely to exit. Each additional percentage point of Medicare share reduced the log odds of exit by $-.33$. This magnitude suggests that moving from the lowest decile of Medicare share ($< 1\%$) to the top decile ($> 10\%$) would reduce an insurer's probability of exit more than 14-fold – from 71% to approximately 5%. This effect is not driven by solely by insurers with large shares. Regressing exit on deciles of Medicare share reveals that a larger share of Medicare has protective, albeit diminishing, effects throughout the distribution (Figure 2.5).

Higher commercial market share also protects insurers from exit. However, all insurers except Humana offered a commercial plan in either all or none of their markets (see Figure 2.2). As some insurers were more likely to exit than others, this may create bias in this coefficient. For this reason, this coefficient is revisiting after adding insurer fixed effects.

Higher physician and hospital market HHI both increase likelihood of exit. Each additional thousand points of hospital HHI increased the probability of exit by a percentage point, while each additional thousand points of physician HHI increased the probability of exit by 2 percentage points. The percent of doctors in a county who were vertically integrated had no significant effect on exit.

Column 2 adds a full set of controls plus insurer and state fixed effects to the regression from Column 1. Fixed effects significantly improve the model's fit in terms of log-likelihood and pseudo r-squared, but impose a small cost on sample size (n=165 or 2.8% of the original sample). Insurer fixed effects result in the exclusion of one insurer

[32] All regressions include controls for missing provider market data.



Results of regressing insurer exit on deciles of insurers' Medicare share, controlling for county characteristics, plan characteristics, and provider market characteristics. Logistic regression with standard errors clustered on the county level; 95% confidence intervals.

Figure 2.5: Effect of insurers' Medicare share on predicted probability of exit

who exits all counties (Medica)[33] and another who stays in all counties (Arkansas Blue Cross Blue Shield). State fixed effects result in the exclusion of five states that insurers exited altogether.[34]

Coefficients on insurer-and provider-market structure variables are qualitatively similar to those in unadjusted regressions. The effect of an insurer's Medicare share is slightly smaller (log odds=-.26), but is still highly protective, even after controlling for enrollment (as a measure of scale), plan's share within Medicare Advantage (as a measure of within-Medicare Advantage market power), and Medicare Advantage penetration (reflecting the overall appeal of managed care.)[35] Average marginal effects suggest that each additional percentage point of Medicare share reduces the probability of insurer

[33]Medica exits all counties where it operated in 2009 and enters an entirely new set by 2012.

[34]New Jersey, Maryland, Massachusetts, DC, and Delaware.

[35]These variables are naturally quite correlated with an insurer's share of all Medicare. $\rho = .33$ for the correlation between enrollment and insurer's share of Medicare, while $\rho = .025$ for the correlation with penetration.

exit by 4 percentage points.

Controlling for insurer identity triples the protective effect of commercial market share. Average marginal effects suggest that each each additional percentage point of commercial share reduces the probability of exit by 1 percentage point. Moving from $0 - 10\%$ commercial market share to the highest decile of market share ($70 - 80\%$ of the market) reduces the predicted probability of exit from 59% to a little under 5%.

Physician HHI's effects are similar, even after controlling for the supply of doctors in the county. Hospital HHI's effect is somewhat stronger (θ_1=.10) than in unadjusted regressions, even controlling for the number of hospital beds in the county. The slightly larger coefficient suggests that moving from the most competitive HSAs to one with only one hospital system changes the predicted probability of exit from .46 to .61. Vertical integration remains insignificant.

Most other controls are of intuitive signs and reasonable magnitudes. Insurers are less likely to leave counties with higher benchmark payments ($\beta_{bench} = -.50$). Insurers do not respond significantly to future benchmark cuts, although coefficients are in the expected direction.[36] Insurers were also more likely to leave counties with higher traditional Medicare costs. Marginal effects suggest that an additional hundred dollars of Medicare spending in a county almost exactly offset an additional hundred dollars of benchmark payments. Insurers were also less likely to cancel larger and older plans, as measured by number of enrollees and contract age. Each additional 100 enrollees reduced the log odds of exit by -.18, and each additional year of contract age reduced log odds of exit by -.46. These findings suggest an important role for insurer experience and economies of scale.

The coefficient on Traditional Medicare risk scores is large and significant ($\beta = -.35$), but in an unexpected direction. This may suggest that Traditional Medicare risk scores

[36] Payment cuts and benchmark payments are strongly correlated ($\rho = .45$), with greater payment cut being implemented in counties where benchmarks were already high.

reflect selection between Medicare Advantage and Traditional Medicare, and that insurers are less likely to leave favorably selected markets. Alternately, risk scores may reflect the diligence with which local providers code diagnoses. As insurers are compensated based on risk scores, providers "up-coding" may make plans substantially more profitable.(43) Coefficients on plan risk and county-level Medicare Advantage risk are insignificant, possibly because they are highly correlated with Traditional Medicare risk ($\rho = .29$ and $\rho = .64$ respectively).

Not shown are controls for plan generosity (expected out-of-pocket cost), plan premiums, county per-capita income, the percent of the county in poverty, the percent of the county that is unemployed, the size of the population over 65, an indicator for whether a plan is urban or rural, and indicators for whether or not there is missing data in provider market variables. Of all these variables, only the percent of the county below poverty and the indicator for missing physician HHI were significant at the $p < .05$ level. Insurers were more likely to leave counties with more poverty, but the effect size is very small – each additional percentage point of poverty increased the probability of exit by .00008. Insurers were much more likely to leave counties where there was insufficient data to calculate physician HHI (β_{mi}=1.10), suggesting that provider supply is quite important.

As discussed in earlier sections, an insurer's bargaining power with providers likely depends on both their size relative to all Medicare and their size within Medicare Advantage. If an insurer is small within the MA market, they may be less able to extract favorable price terms from a provider. Medicare distance is added to test this (Column 3). Coefficients suggest both Medicare distance and Medicare share are important, ($\gamma_{TM\_dist} = .07$ and $\gamma_{TM\_Share} = -.21$, significant at the $p < .05$ level), but the size of coefficients indicate Medicare share is more so.[37]

To test how much of county-level variation in exit is captured by this model, a model

[37] We acknowledge that both variables may plausibly reflect insurer's market power over consumers, rather than consumer's bargaining power with providers. We return to discussing this issue in later sections.

with county-fixed effects is tested (Col 4.) County fixed effects necessitate omission of all variables defined on the county level (physician HHI, hospital HHI). This also results in the exclusion of 47% of observations, as insurers showed substantial consensus about which markets were profitable.[38] County-fixed effects make Medicare Share and commercial market share much more protective than in previous regressions ($\gamma_{TM\_share} = -1.12$ and $\gamma_{comm\_Share} = -.14$).

Columns 5 and 6 test whether similar patterns of exit are observed in counties exempt from the policy. To make the sample comparable, attention is limited to the same insurers, and county-plan observations where the insurer offered an HMO/PPO at baseline are excluded (n=46). Both the sample size (n=2,931) and the proportion of plans that exited (n=921, 31%) are smaller. The sample spans 994 counties, which includes all exempt counties where insurers offered PFFS plans in 2009.

Column 5 repeats the specification from Column 2, excluding controls for within-Medicare market share and Medicare distance. Inclusion of state- and insurer- fixed effects result in the omission of two states where all insurers exited (Delaware and Massachusetts) and two insurers who stayed in all counties (Blue Cross Blue Shield of Tennessee and Blue Cross Blue Shield of Arkansas).

Coefficients on Medicare market share and commercial market share are still significantly protective in non-networked counties, but effect sizes are half as large ($\gamma_{TM\_Share} = -.13$ and $\gamma_{Comm\_Share} = -.01$. Moreover, inclusion of insurer's Medicare Advantage share and Medicare distance (Column 6) makes both of these variables insignificant.[39]

Coefficients on hospital HHI are insignificant in exempt counties, suggesting that provider bargaining power is less important. However, physician HHI remains predictive

[38] All insurers pulled their plans from 40% of these counties (n=630), and all stayed in another 22% (n=346).

[39] An insurer's share within MA is again highly correlated with their Medicare share $\rho = .47$. However, their share is less correlated in these markets than in markets for baseline regressions ($\rho = .62$), where insurers' Medicare Advantage share did not affect the size of the coefficient on Medicare share.

of exit, and effect sizes are similar to those in affected counties. This may suggest that physician HHI is a better measure of provider supply than market concentration. Alternately, physician market structure may primarily affect providers' decision to take patients, rather than prices.[40]

Coefficients on most controls are of similar signs and magnitudes to those in counties affected by the policy. One difference between populations is that county-level Medicare Advantage risk scores significantly predictive of exit in exempt counties. This may suggest that beneficiary health is more important to insurers, as they cannot as easily manage care without a network.

Medicare share is a plausible measure of insurer bargaining power, but may also reflect insurers' ability to increase mark-ups to consumers. If an insurers' share of Medicare primarily reflects bargaining power with providers, then a larger Medicare share should be more important in more concentrated provider markets. This is tested by adding interactions between insurer's Medicare share and provider market variables (Table 2.3 and Figure 2.6).

In networked counties, insurers' Medicare share interacts significantly and negatively with hospital HHI (Column 1 and Figure 2.6), suggesting that Medicare share is more protective in more concentrated hospital markets. Margins on this interaction suggest that, for each additional 1000 points in hospital HHI, an additional percentage point of Medicare share would reduce the probability of exit by .4 percentage points. In the most concentrated hospital markets, moving from the highest decile to lowest decile of Medicare share reduces the probability of exit by 26 percentage points. In the most competitive hospital markets in the sample,[41] Medicare share has no protective effect.

The interaction between physician HHI and Medicare bargaining share is insignifi-

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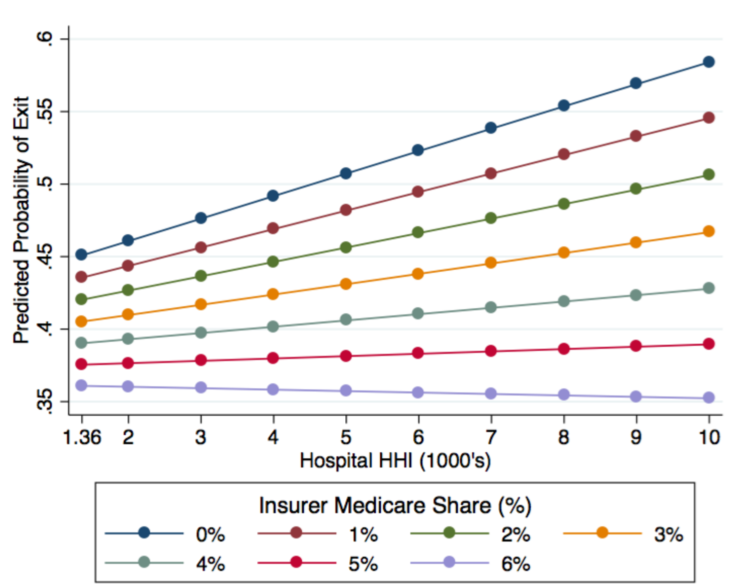
[40]Anecdotal reports suggests that non-networked PFFS plans sometime struggled with physicians turning away patients.(81) This problem may be worse in areas where providers are capacity-constrained.

[41]Minimum HHI=1364.



Predicted probabilities from regressing insurer exit on a continuous interaction between insurers' Medicare share and HSA-level hospital HHI, controlling for county characteristics, plan characteristics, and provider market characteristics. Logistic regression with standard errors clustered on the county level.

Figure 2.6: Interaction between Medicare share and hospital HHI

cant, although physician HHI still predicts exit. This may indicate that the bargaining process between insurers and physicians differs from insurer-hospital bargaining, or that physician market power more closely reflects physician supply.

The next regression (Column 2) adds Medicare distance and interactions between this variable and provider market variables. Medicare distance still significantly predicts exit, but interactions between Medicare distance and provider HHIs are almost identically 0.

The same interactions are tested in non-networked counties (Column 3). None of the interaction coefficients are significant and adding interactions renders the effects on Medicare Share and physician HHI insignificant.

To test how well models predict exit, count statistics and pseudo adjusted R^2 are presented across different models. Count statistics reflect the percent of the time that the model correctly predicts exit.[42] Pseudo R^2, a commonly reported statistic of fit in

[42]The cutoff for success here is .5. There are many flaws of count statistics (reviewed in Train (2007) and Wooldridge (2010)(82, 83)). The most obvious is that they are sensitive to the choice of probability cutoff.

logistic regression models, is the ratio of the log likelihood of the adjusted model to the log-likelihood of the null. Higher numbers indicate better fit, but do not have the same ready interpretation as OLS R^2.[43]

Table 2.4 compares fit across models with only provider and insurer market variables (Column 1)[44], a model that excludes these variables but includes all other controls (Column 2), a model with all controls, insurer-provider market variables, and state fixed effects ("Full Model", Column 3), and a model with county fixed effects plus all plan level variables (Column 4). The last comparison is included to test how much residual county-level variation is unexplained in the full model. Fit statistics are intended for comparison within the same sample, and county fixed effects exclude a substantial portion of the sample. To address this, fit is compared for the whole sample and the subset of plan-county observations included in county fixed effects regressions.

Count statistics (Rows 1 and 3) show that provider and insurer variables by themselves explain more variation than other all other controls, including county-level benchmarks and county-level costs. The full model (Col 4) performs far better than either set of controls (79% correctly predicted vs. 67 and 62%). Without accounting for the fact that models are estimated on different samples, the full model does slightly worse than the county fixed effects model (79% vs. 84%.) However, when estimated on the same sample, these models do almost exactly as well. This suggests that the full model captures a majority of the county-level variation that predicts insurer exit.

However, .5 is thought to be a reasonable cutoff when neither outcome is particularly rare (as in this case.)

[43]($R^2 = 1 - \frac{LL_m}{LL_0}$ where LL_m is the log likelihood including controls and LL_0 is the log-likelihood using only an intercept.) Higher numbers indicate better fit, but Pseudo R^2 do not have the same ready interpretation as OLS R^2. Values of pseudo R^2 are also generally lower than OLS $R^2$and an often cited rule of thumb is that numbers between .2-.4 indicate a "good fit".

[44]Medicare share, Medicare distance, commercial market share, hospital and physician HHI, the percent vertically integrated, and indicators for missing values.

Table 2.2: Predictors of exit for insurers staying in PFFS (1= Exit)

VARIABLES	(1) Baseline No Fixed Effects	(2) Controls Fixed Effects	(3) Insurance Structure	(4) County Fixed Effects	(5) Non- Networked	(6) Non- Networked
Insurer Medicare Share (%)	-0.33**	-0.26**	-0.21**	-1.12**	-0.13**	-0.05
	(0.03)	(0.05)	(0.05)	(0.08)	(0.06)	(0.07)
Commerical market share (%)	-0.02**	-0.06**	-0.07**	-0.14**	-0.01**	-0.01
	(0.00)	(0.02)	(0.02)	(0.02)	(0.01)	(0.01)
Insurer Share of MA (%)		0.01	0.01			-0.01**
		(0.00)	(0.00)			(0.00)
MA penetration (%)		-0.00	-0.03**		-0.02	-0.04
		(0.01)	(0.01)		(0.02)	(0.02)
Medicare distance (%)			0.07**			-0.01
			(0.02)			(0.05)
County-level physician HHI (1000's)	0.08**	0.10**	0.10**		0.07**	0.08**
	(0.02)	(0.03)	(0.03)		(0.03)	(0.03)
HSA-level admission HHI (1000's)	0.07**	0.10**	0.10**		-0.08	-0.08
	(0.02)	(0.03)	(0.03)		(0.05)	(0.05)
Percent of docs vertically integrated	-0.61	0.03	0.07		-0.72	-0.74
	(0.43)	(0.54)	(0.54)		(0.78)	(0.78)
Number of enrollees in plan (100's)		-0.18**	-0.17**	-0.15**	-0.11	-0.09
		(0.03)	(0.03)	(0.07)	(0.09)	(0.09)
Medicare benchmark (100's)		-0.49**	-0.46**		-0.49**	-0.53**
		(0.16)	(0.16)		(0.21)	(0.21)
Future Benchmark cuts (+100's)		0.82	0.72		0.80	0.92
		(0.46)	(0.46)		(0.50)	(0.50)
Contract age (years)		-0.46**	-0.46**		0.43	0.48
		(0.18)	(0.18)		(0.30)	(0.29)
Normalized FFS costs (100's)		0.45**	0.39**		0.55**	0.62**
		(0.13)	(0.13)		(0.16)	(0.16)
Standardized FFS risk		-0.36**	-0.34**		-0.37**	-0.39**
		(0.12)	(0.12)		(0.13)	(0.13)
Standardized MA county risk score		0.13	0.09		0.20**	0.22**
		(0.10)	(0.10)		(0.08)	(0.08)
Standardized plan-level risk		0.05	0.06		-0.02	-0.01
		(0.06)	(0.06)		(0.09)	(0.09)
Doctors per 10,000		-0.01**	-0.01**		-0.02**	-0.02**
		(0.00)	(0.00)		(0.01)	(0.01)
County-level Hospital beds per 10,000		-0.00	-0.00		-0.00	-0.00
		(0.00)	(0.00)		(0.00)	(0.00)
Observations	5,836	5,671	5,671	3,059	2,884	2,884
Insurer FE	No	Yes	Yes	Yes	Yes	Yes
State FE	No	Yes	Yes	Yes	Yes	Yes
County FE	No	No	No	Yes	No	No
Log Likelihood	-3657	-2625	-2610	-925.1	-1504	-1495
Pseudo R^2	0.0811	0.322	0.325	0.564	0.164	0.169

Logit model, standard errors clustered on the county level. Exit=1 if insurers will removed all plans from a market between 2010-2012 and 0 if not. All variables are 2009 values. Sample includes PFFS plans offered by insurers who continued to offer PFFS plans in 2012. Cols 1-4 are PFFS plans offered in counties where insurer offered no HMO/PPO. Cols 5-6 are PFFS plans offered by the same insurers in counties unaffected by policy. Control variables not shown: the number of people in the county over 65, percent of population below poverty, the unemployment rate, per-capita income plan benefits, premiums, plan age, and indicators for missing physician HHI, hospital HHI, or vertical integration data. Additional stats include log-likelihood and McFadden's R^2. $^{**}p \leq .05$.

Table 2.3: Predictors of exit with continuous interactions (1= Exit)

VARIABLES	(1) Networked Interactions	(2) Networked With Distance	(3) Non-Networked Without Share
Insurer Medicare Share (%)	-0.05	-0.05	-0.09
	(0.08)	(0.09)	(0.09)
HSA-level admission HHI (1000's)	0.13**	0.12**	-0.07
	(0.03)	(0.04)	(0.05)
County-level physician HHI (1000's)	0.09**	0.10**	0.07
	(0.04)	(0.04)	(0.04)
Medicare Share * Hospital HHI	-0.02**	-0.02**	-0.01
	(0.01)	(0.01)	(0.01)
Medicare Share * Physician HHI	-0.00	-0.00	0.00
	(0.01)	(0.01)	(0.01)
Medicare distance (%)		0.06**	
		(0.03)	
Medicare Distance * Hospital HHI		0.00	
		(0.00)	
Medicare Distance * Physician HHI		0.00	
		(0.00)	
Insurer Share of MA (%)		0.01	
		(0.00)	
Observations	5,671	5,671	2,884
Insurer FE	Yes	Yes	Yes
State FE	Yes	Yes	Yes
Controls	Yes	Yes	Yes
Log Likelihood	-2618	-2601	-1503
Pseudo R^2	0.323	0.328	0.165

Logit model, standard errors clustered on the county level. All variables are 2009 values. Exit=1 if insurers will removed all plans from this market between 2010-2012 and 0 if not.Control variables not shown. Additional stats include log-likelihood and McFadden's R^2. $^{**}p \leq .05$.

Table 2.4: Fit across models

	Prov./Ins. Vars	**Other Controls**	**County FE**	**Full Model**
All Counties				
Count	0.67	0.62	0.79	0.84
McFadden's R2	0.08	0.04	0.32	0.41
Counties with Overlap				
Count	0.69	0.67	0.84	0.84
McFadden's R2	0.15	0.04	0.40	0.41

Comparison of count statistics and R^2 across models. Column 1 includes insurer and provider market variables (Medicare share, Medicare distance, commercial market share, hospital and physician HHI, the percent vertically integrated, and indicators for missing values), Column 2 includes all other controls, Column 3 includes the full model (all controls + state and insurer fixed effects), and Column 4 includes county fixed effects and all controls. Rows 1 and 2 estimate models in the whole sample, and Rows 3 and 4 in the subsample that supports county fixed effects.

2.6 Conclusions

Competition in health insurance markets is a matter of critical importance, as competition may promote lower premiums and more generous policies.(9–14, 16, 17, 61) However, levels of competition in most health insurance markets are low and seem to be declining.(57) Insufficient entry may contribute to high market concentration, but why entry is so low is an open question. Many experts cite the difficulty of forming provider networks as a key deterrent, but there is little empirical evidence on whether this is the case.(22)

To test whether provider market concentration suppresses insurer market participation, this study examines selective exit among Medicare Advantage insurers who are forced to form provider networks from scratch. Recognizing that plans may use Traditional Medicare's "market power" to aid in bargaining with providers, insurers' bargaining power is measured using their market share of all Medicare. Both physician and hospital concentration are found to promote exit. Insurer bargaining power is found to protect against exit, with each additional percentage point of Medicare share reducing the probability of exit by 4 percentage points and each additional percentage point of an insurer's commercial share reducing the probability of exit by 1 percentage point. Interactions between insurer and provider variables show that Medicare share's protective effect is strongest in the counties with the least competition between hospitals. This supports the idea that insurer market power protects against exit by increasing insurer bargaining power.

These findings have several implications. The first is that both insurer and provider market structure matter for entry. For instance, in addition to raising insurer prices, mergers in hospital markets may also deter future insurer entry. The second is that insurers and hospitals seem to be engaged in bilateral bargaining over prices. This finding supports other studies finding that higher insurer market power has a countervailing effect on provider prices. Lastly, results suggest that administered prices can have strong spillover effects onto private markets. If Medicare Advantage insurers are bargaining using Traditional Medicare as leverage, than Traditional Medicare may be exerting significant downward pressure on provider prices. Without Traditional Medicare, Medicare Advantage premiums might significantly increase. This finding is of crucial importance for proposals restructuring Medicare.

Though this study uses a novel strategy to identify barriers to insurer market participation, it also has limitations. First, the factors that make an insurer more powerful with providers also increase their market power with consumers. We take the fact that insurer market power matters more in concentrated hospital markets as evidence that provider market power induced exit, rather than insurers' inability to increase premiums. The importance of commercial market share also lends validity to the bargaining model. However, absent stronger assumptions about consumer demand, effects cannot be fully separated.

Second, results suggest that physician market concentration matters both in markets where insurers were forced to form networks and markets where they were not. This may suggest that physician HHI, as measured here, captures more about variation in supply than variation in concentration or that physician market structure matters not just for bargaining over prices, but for whether physicians accept patients at all. Alternately, it may be spillover from insurers canceling plans that span multiple counties.

Additionally, the study design does not allow for separate identification of fixed and variable costs. Provider market concentration may affect these two parameters in opposite directions. Specifically, it may be easier to set up networks in concentrated markets, as there are fewer providers to negotiate with. However, the overall question of interest is about market participation, and any entering insurer incurs both fixed and variable costs of entering a market. Estimates likely reflect this overall net effect.

Lastly, having more insurers in a market is not unambiguously good. Though many studies emphasize the benefits of competition, excessive entry can lead to duplication of fixed costs.(84) In the case of Medicare Advantage, more insurer market participation also comes at a cost of public funds. However, as past results suggest that insurer exit came at cost to consumers in the form of reduced benefit generosity, insurer market concentration is still likely suboptimal in many markets.(61)

In summary, results suggest that both insurer and provider market power matter for insurer market participation, and that policies aimed at fostering entry must focus on both. In implementing coverage expansions, giving insurers additional tools for bargaining with providers may be at least as valuable as guaranteeing that insurers compete amongst themselves, particularly in markets where provider concentration is high.

Chapter 3

Trends in Medicare Advantage enrollment

3.1 Introduction

Medicare Advantage enrollment grew to its highest level in 2014, following a decade of continuous upward growth.(26) Growth in earlier years (2004-2008) coincided with increased plan availability and generosity. However, since 2009, policy changes have led to decreased plan availability and potentially less generous benefits.(26, 61, 85, 86) Despite these changes, enrollment in Medicare Advantage has continued to grow, defying experts' projections.(87, 88)

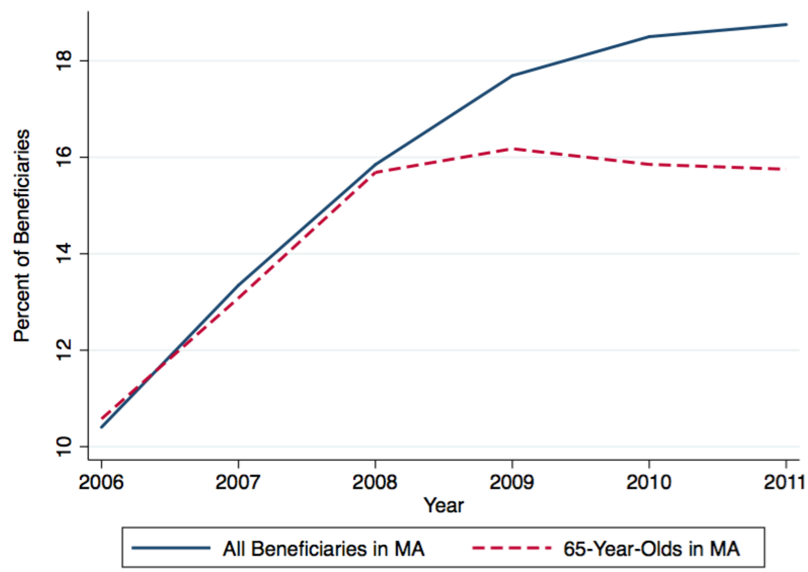
Whether increased enrollment is the result of more generous and widely available Medicare Advantage plans, changes in other insurance markets (i.e., Part D), or evolving beneficiary demographics and preferences is unknown. Understanding why enrollment has increased and whether increases will be sustained matters greatly for policy. The Medicare Advantage program currently enrolls a third of all Medicare beneficiaries and consumes a third of the Medicare budget. Though Medicare Advantage may provide beneficiaries with greater financial protection and more efficient care, it has cost more to cover a beneficiary in Medicare Advantage than in Traditional Medicare.(52, 86, 89–92) Historically, reductions in plan payments have resulted in reduced plan availability and decreased enrollment.(28) Understanding why enrollment has continued to grow, despite payment cuts and declining plan availability, may help policy-makers improve the program's efficiency.

Whether high Medicare Advantage enrollment will be sustained depends largely on the behavior of new enrollees. Initial choices in Medicare have been found to be very persistent(93, 94), so enrollment changes among younger cohorts may predict future program growth. To better understand how and why Medicare Advantage enrollment has changed, 65-year-olds' behavior is examined.

65-year-olds do not show the same continuous upward enrollment growth observed among all beneficiaries. Figure 3.1 compares average, county-level Medicare Advantage enrollment among 65-year-olds and the overall Medicare population. Data span years in which plan availability and payments both grew (2006-2009) and declined (2009-2011). 65-year-olds' enrollment increased at the same rate as overall enrollment between 2006-2009, but diverged from older cohorts and significantly leveled off in 2009.

This paper describes theories that may explain recent enrollment growth and tests which theories might fit 65-year-olds' enrollment trends. Changes in Medicare Advantage benefit generosity and premiums most plausibly explain the leveling in 65-year-olds' enrollment. Additionally, 65-year-olds' take-up of Medicare Part A (hospital insurance) during this time period changes in ways that suggest beneficiaries were losing jobs and employer-sponsored insurance. Absent these changes, enrollment in Medicare Advantage might have declined more sharply. Should benefit generosity further decline and the economy improve, results suggest that enrollment growth may soon slow.

Section 3.2 describes Medicare Advantage and supplemental insurance in Medicare, and Section 3.3 discusses theories for why Medicare Advantage enrollment increased. Section 3.4 describes data and methods used to test whether these theories fit 65-year-olds'



Calculated from enrollment files. The percent of beneficiaries enrolling is averaged at the county level; levels are therefore lower than other reported statistics.

Figure 3.1: Average, county-level percent of all beneficiaries and 65-year-olds enrolled in Medicare Advantage

enrollment patterns, Section 3.6 describes results, and Section 3.7 concludes.

3.2 Medicare Advantage and supplemental insurance in Medicare

The Medicare program provides health insurance for the disabled and elderly. At age 65, most Americans become eligible for coverage in the Medicare Part A (hospital) and Part B (medical) insurance programs. Medicare covers a wide range of services, including inpatient hospital stays, skilled nursing care, physician visits, and many outpatient services. However, there are significant gaps in Medicare benefits. Most services are covered with high cost-sharing, there are no caps on total beneficiary out-of-pocket spending, and many categories of care are not covered at all.[1] As a result, most beneficiaries (82% in 2010) obtain some sort of supplemental coverage.(35)

A variety of options exist to fill the gaps in Medicare coverage. For 20% of beneficiaries, the Medicaid program supplements Medicare. However, to qualify for Medicaid, beneficiaries must have very low incomes and assets.(95) Another 20-30% of beneficiaries are covered by employer-sponsored retiree health insurance. These plans are usually subsidized by employers and reduce Medicare cost-sharing or cover additional services. However, the percent of employers offering this coverage has declined by as much as 40% over the past thirty years.(96, 97)

Beneficiaries can also purchase individual supplemental insurance plans (called Medigap plans) or stand-alone prescription drug plans (Part D plans) to fill holes in Medicare coverage. Medigap plans cover a portion of deductibles, copayments, and coinsurance, but often have high premiums and do not limit out-of-pocket spending.[2] Additionally, as of 2006, newly issued plans cannot cover prescription drugs.

[1]For instance, Medicare Part A has a high deductible ($1,216 in 2014), and Medicare Part B has a 20% coinsurance on most services, after a $147 deductible. Prescription drugs, vision services, and long-term care are not covered under Medicare Parts A and B.

[2]Insurers in most states can only sell policies with federally standardized benefit packages, identified by letters A through N. Only plan types K and L, which together enrolled less than 1% of Medigap beneficiaries in 2010, cap out of pocket spending.(98)

The Medicare Part D program, created in 2006, allows private insurers to sell plans covering only prescription drugs. These plans are subsidized by Medicare payments, so premiums are less than the cost of coverage. However, plan generosity has declined since the program's inception, and Part D plans do not cover any portion of medical spending.(99)

In Medicare Advantage, beneficiaries enroll in private plans that replace rather than supplement Traditional Medicare. The Medicare program pays private insurers a per-beneficiary fee for accepting enrollees, and Medicare Advantage enrollees receive all coverage for Medicare benefits from that insurer. These plans usually provide additional benefits, reducing cost sharing, limiting total out-of-pocket expenditure,[3] and/or covering additional services. However, beneficiaries who enroll in Medicare Advantage sacrifice provider choice for financial protection. While beneficiaries in Traditional Medicare can visit any doctor accepting Medicare, Medicare Advantage beneficiaries must visit providers in their insurers' network or face higher out-of-pocket costs.[4]

The share of beneficiaries enrolling in Medicare Advantage (also called Medicare Part C and previously, Medicare + Choice) has grown and shrunk over the program's history, correlated largely with plan availability and Medicare payments to plans.[5] Payment increases are generally associated with greater plan availability(19, 28, 66), better benefits(19, 29, 100), and lower premiums(19, 40, 100), which, in turn, have historically been associated with more beneficiaries enrolling in Medicare Advantage.

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[3] Medicare required all Medicare Advantage plans to have out-of-pocket limits on spending for Medicare-covered services starting in 2011. Limits are allowed to be higher for out-of-network services.(86)

[4] HMO plans usually do not offer out-of-network coverage. PPOs do, but often charge beneficiaries substantially higher out-of-network copays. For a limited time (1997-2008), insurers could offer private fee-for-service plans, in which insurers was not required to form provider networks and beneficiaries could visit any provider in Medicare at no extra cost. The 2008 Medicare Improvements for Patients and Providers Act changed the law and forced PFFS plans to limit provider networks in the majority of counties.

[5] McGuire, Newhouse, and Sinaiko (2011) review the history of enrollment and plan availability in Medicare Advantage.(28)

Recent policy changes were projected to lead to decreased enrollment.(87, 88) Medicare payments to plans were reduced under the Medicare Improvements to Patients and Providers Act of 2008 (MIPPA) and further cut under the Affordable Care Act in 2010.[6] MIPPA also forced private-fee-for-service (PFFS) plans, a type of Medicare Advantage plan which offered unlimited provider choice, to form limited provider networks. These new requirements and payments cuts were projected to decrease plan availability, especially among PFFS plans.(102) Decreased availability was, in turn, projected to negatively impact enrollment, particularly as PFFS plans had been responsible for dramatic enrollment growth in many areas.(30, 87, 88)

Consistent with predictions, plan availability has declined since 2009. Much of the decline in availability is due to widespread cancellation of PFFS plans, though payment cuts also led to exit among HMOs and PPOs.(85, 86) There is also evidence that plan benefit generosity has declined, though plan premiums have largely stayed steady.(61, 103) Contrary to predictions, Medicare Advantage enrollment has continued to climb, with a third of all beneficiaries enrolling in Medicare Advantage in 2014.(26) Enrollment seems unaffected by recent policy changes, staying high even in counties where payment and plan availability declined.(85)

3.3 Why did Medicare Advantage enrollment increase?

There are several plausible theories for why Medicare Advantage enrollment increased over the last decade. Some, but not all, also plausibly explain why it has stayed high. Theories can be organized into those pertaining to changes in Medicare Advantage market, changes in related insurance markets (Medigap, Part D), and changes in characteristics

[6] MIPPA reduced benchmark payments by gradually eliminating duplicate payments for indirect medical education. Reducing indirect medical education payments was projected to reduce plan payments by 2.2%.(33) The Affordable Care Act froze payment levels in 2011 and phased in payment reductions, starting in 2012.(101)

or preferences of Medicare beneficiaries. Table 3.1 summarizes theories and evidence for and against each.

Increased plan availability is still frequently cited as a cause of increased enrollment.(104) Past research supports the idea that greater plan availability would increase enrollment, as the two have historically been correlated. Why enrollment might stay high amidst declining availability is less clear, though there are a few plausible explanations. Beneficiaries tend to show substantial inertia in coverage choices, so sustained enrollment might be an artifact of expanded plan availability.(85) There is also evidence that changes in the number of plans have little effect on enrollment, provided beneficiaries have access to "enough" options.(93, 105) As the number of markets with at least five or ten plans has grown and stayed high, decreases above that number may not affect enrollment (see Appendix Figure C.1).

Lower premiums and better benefit generosity are also frequently cited as causes for increased enrollment. Medicare Advantage proponents note that plan premiums have stayed low, despite payment cuts, and that plans protect beneficiaries from high out-of-pocket expenses. Here, too, the literature supports the idea that lower premiums/ better benefits would increase enrollment, as a large literature on Medicare Advantage plan choice shows that low premiums/generous benefits increase plans' market share.(14, 56, 67, 105–109)

There are several plausible explanations for why enrollment might have stayed high in the face of declining generosity. Low premiums might lead to beneficiaries continuing to enroll, particularly if premiums are the most important or salient characteristic of plans. Alternately, observed benefit decreases may primarily affect sicker beneficiaries. For instance, Gold (2014) finds that insurers are mainly adjusting out-of-pocket limits.(26) As few beneficiaries hit these limits, changes in this characteristic might not substantially impact enrollment.

Changes in other supplemental insurance markets could also explain increased

Medicare Advantage enrollment. It is often noted that increased enrollment coincided with the introduction of Medicare Part D. Why the introduction of Part D would increase Medicare Advantage enrollment is not clear, especially as no research to date has tested whether enrollees treat the two programs as substitutes. Passage of Part D might have alerted beneficiaries to the existence of Medicare Advantage as an option for covering medications. Alternately, Medicare Part D plans' generosity has declined since the program's inception, which might cause beneficiaries to switch to Medicare Advantage.(99) However, research pre-dating Part D found that drug coverage was the most valuable feature of Medicare Advantage.(14, 107, 109) If Medicare Advantage plans are a substitute for Part D, then Part D's implementation should have reduced Medicare Advantage enrollment.

Changes in Medigap could also affect Medicare Advantage enrollment. Past research shows that beneficiaries treat Medigap and Medicare Advantage as substitutes, (105, 107, 108, 110, 111) and the introduction of Part D significantly changed Medigap benefit generosity. Following introduction of Part D, new Medigap plans were no longer allowed to offer drug coverage. A beneficiary needing drug and medical coverage would need to purchase both a Part D and Medigap plan, which might result in new beneficiaries "crowding in" to Medicare Advantage. However, enrollment in Medigap market has stayed largely steady throughout this time period. This makes it less likely that increased Medicare Advantage enrollment is due primarily to decreased Medigap enrollment.(98)

The last supplemental market that might affect Medicare Advantage enrollment is retiree supplemental insurance. Though there is very little evidence on substitution between retiree health insurance and Medicare Advantage, and based on their characteristics, retiree health plan enrollees would likely opt into Medigap plans if they lost their coverage.(96) However, the availability and generosity of retiree health plans has declined dramatically in recent decades,(96, 97, 112) and declining availability of retirement benefits might increase Medicare Advantage enrollment.

Demographic or economic changes might contribute to increased Medicare Advantage enrollment. Younger beneficiaries are more likely to join Medicare Advantage and currently make up a higher proportion of Medicare than ever before.(93, 105, 113, 114) Additionally, new cohorts entering Medicare Advantage have spent more of their lives insured by HMOs and may therefore be more comfortable with managed care. Both of these phenomena might lead to increasing upward enrollment trends.

Worsening economic outcomes for the middle class might also explain high enrollment. Medicare Advantage generally appeals to poorer beneficiaries, providing coverage for those too rich for Medicaid and too poor for Medigap.(105, 114–116) As retirees lost income and assets during the recession, the percent of the population that found Medicare Advantage appealing may have expanded.

For the young elderly, changing employment patterns might also explain rising Medicare Advantage enrollment. If the recession led to the young elderly (those 65-70) losing their jobs and employer-sponsored insurance, this might increase the population enrolling in Medicare Advantage. However, neither this nor the prior theory perfectly fit the evidence, as the rise in Medicare Advantage enrollment pre-dates the recession.

Data are assembled to test which of these theories might fit enrollment patterns observed among 65-year-olds. Data capture variation in plan availability, premiums, and benefit generosity in Medicare Advantage, Medigap and Medicare Part D plans. Information on employment, income, and employer-sponsored insurance options are not available in Medicare data. In lieu of this data, aggregate data on county-level income and employment are used, combined with indicators constructed from Medicare data that may reflect shifts in employment or income in the Medicare population.

Table 3.1: Summary of theories regarding increased Medicare Advantage enrollment

Theory	Evidence for	Evidence against
Changes in the Medicare Advantage market		
Plan availability	Increased availability coincided with increased enrollment	Availability has declined without impacting enrollment
	Inertia might keep enrollment high	
Lower premiums	Reduced premiums coincided with increased enrollment	
	Premiums have stayed low, despite declining plan availability	
Better benefits	Benefit improvements coincided with increased enrollment	Declining generosity has not impacted enrollment
	Generosity declines may not affect healthy enrollees	
Changes in related markets		
Part D market	Part D implementation coincided with increased enrollment	Part D might be a substitute for MA, which would lead to decreased enrollment
Medigap	Medigap and Medicare Advantage are substitutes	Medigap enrollment has stayed steady since 2006
Medigap	Medigap benefits became less generous due to Part D	
Retiree plans	Retiree health insurance availability is declining	Retiree health plan beneficiaries likely prefer Medigap
Changes in population characteristics		
Age	Younger beneficiaries prefer Medicare Advantage and population of younger beneficiaries increasing	
Income	Poorer beneficiaries prefer Medicare Advantage and incomes are falling during the recession	Increased enrollment pre-dates recession
Employment	Loss of employment/employer-sponsored insurance among young elderly might increase Medicare Advantage take-up	Increased enrollment pre-dates recession

3.4 Methods and data

There are three goals of this analysis: 1) to test which factors are correlated with 65-year-olds' enrollment in Medicare Advantage, 2) to test which factors could plausibly explain the enrollment trend break, 3) to test which factors are correlated with initial upward enrollment trends. Visual evidence suggests there are two distinct periods in 65-year-olds choosing Medicare Advantage: 2006-2009 and 2009-2011 (see Figure 3.1).[7] To test which factors might explain trends, a regression with a linear time trend, a time break, and county-level time-varying covariates is specified. The specification is:

$$Percent\_65yolds_{mt} = \beta_0 + \beta_1\tau_t + \beta_2\tau_t * (t \geq 2009) + \beta_3(t \geq 2009) + \gamma_1 X_{mt} + \theta_m + \varepsilon_{mt} \tag{3.1}$$

where $Percent\_65yolds_{mt}$ is the percent of 65-year-olds in market m at time t enrolled in Medicare Advantage, τ is a linear time trend, and θ_m are county fixed effects. X_{mt} includes average, county-level characteristics of Medicare Advantage, Medigap, and Part D plans and economic conditions in market m at time t.

The coefficient β_1 captures the overall trend in the percent of 65-year-olds choosing Medicare Advantage, while coefficients β_2 and β_3 capture whether there is a break in enrollment trends in 2009.[8] Coefficients γ_1 show which factors are correlated with enrollment, and county-fixed effects absorb persistent differences in 65-year-olds' enrollment. All models are weighted by the average number of beneficiaries in a county across all years of the sample, so that results are less sensitive to noise in small counties.

Data are combined from a range of person- and plan-level Medicare files and supple-

[7] 2006 was chosen as the baseline year for analysis because plan-level Medicare Advantage data were not available prior to 2006. 2011 was the most recent year of Medicare enrollment data available at the time of this research.

[8] Data suggest a trend break if β_2 and β_3 are jointly different from 0. A Quandt-likelihood ratio test is used to test whether the data suggest that the break occurred in 2008 or 2010, rather than 2009. Coefficients $\beta_1 + \beta_2$ capture the slope of enrollment trends, following the break.

mented by county-level data from the Bureau of Economic Analysis, Bureau of Labor Statistics, and Area Resource File. Data are reviewed briefly here and discussed further in the data appendix.

The percent of 65-year-olds enrolling in Medicare Advantage in each county in a year is calculated from Medicare enrollment records. Beneficiaries are counted as enrolled in Medicare Advantage if they enrolled in a plan at any point during the year. Beneficiaries who are dual eligible or classified as having end-stage renal disease are excluded, as is any county-year observation with fewer than twenty 65-year-olds (to reduce noise).

To test whether variation in plan availability affects enrollment, the number of Medicare Advantage and Medicare Part D plans available in a county are added as controls. The number of plans in a county is calculated from aggregate Medicare enrollment data. The number of Medicare Advantage plans includes HMOs, PPOs, PFFS plans, and regional PPOs, and the number of Part D plans includes all stand-alone Part D plans (PDPs).[9] Counts of plans exclude any plan with fewer than 11 enrollees (the level at which Medicare censors aggregate data), and the study population is limited to any county-year with at least one Medicare Advantage plan.[10] The final set of counties analyzed included a total of 16,848 county-years (between 2,937 and 2,952 counties each year).

To test whether variation in plan characteristics affects 65-year-olds' enrollment, data on plan premiums and benefit generosity in Medicare Advantage, Medigap, and Medicare Part D are constructed from Medicare out-of-pocket cost datasets and landscape files. To capture the characteristics of enrollees' choice sets, average premiums/benefits for the

[9]Employer-sponsored Medicare Advantage plans are excluded from counts of plans, as they are only available to a subset of the population. However, as plan identifiers are encrypted in enrollment files, enrollees in employer-sponsored plans are not excluded from the dependent variable. To address this, the percent of Medicare Advantage enrollees in a county in an employer-sponsored plan is calculated from aggregate Medicare enrollment and added as a control variable.

[10]In the appendix, this assumption is varied by constructing a balanced panel of counties and setting the number of plans equal to 0 for years when no plans were offered in a county.

set of plans offered in a county (or state for Medigap) are calculated. Averages are unweighted, so that they reflect the characteristics of available plans, not the characteristics of plans chosen by enrollees.[11]

Medicare Advantage premiums are split into medical (Part C) and drug (Part D) portions, and Part C premiums are adjusted to reflect whether insurers reduce beneficiaries Part B premiums.[12]

Medigap premiums for a representative beneficiary are estimated used average, statewide premiums for plans F and C. A broader range of Medigap options are available, but many benefit packages enroll relatively few beneficiaries. Plans F and C together account for 54% of all Medigap enrollment nationwide and may therefore be representative of Medigap plan generosity. Medigap premium data is drawn from a industry representative source and reflects average premiums for beneficiaries 65 and above.

Benefit generosity is measuring using a summary statistic called out-of-pocket cost, which captures a representative beneficiary's spending under each plan's benefit design. It reflects variation in copays, deductibles, and coverage for a range of services, including inpatient, outpatient, lab and diagnostic services, skilling nursing services, and prescription drugs. Out-of-pocket cost is calculated on the plan-level for Medicare Advantage and Part D plans and on the benefit package level for Medigap. (Medigap plan generosity does not vary across states.) For comparison, Traditional Medicare out-of-pocket cost is calculated for a beneficiary with no supplemental insurance.

Out-of-pocket cost data is missing for all Medicare Advantage and Medigap plans in 2006, as is 2006 premium data for Medicare Advantage plans that do not offer drug coverage (60% of plans in 2006). For Medicare Advantage plans, these data are imputed using average OOPC/premiums for the same plan in 2007 and 2005. For Medigap, out-of-pocket costs are imputed using average levels for each state-plan combination in

[11] Enrollment-weighted averages primarily reflect the choices of older cohorts, not 65-year-olds.

[12] Very few plans do this – $2.3 - 4.7\%$ depending on year.

2005 and 2007. (Further detail on constructing 2006 data is in the Data Appendix.)

Out-of-pocket cost data for stand-alone Part D plans are not available until 2008. These variables are included in some regressions to test whether PDP out-of-pocket costs affect Medicare Advantage enrollment. However, these regressions cannot be used to test for changes in trends because of the limited number of years.

Though nationally representative data on the elderly's income, employment, and health insurance exist, none cover all markets analyzed here.[13] Instead, a combination of county-level data for the whole population and indicators from Medicare enrollment files are used to capture changes in economic conditions. County-level data include nominal per-capita income (from the Bureau of Economic Analysis local area personal income estimates), unemployment rates for the population over 18, and the percent of a county's population below poverty (both from the Area Resource File).

Several variables which may reflect systematic changes in Medicare beneficiaries' employment or income are also calculated. They include: the percent of 66-year-olds who are enrolled in Medicare in the prior year, the percent of beneficiaries enrolled in Medicare Part A only, the percent of beneficiaries who are dual eligible, and the percent of beneficiaries who originally qualified for Medicare due to disability.

The percent of 65-year-olds enrolled in only Medicare Part A and the percent of 66-year-olds observed in the prior year's data may both plausibly reflect changes in employment and/or employer-sponsored insurance. Changes in employer-sponsored insurance affect Medicare Advantage enrollment because a beneficiary cannot enroll in Medicare Advantage if they are covered by their employer. Employment changes are difficult to observe because only beneficiaries who are retired and collecting social security are automatically enrolled in Medicare.

The percent of beneficiaries enrolled in only Medicare Part A may reflect continued

[13]The choice was made to analyze the universe of Medicare data to better capture changes in small markets. Medicare Advantage enrollment in large counties is often very stable, whereas enrollment in small markets is more sensitive to changes in market characteristics.

beneficiary employment and employer-sponsored insurance. 65-year-olds who are actively working and deferring social security are not automatically enrolled in Medicare, but are contacted by Medicare and invited to do so. These individuals are often advised to enroll in Medicare Part A and defer taking up Part B, because it is costless for them to do so.

The percent of 66-, 67-, and 68-year-olds observed in the prior year's data may also reflect changes in employment. If employed 65-year-olds defer enrollment in both Medicare Parts A and B, they will not appear in Medicare files. However, if they enroll at later ages, this can be detected by testing whether they appeared in prior year files.[14] In some regressions, the percent of 66-year-olds enrolled in Medicare in year $t+1$ is included as a control, as it reflects the percent of 65-year-olds enrolled in Medicare in current year t.

Lastly, two characteristics of the Medicare population may indicate changes in income and preferences: the percent of beneficiaries who are dual eligible and the percent of beneficiaries who qualified for Medicare because of disability (rather than age.) Both Medicaid eligibility and disability increased during the recession.(117, 118) Changes in the dual eligible population change the set of beneficiaries appearing in the data and may change the percent of 65-year-olds enrolled in Medicare Advantage, as poorer 65-year-olds often prefer Medicare Advantage. Beneficiaries who originally qualified for Medicare through disability are not excluded from the study population, but increased disability enrollment may systematically change beneficiary preferences. For instance, the disabled are systematically more likely to dis-enroll from Medicare Advantage plans. Hence, increases in disability might decrease the percent of beneficiaries choosing Medicare Advantage.(113)

[14] Due to change in beneficiary ID, beneficiaries cannot be matched between 2005 and 2006, so this statistic cannot be generated for 2006.

3.5 Descriptive statistics

Table 3.2 summarizes Medicare Advantage enrollment and market characteristics. Rows 1 and 2 summarize the average, county-level percent of 65-year-olds and beneficiaries enrolling in Medicare Advantage. From 2006-2009, 65-year-olds' enrollment grew at a rate similar to enrollment in all Medicare. The average, county-level percent of beneficiaries in Medicare Advantage grew from about 10% in 2006 to about 15% in 2009. Following 2009, the percent of all beneficiaries enrolled in Medicare Advantage continued to increase, growing from 15 to 18%, while 65-year-olds' enrollment leveled off.[15]

The average number of Medicare Advantage plans available in a county rose swiftly between 2006-2009 and decreased following implementation of payment freezes and restrictions on PFFS plans (Rows 5-7). However, the fall-off in plan availability during this time period is largely driven by insurers' removal of PFFS. Availability of other plan types generally continued to increase.

Throughout this time period, there were consistently more stand-alone Part D drug plans (PDPs) available to beneficiaries than Medicare Advantage plans (Row 8). However, trends in PDP availability are similar to those in Medicare Advantage. Availability grew through 2008 and then decreased. By 2011, the average number of PDPs in a county (n=21) was lower than it was in 2006 (n=23.7).

Rows 9-12 summarize average, monthly premiums for Medicare Advantage, Medicare Part D, and Medigap plans, with Medicare Advantage premiums broken into Part C (medical) and Part D (prescription drug) components. The most notable pattern is the consistent, $100 gap between average monthly Medigap and Medicare Advantage premiums. Average premiums in Medigap are around $130 per-beneficiary-per-month,

[15]County-level average enrollment statistics are lower than the overall percent of Medicare enrolled in Medicare Advantage (Rows 3 and 4), consistent with the fact that Medicare Advantage enrollment is driven by larger counties. However, broad patterns are similar.

Table 3.2: Enrollment and benefits in Medicare Advantage (2006-2011)

	2006	2007	2008	2009	2010	2011
MA Enrollment (avg. across counties)						
65-year-olds (%)	9.7	12.7	15.3	15.8	15.6	15.5
All Beneficiaries (%)	9.5	12.9	15.4	17.3	18.2	18.5
MA Enrollment (aggregated across Medicare)						
65-year-olds (%)	15.1	17.4	20.3	21.3	21.1	21.2
All Beneficiaries (%)	18.5	20.9	23.4	25.4	26.1	26.6
Avg. Number of Plans in a County						
Number of Medicare Advantage plans	4.4	7.0	9.4	10.7	9.2	7.5
Number of PFFS plans	2.0	4.3	5.9	6.3	4.2	1.9
Number of HMO/PPO/RPPOs	2.4	2.7	3.6	4.4	4.9	5.6
Number of Part D Plans	23.7	26.0	28.7	27.9	26.1	21.0
County Avg. Premiums						
MA Part C Premium	25.2	22.2	23.0	21.1	32.7	28.4
MA Part D Premium	10.2	9.6	11.2	15.3	14.3	15.8
Part D Plan Premium	31.9	33.9	34.7	38.9	40.9	47.4
Medigap Premium	122.6	129.8	126.3	133.1	123.6	129.4
County Avg. Benefits (excluding premiums)						
MA out-of-pocket cost	257.0	225.4	172.6	179.6	241.2	269.6
Medigap out-of-pocket cost	255.6	208.8	217.7	157.4	269.1	336.8
Part D Plan out-of-pocket cost	.	.	55.1	55.4	109.7	122.0
Traditional Medicare out-of-pocket cost	322.6	322.6	352.0	277.6	403.4	485.3

Expected out-of-pocket costs exclude Part B premiums for all plans, Part C and D premiums for Medicare Advantage plans, and Part D premium for stand-alone Part D plans. Medicare Advantage Part C premiums reflect Part B reductions.

while Part C Medicare Advantage premiums average between $\$20-30$.[16] Both are largely steady over time.

Part D premiums in both Medicare Advantage and PDPs rose steadily throughout this time period, though PDP premiums are consistently higher. PDP premiums rose from an average of $31 per-member-per-month to $47 per-member-per-month, while MA-PD premiums rose from $10-15. Dividing premiums into components perhaps understates the generosity gap between Medicare Advantage and other options, particularly as Medigap plans issued after 2006 do not cover prescription drugs. A beneficiary who wanted both medical and drug coverage in Medigap would also need to pay $\$30-45$ a month for PDP.

Medicare Advantage, Medigap, and PDP generosity, as measured by average expected out-of-pocket costs, are compared to the estimated out-of-pocket cost for an enrollee without supplemental coverage (Rows 13-16). (Estimates exclude Part B, C, and D premiums.) Average expected out-of-pocket costs for a beneficiary in Medicare Advantage fell from $256 per-member-per-month to $157.4 between 2006-2008, before rising back to $336 by 2011. However, expected out-of-pocket costs in Medicare Advantage are generally lower than Medigap throughout this time period. Additionally, monthly out-of-pocket costs with both types of supplemental coverage are $67-216 less than with no supplemental insurance. Part D stand-alone costs are lower as they only reflect drug spending. However, these too increased by more than 200% between 2008-2011.

Table 3.3 summarizes county-level economic indicators and Medicare population characteristics. The unemployment rate most clearly reflects the recession, increasing from 5.9% in 2008 to 8.7% in 2011 (Row 1). The average percent of county population below poverty also increased throughout this time period, from 15.4% in 2006 to 17% in

[16]Estimates of average monthly Medigap premiums are about $50 a month lower in this data than other estimates.(98) One potential reason for this is that these estimates are constructed without accounting for enrollment. Another potential difference is that these estimates exclude premiums for beneficiaries under 65, who are generally charged substantially more for Medigap plans than older beneficiaries.

Table 3.3: Employment and income statistics

	2006	2007	2008	2009	2010	2011
County Characteristics						
County-level unemployment rate	5.0	4.9	5.9	9.2	9.4	8.7
County below poverty (%)	15.4	15.1	15.3	16.4	16.9	17.4
Per capita income (1000's)	30.0	31.8	33.7	32.9	33.7	36.1
65-year-olds' Characteristics						
% of 65-yr-olds who are dual eligible	10.6	10.5	10.5	11.3	12.3	12.0
% of 65 yr-olds qualified by disability	10.5	10.6	11.0	11.3	11.4	11.4
% of 65-yr-olds in Part A only	18.2	19.3	19.7	14.9	15.3	15.4
All Beneficiaries' Characteristics						
% of beneficiaries who are dual eligible	24.8	24.9	24.8	25.2	25.7	26.2
% of beneficiaries qualified by disability	16.2	16.3	16.4	16.5	16.7	16.8
% of beneficiaries in Part A	6.9	7.4	7.8	6.7	6.8	6.9
Percent of 66+ year-olds observed in Medicare last year						
% of 66-yr-olds in Medicare last year	.	97.0	96.6	96.6	98.5	98.5
% of 67-yr-olds in Medicare last year	.	99.3	99.2	99.3	99.6	99.6
% of 68-yr-olds in Medicare last year	.	99.6	99.5	99.6	99.8	99.8

Percent dual eligible, percent qualified through disability, percent of beneficiaries in Part-A-only, and the percent of age groups observed in last year's Medicare files reflect overall statistics in Medicare, not county-level averages. Construction of these variables is discussed in the data section.

2011 (Row 2). There is no clear trend in nominal income (Row 3).

Both the percent of 65-year-olds who are dual eligible and who originally qualified for Medicare due to disability also increased (Rows 4-5). The percent of 65-year-olds who were dual eligible was steady at around 10.5% between 2006-2008, before increasing to 11-12% in 2009-2011. The percent of 65-year-olds who qualified for Medicare through disability increased steadily from 10.5 to 11.4%. 65-year-olds were less likely to be dual eligible or disabled then the overall Medicare population (Rows 7-8), but trends were similar.

The percent of 65-year-olds in only Medicare Part A dropped sharply around the recession, decreasing from 18.2% in 2006 to 14.9% in 2009 (Row 6 and Figure 3.2). This may suggest that many individuals lost their jobs and employer-sponsored insurance during this time period. The percent of 65-year-olds enrolled in Part A only is very high, relative to all Medicare (15-20% vs. 6-8%), which is consistent with this statistic reflecting continued employment.

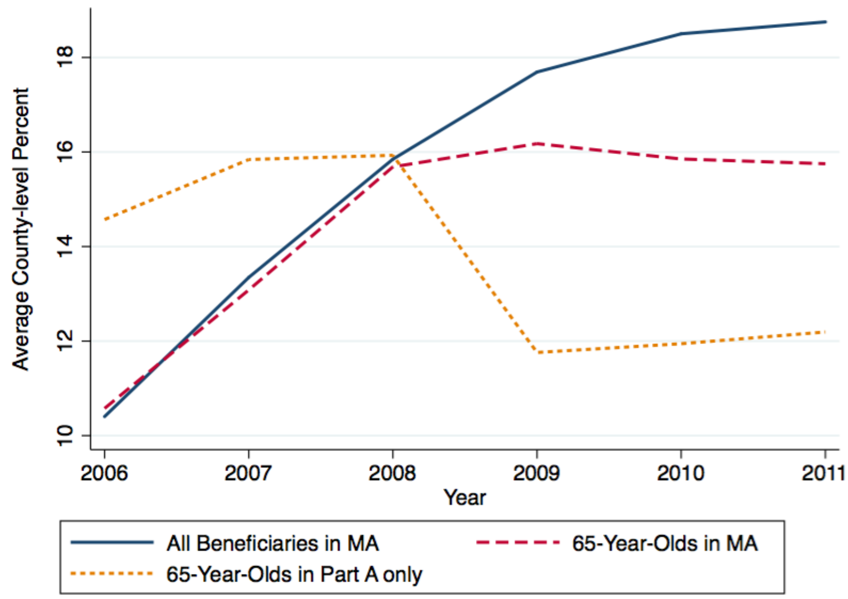


Figure shows the percent of 65-year-olds enrolled in only Medicare Part A, overlaid on the percent of 65-year-olds enrolled in Medicare Advantage. 65-year-olds are counted as being in Part A only if they are enrolled in just Part A or out of Medicare for the entire year.

Figure 3.2: Trends in Medicare Advantage enrollment and beneficiaries in Part-A-only

Patterns in 66-, 67-, and 68-year olds enrolled in Medicare in the prior year are less obviously pro-cyclical, and the percent of these cohorts observed in last year's data is generally quite high (> 96%). The percent of 66-year-olds observed in the prior year's data is generally about 2 percentage points lower than the percent of 67- or 68-year olds, consistent with these statistics reflecting delayed Medicare enrollment.

To test whether dual eligibility, disability, delayed enrollment, or Part-A-only enrollment plausibly reflect economic changes, I test for correlations between these variables and county-level unemployment. Table 3.4 shows the results of regressing each variable on unemployment rates in regressions with county- and year-fixed effects. Three of the four variables are significantly correlated with economic conditions. An additional percentage point of county-level unemployment is associated with a .12 percentage point increase in the percent of 65-year-olds who are dual eligible, a .03 decrease in the percentage of 65-year-olds who qualified for Medicare due to disability, and a .15 decrease in the percent of 65-year-olds in Part A only. The percent of 65-year-olds who can be matched to next year's data is not significantly correlated with unemployment, perhaps because this number is generally high.[17]

These results are consistent with the idea that worse economic conditions increase the percent of 65-year-olds in poverty (as measured by dual eligibility) and increase the percent of 65-year-olds who are unemployed (as measured by enrollment in only Part A). The sign on the percent of 65-year-olds who are disabled is counterintuitive, but may reflect the fact that the disabled are out of the workforce and not counted in the unemployment rate.

Table 3.4: Correlation between Medicare population characteristics and unemployment rates within counties over time

VARIABLES	(1) % Duals	(2) % Disbld	(3) % PartA	(4) % Age 66
Unemployment rate	0.12**	-0.03**	-0.15**	0.04
	(0.05)	(0.01)	(0.01)	(0.04)
Observations	16,848	16,848	16,848	11,354
County FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes

Outcome variables are 65-year-olds' characteristics: the percent of 65-year-olds who are dual eligible, disabled, enrolled in Part A only, or enrolled in Medicare this year. Standard errors clustered on the county level. Observations on the county-year level. Weighted by number of beneficiaries in a county, averaged across years. $^{**}p \leq .05$.

[17]Using a lead of the percent of 66-year-olds in the data results in 2011 being omitted. 2006 is already omitted because beneficiary IDs cannot be matched across 2005 and 2006.

3.6 Explaining trends

To test which factors are correlated with 65-year-olds' enrollment, a model with linear time trends and time-varying controls is fitted to the data. Column 1 of Table 3.5 shows unadjusted trends in 65-year-olds' enrollment. Consistent with patterns in Figure 3.1, the percent of 65-year-olds choosing Medicare Advantage grew between 2006-2009 by an average of 2 percentage points a year. In 2009, enrollment leveled off, decreasing by a statistically significant $-.14$ percentage points a year through 2011.[18]

To test whether trends are similar across counties, counties are divided into quartiles based on the 2006 percent of 65-year-olds enrolled in Medicare Advantage. Figure 3.3 shows trends across groups, where trends in Panel B are normalized to 2006 levels. Enrollment trends are largely similar across groups of counties, with enrollment rising through 2008, and leveling off in 2009.[19]

Plan availability in a county is correlated with enrollment, but the trend break remains significant after controlling for the number of plans in a market (F=32.48, Col. 2 of Table 3.5) Each additional Medicare Advantage plan in a county increases the percent of 65-year-olds who join Medicare Advantage by .15 percentage points, while each additional PDP reduces enrollment by a statistically significant -.07 percentage points. Changes in plan availability reduce but do not, by themselves, fully explain trends.

Including average Medicare Advantage premiums and benefits reduces the magnitude of residual upward trends from 2006-2009 and makes the 2009 trend break insignificant (Column 3). Increasing average monthly out-of-pocket costs by $10 reduces the percent of 65-year-olds enrolling in Medicare Advantage by .25 percentage points, while an

[18] The break is statistically significant at $F = 139.1$. Trend breaks in 2008 and 2010 are also tested, using a Quandt-likelihood ratio test. Joint F-statistics (β_2=β_3=0) are highest in 2009 (139.1 vs. 111.4 and 98.4 respectively) and well above the level at which QLR statistics indicate a significant trend break.

[19] Trend breaks in all quartiles are jointly statistically significant at $F = 79.09$. There are statistically significant differences in trends across counties. However, trends are generally upward between 2006-2008 (between 2.2-3.1% per year) and generally flat or decreasing thereafter.

Table 3.5: Trends in 65-year-olds' enrollment in Medicare Advantage with county-level controls (2006-2011)

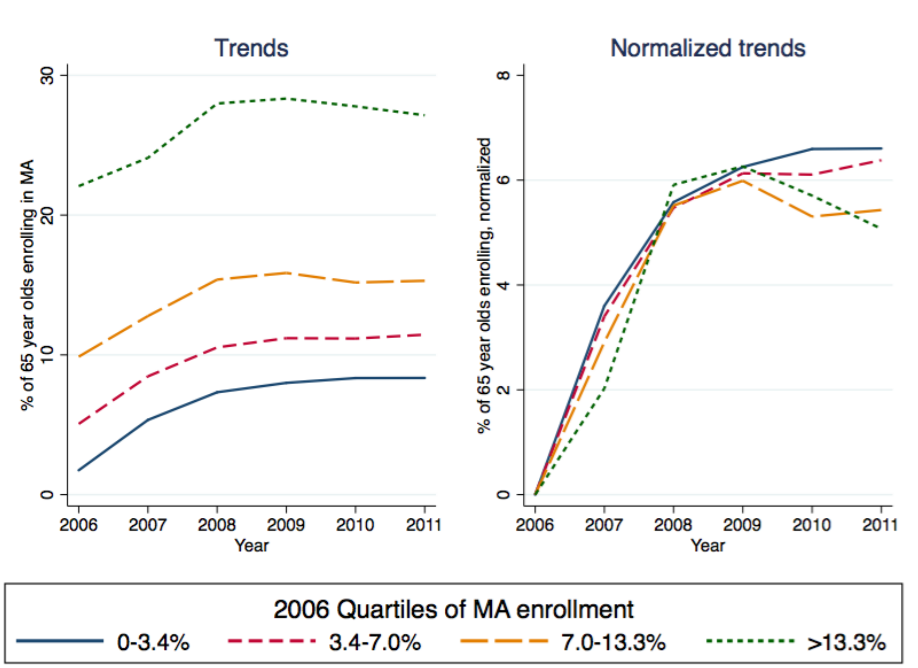
VARIABLES	(1) Trends Enrollment	(2) Trends +N Plans	(3) Trends +Benefits	(4) Trends +Benefits & Plans	(5) Trends +Related Markets	(6) Trends +Everything
Trend	2.09**	1.75**	1.11**	0.93**	0.88**	0.78**
	(0.10)	(0.15)	(0.16)	(0.18)	(0.18)	(0.29)
Trend 2009 and later	-2.23**	-1.75**	0.20	0.28	0.45	
	(0.14)	(0.22)	(0.37)	(0.35)	(0.35)	
Year 2009 or later	1.30**	0.65**	0.16	-0.21	-0.29	
	(0.11)	(0.16)	(0.21)	(0.21)	(0.20)	
Number of Plans		0.15**		0.14**	0.13**	0.05**
		(0.02)		(0.02)	(0.02)	(0.02)
Number of PDPs		-0.07**		-0.05**	-0.05**	-0.03
		(0.01)		(0.01)	(0.01)	(0.02)
MA OOPC (10's)			-0.25**	-0.20**	-0.19**	-0.38**
			(0.04)	(0.03)	(0.03)	(0.05)
MA Premium (10's)			-0.40**	-0.31**	-0.29**	-0.49**
			(0.06)	(0.05)	(0.05)	(0.06)
MA Part D premium (10's)			-0.40**	-0.42**	-0.41**	-0.85**
			(0.20)	(0.18)	(0.17)	(0.16)
Part D Plan premiums (10's)					-0.23	0.28
					(0.27)	(0.40)
Medigap Premium (10's)					0.36**	0.20**
					(0.08)	(0.07)
Medigap OOPC (10's)						-0.03
						(0.02)
Part D OOPC (10's)						0.45**
						(0.07)
Observations	16,848	16,848	16,848	16,848	16,848	11,264
County FE	Yes	Yes	Yes	Yes	Yes	Yes
F-test for time break	139.1	32.48	2.813	2.813	1.047	

The outcome variable in all regressions is the percent of 65-year-olds enrolling in Medicare Advantage in a county in a year. Regressions include linear time trends, trend breaks, and county fixed effects. F-tests for time breaks are included bottom line. OOPC=out-of-pocket cost. Standard errors clustered on the county level. Observations on the county-year level. $^{**}p \leq .05$. *All regressions control for the percent of Medicare Advantage in Medicare Advantage employer plans (coefficients not shown). All regressions weighted by number of beneficiaries in a county, averaged across years.*

additional $10 of Medicare Advantage Part C or Part D premiums each reduce the percent of 65-year-olds enrolling by .4 percentage points.

Including both plan availability and premiums/benefits absorbs substantial residual variation in enrollment and reduces upward trends by 50% of the baseline amount (Col. 4). Coefficients also show that plan availability and premiums/ benefits are correlated. Including plan availability reduces effect sizes on out-of-pocket costs and Part C premiums. However, benefits/premiums still have an effect on enrollment that is not fully captured by plan availability.

Columns 5 and 6 show that some dimensions of related markets (Medigap and



Counties are divided into quartiles based on 2006 levels of 65-year-olds' enrollment in Medicare Advantage. The left panel shows trends in the percent of 65-year-olds enrolling, and the right panel shows trends where enrollment is normalized to 2006 levels.

Figure 3.3: Trends in 65-year-olds enrolling in Medicare Advantage, by 2006 quartiles of enrollment

Medicare Part D) affect Medicare Advantage enrollment, but that changes in these markets cannot explain trends. Medigap premiums (Col. 5) significantly affect enrollment, with each additional $10 of Medigap premiums increasing the percent of 65-year-olds who enroll in Medicare Advantage by .36 percentage points. Average Medicare Part D premiums do not affect aggregate enrollment, and neither of these variables absorbs more of the initial upward enrollment trend.

Column 6 adds measures of Part D and Medigap generosity, as measured by out-of-pocket cost. Part D out-of-pocket cost is only available for 2008 and later, which limits sample size and precludes tests for time breaks. However, results suggest that PDPs and Medicare Advantage plans are substitutes. An additional $10 in average PDP out-of-pocket cost increases the percent of 65-year-olds in Medicare Advantage by .43 percentage points. Estimated Medigap out-of-pocket costs have a statistically insignificant effect on enrollment, likely because variation in Medigap generosity is driven almost entirely by changes in the Traditional Medicare benefit.

Results suggest that changes in 65-year-olds' enrollment trends are most plausibly explained by changes in premiums and benefits. However, upwards trends observed between 2006-2009 cannot be fully explained by the characteristics of Medicare Advantage plans or related markets.

Table 3.6 tests how economic variables affect 65-year-olds' enrollment. Column 1 adds county-level controls reflecting economic conditions in the whole population (per capita income, unemployment rates, and the percent of the county population below poverty). Of these variables, only the percent below poverty has a significant effect on enrollment, with each additional percentage point of the population in poverty increasing the percent of 65-year-olds enrolling Medicare Advantage by .18. However, these variables do not absorb the trend break (F=111.4).

Columns 2-4 control for the percent of 65-year-olds who are dual eligible, originally qualified based on disability, and who are observed in Medicare files this year (as measured by the percent of next year's 66-year-olds observed in Medicare in the prior year.) Of these, only the percent of beneficiaries qualified by disability has a statistically significant effect on the percent of 65-year-olds enrolled in Medicare ($\beta = .12$). However, including disability controls does not substantially change enrollment trends.

Columns 5-7 test how Part-A-only enrollment is related to Medicare Advantage enrollment. Beneficiaries must be enrolled in both Medicare Parts A and B to opt into Medicare Advantage, so there is a mechanical, negative relationship between Part-A only enrollment and Medicare Advantage enrollment (Col 5). Cols. 6 and 7 test how excluding Part-A-only beneficiaries affects Medicare Advantage trends. Excluding Part A beneficiaries makes the upward trend and subsequent trend break in Medicare Advantage enrollment look much sharper (Figure 3.4).

Results suggest that, absent the decrease in Part-A-only beneficiaries, the downward trend in 65-year-olds joining Medicare Advantage would have been much sharper ($-.2$ percentage points in unadjusted regressions, with a significant trend break (F=160.8)). Adding controls for plan characteristics in Medicare Advantage, Part D, and Medigap attenuates negative trends. However, the trend break in 2009 is still statistically significant ($F = 27.90$). If Part-A-only enrollment truly reflects changes in employment, results suggest that without the recession, the break in 65-year-olds' Medicare Advantage enrollment might have been sharper.

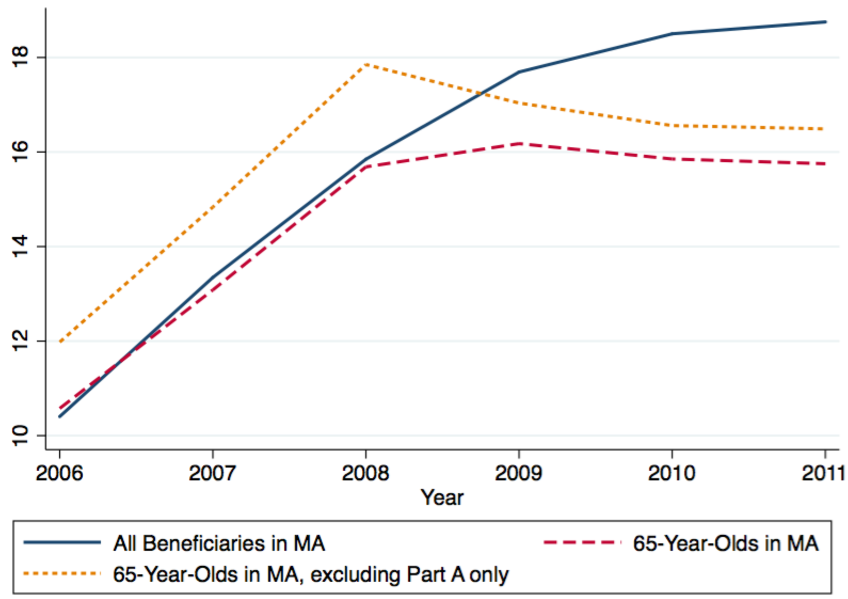


Figure shows 65-year-olds' Medicare Advantage enrollment with and without Part-A-only beneficiaries. 65-year-olds are counted as being in Part A only if they are enrolled in just Part A or out of Medicare for the entire year.

Figure 3.4: Medicare Advantage enrollment and beneficiaries in Part-A only

Table 3.6: Trends in 65-year-olds' enrollment in Medicare Advantage (2006-2011) and Medicare population characteristics

VARIABLES	(1) Trends Controls	(2) Trends % Disabled	(3) Trends % Disabled	(4) Trends % Lead 66	(5) Trends Part-A Only	(6) Trends No Part-A	(7) Trends No Part-A
Trend	2.09**	2.09**	2.07**	2.06**	2.28**	2.80**	1.21**
	(0.16)	(0.16)	(0.17)	(0.19)	(0.17)	(0.20)	(0.30)
Trend 2009 and later	-2.31**	-2.31**	-2.29**	-2.54**	-2.45**	-3.00**	0.44
	(0.16)	(0.16)	(0.16)	(0.23)	(0.16)	(0.20)	(0.45)
Year 2009 or later	0.77**	0.77**	0.71**	0.42	-0.25	-0.73	-2.74**
	(0.32)	(0.32)	(0.32)	(0.33)	(0.34)	(0.40)	(0.43)
% dual eligible		-0.01					
		(0.03)					
% qualified by disability			0.12**				
			(0.03)				
% Medicare next year				0.03			
				(0.08)			
% in Part A only					-0.23**		
					(0.04)		
Per capita income (1000's)	-0.03	-0.03	-0.04	-0.03	-0.04	-0.09	-0.07
	(0.06)	(0.06)	(0.06)	(0.05)	(0.06)	(0.07)	(0.07)
Unemployment rate	0.08	0.08	0.08	0.28**	0.05	0.01	0.06
	(0.08)	(0.08)	(0.08)	(0.08)	(0.08)	(0.09)	(0.09)
% of county below poverty	0.18**	0.18**	0.19**	0.08	0.18**	0.22**	0.15**
	(0.05)	(0.05)	(0.05)	(0.05)	(0.06)	(0.07)	(0.06)
MA OOPC (10's)							-0.21**
							(0.04)
MA Premium (10's)							-0.37**
							(0.06)
MA Part D premium (10's)							-0.35
							(0.22)
Part D Plan premiums (10's)							-0.32
							(0.36)
Medigap Premium (10's)							0.47**
							(0.11)
Number of Plans							0.16**
							(0.03)
Number of PDPs							-0.05**
							(0.01)
Constant				8.54			
				(7.88)			
Observations	16,848	16,848	16,848	11,354	16,848	16,848	16,848
County FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
F-test for time break	111.4	110.9	110	75.81	131.7	160.8	27.90

Outcome variable is the percent of 65-year-olds joining MA in a county in a year. All regressions have county-fixed effects. F-tests capture significant trend breaks. Standard errors clustered on the county level. All regressions control for the percent of Medicare Advantage in Medicare Advantage employer plans (coefficients not shown). Observations on the county-year level. Weighted by number of beneficiaries in a county, averaged across years. $^{**}p \leq .05$

3.7 Conclusions

Trends in Medicare Advantage enrollment matter for policy, as a substantial portion of federal funds are spent on the Medicare Advantage program. Enrollment in Medicare Advantage reached its highest point in the program's history in 2014, despite payment cuts, declining plan availability, and potential reductions in plan generosity. Understanding why enrollment increased despite these changes may be a key part of efficiently financing the program in the future.

To better understand future Medicare Advantage enrollment, this research describes 65-year-olds' enrollment trends between 2006-2011. Enrollment among 65-year-olds may be particularly important for the future of Medicare Advantage, as 65-year-olds are now a larger portion of new Medicare Advantage cohorts than any other age group.(113)

65-year-olds' enrollment in Medicare Advantage in 2006-2009 increased at rates similar to those observed in the broader Medicare population. After 2009, however, 65-year-olds' enrollment diverged and stayed flat. A number of factors are correlated with 65-year-olds' enrollment, including plan availability, Medicare Advantage plan generosity and premiums, characteristics of plans in related markets (Medigap and Medicare Part D), and county economic conditions. Of these variables, changes in Medicare Advantage plan generosity/ premiums most plausibly explain flattening trends in 65-year-olds' enrollment. Additionally, the percent of 65-year-olds enrolled only in Medicare Part A, which may plausibly reflect continued employment, decreased sharply in 2009. This result may suggest that, absent the recession, Medicare Advantage enrollment would have leveled off more sharply. Results suggest that, if benefit generosity continues to decline and the economy recovers, then enrollment growth may slow.

Analysis is limited in a number of ways. First, data end in 2011, before most payment cuts were implemented. Future analysis should incorporate more years of data to test whether slowed enrollment persists. Second, data are on the aggregate, rather than individual level. This allows for analysis of a broader set of markets than those captured in most survey data. However, use of aggregate data limits the degree to which causal statements about enrollee choices can be made. Lastly, many factors relevant to individual enrollment decisions (i.e., employment status, availability of retiree supplemental insurance, health) are unavailable or poorly measured in Medicare data.

There are also several plausible theories explaining upward enrollment trends that cannot be fully tested with the data here. First, Part D implementation occurred during this time period. Though Part D plan premiums and availability do not explain residual trends in enrollment, the implementation of Part D and concurrent changes in Medigap policies might have had a "crowding-in" effect on Medicare Advantage. Future research could explore this hypothesis using more granular data on beneficiary prescription drug use and choices between supplemental options. As no papers to date have explored interactions between Part D and other supplemental programs, research in this area would be a contribution.

Second, retiree health insurance coverage has steadily declined in recent decades. Declines in supplemental coverage cannot be observed in Medicare data. However, if Medicare Advantage and supplemental employer coverage are substitutes, than declines in retiree health insurance may explain upward trends in Medicare Advantage enrollment. Future research using more granular data on supplemental coverage could test this hypothesis.

Analysis here also suggests that trends in employment and income may affect Medicare Advantage take-up. Though upward trends in Medicare Advantage enrollment pre-date the 2008 recession, declines in income and employment may have kept Medicare Advantage enrollment high. The data here have limited information on income and employment, but future research should explore if the recession can plausibly explain sustained Medicare Advantage enrollment.

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Appendix A

Appendix to Chapter 1

A.1 Supplementary analyses

A.1.1 Data and specification assumptions

Baseline regressions were of the following form:

$$Y_{jmt} = \beta_0 + \beta_1 Post_{mt} * S_{m(2009)} + \eta M_{mt} + \theta X_{jt} + \gamma_m + \tau_t + \varepsilon_{jmt} \quad \text{(A.1)}$$

where Y_{jmt} measures premiums or plan generosity for plan j in market m at time t, $S_{m(2009)}$ is the 2009 share of plans cancelled everywhere, and $Post_{mt} = 1$ in 2010 and later, the year the first PFFS plans were cancelled. M_{mt} are county-level controls, X_{jt} are indicators for plan type, γ_m are county fixed effects, and τ_t are time fixed effects.

Several assumptions of this specification warrant further testing. First, the law limiting PFFS networks was passed in July 2008, after insurers had already submitted contracts to CMS for the 2009 calendar year. However, it is still possible that insurers anticipated the law. $S_{m(2009)}$ might be endogenous if insurers preemptively increased or decreased their market shares based on PFFS presence in their county. To insure that the specification in Equation (A.1) is robust to this, I alternately define nationally cancelled PFFS plans shares based on 2008 values ($S_{m(2008)}$). Column 1, Table A.1 shows the results

Table A.1: Robustness checks

VARIABLES	(1) Share 2008 OOPC	(2) Time-Varying Post OOPC	(3) Unbalanced Panel OOPC	(4) Zero-Premium Premium	(5) Zero-Premium OOPC
S_{2008}*Post	8.10				
	(9.60)				
S_{2008}*PPO*Post	31.92**				
	(12.18)				
S_{2008}*PFFS*Post	83.68**				
	(15.13)				
$S_{2009} * Post_{mt}$		-4.95			
		(9.29)			
$S_{2009} * Post_{mt} * PPO$		43.65**			
		(11.22)			
$S_{2009} * Post_{mt} * PFFS$		104.18**			
		(16.40)			
$S_{2009} * Post$			20.72	0.83	24.52**
			(10.81)	(8.55)	(9.36)
S_{2009} * PPO			31.01**		
			(14.26)		
S_{2009} * PFFS			53.72**		
			(15.39)		
$S_{2009}*$0 Premium*Post				8.27	53.44**
				(7.46)	(14.01)
0 Premium				-42.93**	-3.86
				(1.92)	(2.64)
Observations	128,570	128,570	171,915	118,707	107,454
Excludes Exiters	Yes	Yes	Yes	Yes	Yes
Excludes 2007	No	No	No	No	Yes

OOPC=Out-of-pocket cost, less premiums. Higher numbers indicate less generosity. Premium is the plan premium, less any reductions in the Medicare Part B premium. Zero-premium-indicator flags plans with zero premium in prior year. Plans that are new in year t are excluded by construction. Observations on the plan-county-year level. Weighted by average enrollment over all years. All regressions include year and county fixed effects. SEs clustered on the plan-level. $^{**}p \leq .05$.

of this specification check. The coefficient on $S_{m(2008)}$*PPO*Post is 31.91, similar to the coefficient on $S_{m(2009)}$*PPO*Post from baseline results (36.97). Likewise, the coefficient on $S_{m(2008)}$*PFFS*Post is 83.68, similar to the coefficient on $S_{m(2009)}$*PFFS*Post from main results.

Another assumption made in Equation A.1 is that $Post_{mt} = 1$ in 2010 and later, the year the first PFFS plans were cancelled. This specification was chosen to avoid endogeneity due to selective exit. However, it may understate the true magnitude of cancellation's effects, as many counties were not affected by cancellation until 2011 or 2012. Appendix Table A.1 Column 2 shows the results of regressions allowing $Post_{mt} = 1$

to vary across counties based on the year in which plans were first cancelled in the county. The coefficient on $S_{m(2009)}$*PPO*Post is 43.65, statistically similar to those of baseline regressions. The coefficient on $S_{m(2009)}$*PFFS*Post is 104.18, slightly larger than baseline specifications. A larger coefficient may reflect either that insurers selectively timed their exit, or it might reflect the fact that main results are attenuated by setting $Post = 1$ in some counties before cancellation occurs.

The sample used in main analysis was limited to a balanced panel of counties (each county contains plans in all six years of the sample). This limitation was imposed because, in some small counties, only PFFS plans were operating. Their cancellation led to the exit of all plans. Appendix Table A.1 Column 3 replicates results on the full sample (unbalanced panel). Here, the coefficient on HMOs is larger but not different from 0 at conventional levels. The total effect among PPOs (the coefficient on $S_{m(2009)}$*PPO*Post+$S_{m(2009)}$*Post) is equal to 51.73, similar but slightly larger than effects seen in the balanced panel. The total effect among PFFS plans is slightly smaller (20.72+53.72=74.44), but is not statistically different from the total effect from main results. (85.290+3.73=89.02).

Tests in Section 5 showed that insurers modified benefits by more if they charged \$0 in the previous period. This may be because consumers are more sensitive to premium increases when they are initially charged \$0, so insurers modify benefits instead. This result also suggests that insurers charging positive premiums may adjust premiums by more than insurers who charged \$0 premiums. Column 4, Table A.1 shows results from tests of this hypothesis. Regressions are of the form:

$$Premium_{jmt} = \beta_0 + \beta_1 Post_{mt} * S_{m(2009)} + \beta_2 Post_{mt} * S_{m(2009)} * Zero\_Premium_{jm(t-1)} + \beta_3 Zero\_Premium_{jm(t-1)} + \eta M_{mt} + \theta X_j + \gamma_m + \tau_t + \varepsilon_{jmt}$$

where $Zero\_Premium_{jm(t-1)}$ is an indicator equal 1 if the plan charged \$0 in the prior period.

Results do not suggest that insurers charging positive premiums modify their benefits by more than those charging \$0 premiums. However, the zero-premium indicator ($Zero\_Premium_{jm(t-1)}$) may absorb most premium variation, as premiums among all plans are highly autocorrelated ($\rho = .91$). Consistent with this, the sign on the indicator is negative and highly significant, reflecting the fact that plans with \$0 premiums in time $t-1$ had lower premiums in time t.

To include observations from 2007 when testing whether zero-premium insurers modified benefits by more, the assumption was made that insurers charged the same premium in 2007 and 2006. This assumption is reasonable, given the high level of autocorrelation in premiums. However, it is tested by replicating Column 7 from Table 4 in the main text, omitting observations from 2007 (Column 5, Table A.1). Coefficients are very similar to those in the main specification. Results still suggest that insurers modified out-of-pocket costs by more when they charged 0 premiums in the year before.

Further robustness checks were made in Section 5. To test whether linearity drove cancellation's effects, counties were divided by quantile of 2009 share, and indictors for quantile were interacted with S_{2009} and $Post$. Table A.2 shows margins of out-of-pocket costs and premium, by level of cancellation. Figure 4 in the main text displays these results graphically. Plans in counties with $< 1\%$ cancellation and $1-3\%$ cancellation do not significantly adjust benefits, while plans in all other counties do ($p < .001$). Premiums, in contrast, are noisy, and effects seem largely driven by the top quantile of cancellation.

A.1.2 Tests of mechanisms

Sections 6 and 7 tested alternate mechanisms explaining main results. In Section 6, I test whether changes are related to competition in a county by constructing a measure of competition among PFFS plans' substitutes (networked HHI.) Counties are then divided into groups by 2009 levels of HHI and indicators for group are interacted with $S_{2009} * Post$.

Table A.2: Margins for out-of-pocket costs and premiums, by level of cancellation

	(1) Out-of-Pocket Cost Margins	(2) Premium Margins
$< 1\%$	301.1	22.94
	(1.383)	(1.050)
$1-3\%$	301.5	23.42
	(3.093)	(1.375)
$3-12\%$	311.7	24.39
	(2.999)	(1.510)
$12-20\%$	317.4	26.40
	(3.742)	(1.647)
$20-36\%$	318.6	26.01
	(3.534)	(1.795)
$> 36\%$	320.3	31.38
	(4.244)	(2.827)
Observations	121930	121930

Margins from regressing benefits on quantile of Share, interacted with Post. Excludes ever-cancelled PFFS plans. Observations on the plan-county-year level. Weighted by average enrollment over all years. SEs clustered on the plan-level. Includes county and year fixed effects.

Results are displayed graphically in main text (Figures 10 and 11) and in Appendix Table A.3.

Column 1 of Table A.3 shows the results of the regression:

$$OOPC_{jmt} = \beta_0 + \sum_{k=2}^{5} \beta_k Post_{mt} * S_{m(2009)} * D(m \in k) + \eta M_{mt} + \theta X_j + \gamma_m + \tau_t + \varepsilon_{jmt} \quad \text{(A.2)}$$

where $D(m \in k)$ is an indicator equal 1 if a plan is offered in a county in group k. Column 2 shows the results, where $Post_{mt} * S_{m(2009)} * D(m \in k)$ are further interacted by plan type. Results are as described in the main text. Insurers modified benefits the most in counties with less competition at baseline. This pattern is true across all plan types, and is particularly true for counties where there was only one firm offering networked plans at baseline (Group 4). (Overall effect across plans was $\beta_k = 84.23$. For HMOs, the effect was 51.55 and for PPOs, (51.55+17.23=68.78.)

One concern is that counties might have few networked insurers at baseline because

Table A.3: Response to cancellation, by category of networked HHI

VARIABLES	(1) OOPC	(2) OOPC	(3) OOPC	(4) OOPC
	Networked HHI	Networked HHI and Plan Type	Excluding New Plans	Excluding New Plans By Plan Type
Group 1*Share*Post	23.55	-19.01	20.34	-10.04
	(12.69)	(14.25)	(12.18)	(13.49)
Group 2*Share*Post	33.62**	4.43	21.51**	-0.43
	(10.25)	(11.09)	(10.67)	(11.11)
Group 3*Share*Post	36.47**	18.54	31.34**	16.05
	(11.12)	(13.30)	(12.46)	(15.08)
Group 4*Share*Post	84.28**	51.55**	91.12**	42.68
	(11.02)	(16.79)	(14.89)	(32.08)
Group 5*Share*Post	92.31**	42.79**	104.07**	
	(14.78)	(14.47)	(21.77)	
PPO*Group 1*Share*Post		46.33**		46.15**
		(15.39)		(18.84)
PPO*Group 2*Share*Post		40.67**		34.14**
		(14.11)		(14.59)
PPO*Group 3*Share*Post		12.03		11.96
		(15.68)		(26.82)
PPO*Group 4*Share*Post		17.23		17.95
		(16.11)		(36.25)
PPO*Group 5*Share*Post		42.87**		
		(15.15)		
PFFS*Group 1*Share*Post		144.39**		125.93**
		(24.75)		(23.70)
PFFS*Group 2*Share*Post		69.35**		80.97**
		(18.92)		(21.56)
PFFS*Group 3*Share*Post		56.90**		61.08**
		(19.28)		(20.39)
PFFS*Group 4*Share*Post		39.26**		58.10
		(19.42)		(36.39)
PFFS*Group 5*Share*Post		52.13**		
		(16.77)		
Observations	128,570	128,570	92,356	90,671
R-squared	0.66	0.66	0.67	0.67
Excludes Exiters	Yes	Yes	Yes	Yes
Excludes New Plans	No	No	Yes	No

OOPC=Out-of-pocket cost, less premiums. Higher numbers indicate less generosity. All regressions include year and county fixed-effects. Observations on the plan-county-year level. Weighted by average enrollment over all years. SEs clustered on the plan-level. $^{**}p \leq .05$. *Coefficients from interacting county group with share and plan type. 95% confidence intervals. Counties are divided by baseline HHI without PFFS. Group 1: HHI*$<$ 5031, *Group 2:* 5031 $\leq$*HHI*$<$ 6389, *Group 3:* 6389 $\leq$ *HHI*$<$ 10000, *Group 4: HHI*$=$ 10000, *Group 5: No networked plans at baseline.*

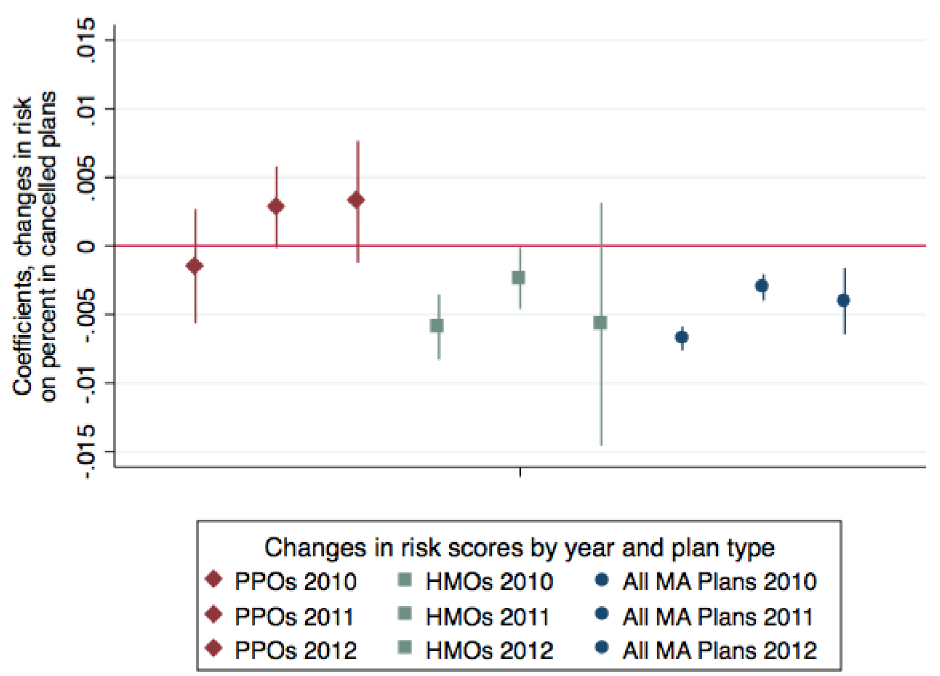
of high provider costs. Insurers might have replaced PFFS plans with PPOs and HMOs in these counties, but offset the costs of building networks with less generous networks. To test this, I use the same approach used in Section 5 and restrict the sample to plans introduced before 2010. This reduces the sample by about a third, but patterns are similar. Without interactions, coefficients are statistically the same as before. In regressions divided by plan type (and omitting Group 5, as it contained no HMOs or PPOs at baseline), point estimates are similar, but results are much noisier.

Analysis presented in Section 7 tested whether changes in benefits were driven by actual or expected changes in risk. To test whether insurers actually enrolled riskier beneficiaries, I regressed aggregate changes in risk scores on the percent of all Medicare beneficiaries in cancelled plans in a county.[1] This tests whether 1) HMOs and PPOs were enrolling worse risks and 2) whether PFFS plans' cancellation changed overall average risk in the market.

Results of these regressions were displayed in Figure 5 in the main text, and in Appendix Table A.4. Columns 1-3 show the overall relationship between cancellation and changes in risk among all MA plans. Cancellation reduced average, county-level risk scores by .001 to .004. Though the magnitude of this change is small, it is statistically significant. Additionally, as .003 is the average change in risk score among all counties in the sample between 2010-2012, it suggests that cancellation led to substantial reductions in overall levels of risk among MA plans. Changes in average PPO plan risk were not significantly different from 0. HMOs also saw a reduction in risk, which would be consistent with them drawing cancelled PFFS plans' enrollees.

Regressions were not weighted by enrollment, because the distribution of Medicare Advantage enrollment and total Medicare population is highly skewed across counties. To test whether this affects conclusions, the same regressions are run, weighting by the

[1] Note that causal regressions use the share of MA enrollment in nationally cancelled plans as an independent variable. This regression uses the share of all Medicare beneficiaries, rather than the share of MA enrollment, and includes all cancelled plans rather than just nationally cancelled plans.



Coefficients from regressing changes in aggregate county-level risk scores on the percent of enrollees in cancelled plans, weighting by the number of Medicare eligibles in a county. Higher numbers indicate greater risk (less health.) Regressions are on the county level. All regressions control for changes in employer plans' shares, as aggregate risk scores are not published separately for employer and non-employer plans.

Figure A.1: Weighted effect of cancellation on risk scores

number of Medicare eligibles in a county, averaged across all years (Figure A.1 and Table A.5.) Weighted coefficients are more strongly negative for all MA plans and for HMOs. This suggests that in larger counties, canceling PFFS plans improved overall risk by more, as sicker enrollees opted to return to traditional Medicare. It also suggests that, in larger counties, HMOs were drawing healthier PFFS enrollees.

I hypothesize that HMOs experienced negative shifts in risk because they were drawing PFFS enrollees, who happened to have lower risk scores than the average HMO enrollee. To test whether the evidence is consistent with this, I also test whether cancellation is associated with shifts in enrollment.[2] I test this by regressing HMO or PPO penetration – the percent of *beneficiaries* in a county in an HMO or PPO – on the

[2]If cancellation reduces HMO risk scores *and* HMO enrollment, this might be evidence that HMOs were generally engaging in strategies to select healthier beneficiaries.

Table A.4: Effect of cancellation on risk scores, by plan type and overall

VARIABLES	(1) PPOs 2010	(2) PPOs 2011	(3) PPOs 2012	(4) HMOs 2010	(5) HMOs 2011	(6) HMOs 2012	(7) All Plans 2010	(8) All Plans 2011	(9) All Plans 2012
% in cancelled plans	-0.003	-0.001	-0.001	-0.004**	-0.002**	0.000	-0.004**	-0.001**	-0.003**
	(0.002)	(0.001)	(0.003)	(0.001)	(0.001)	(0.004)	(0.000)	(0.000)	(0.001)
% in MA-employer plans	0.000	0.000	-0.000	-0.000	-0.000	-0.003	-0.000	-0.000**	-0.001
	(0.000)	(0.000)	(0.001)	(0.000)	(0.000)	(0.001)	(0.000)	(0.000)	(0.000)
Observations	951	1,020	367	1,233	1,005	176	2,004	1,366	430

Results of regressing changes in risk scores on the percent of enrollees in cancelled plans. All Plans is the county-level risk score. HMO and PPO are average, county-level risk scores for HMOs and PPOs, respectively. % of beneficiaries in cancelled plans captures all beneficiaries in the county in any cancelled plan. % of beneficiaries in MA employer plans captures shifts in enrollees in MA plans sponsored by employers. Regressions are unweighted and run separately by year. Observations on the county level. Only counties with positive amounts of cancellation are included in each year. $^{**}p \leq .05$

Table A.5: Effect of cancellation on risk scores, weighted by number of Medicare eligibles in county

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	PPOs	PPO	PPO	HMO	HMO	HMO	All Plans	All Plans	All Plans
VARIABLES	2010	2011	2012	2010	2011	2012	2010	2011	2012
% in cancelled plans	-0.001	0.003	0.003	-0.006**	-0.002**	-0.006	-0.007**	-0.003**	-0.004**
	(0.002)	(0.002)	(0.002)	(0.001)	(0.001)	(0.004)	(0.000)	(0.000)	(0.001)
% in MA-employer plans	0.001	0.000	-0.001	-0.000	-0.000	-0.004**	-0.000	-0.001**	-0.000
	(0.001)	(0.000)	(0.001)	(0.000)	(0.000)	(0.001)	(0.000)	(0.000)	(0.000)
Observations	951	1,020	367	1,233	1,005	176	2,004	1,366	430

Results of regressing changes in risk scores on the percent of enrollees in cancelled plans. All Plans is the county-level risk score. HMO and PPO are average, county-level risk scores for HMOs and PPOs, respectively. % of beneficiaries in cancelled plans captures all beneficiaries in the county in any cancelled plan. % of beneficiaries in employer-sponsored plans captures the percent of the market in an MA-employer plan. Only counties with positive amounts of cancellation are included in each year. Regressions are weighted by the average number of Medicare eligibles in a county, averaged across all years. Observations on the county level. $^{**}p \leq .05$

share of Medicare beneficiaries in all cancelled plans. The specification is:

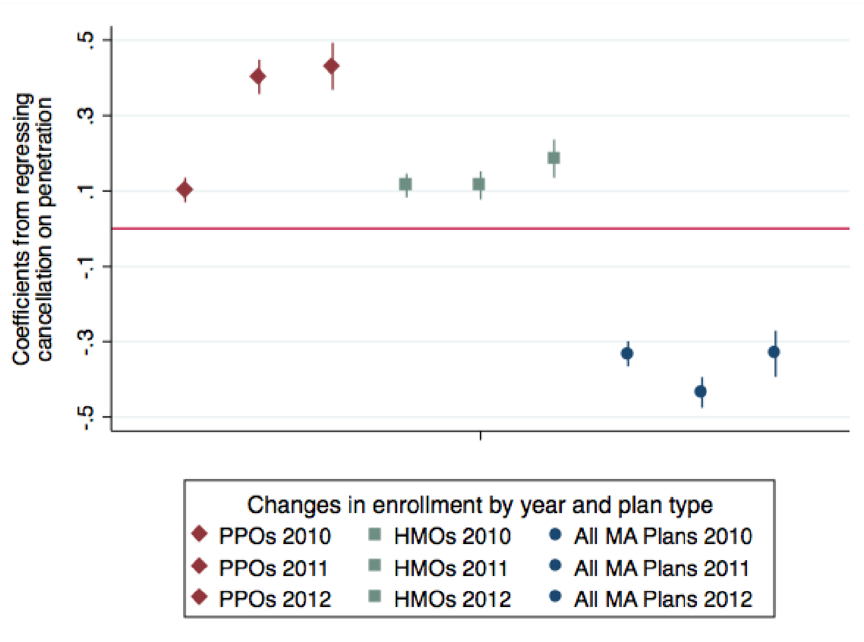
$$\Delta P_{jmt} = \beta_0 + \beta_1 Cancelled\_Share_{m(t-1)} + \varepsilon_{mt(t-1)}$$

where P_{jmt} is the change in penetration in county m for plan type j between time t and $t-1$. β_1 is between 0 and 1 and captures the percent of enrollees in cancelled PFFS plans who join HMOs or PPOs. For completeness, overall MA penetration, or the percent of Medicare beneficiaries in any HMO, PPO, or PFFS plan, is also regressed on cancelled plans' shares. When MA penetration is used as an outcome, β_1 is between -1 and 0 and captures the percent of enrollees in cancelled plans who left MA. Regressions are run separately by year and include only counties in which plans were cancelled. Both unweighted models and models weighted by the number of Medicare eligibles in a county are used.

Figures A.2 and A.3 graphically summarize results of unweighted and weighted regressions, respectively, and Tables A.6 and A.7 display coefficients. Coefficients generally reveal that cancellation increased HMO and PPO enrollment. Each additional percentage point of cancellation increased PPOs market share by between .1 and .4 percentage points, depending on the year and weighting scheme, while HMO enrollment increased between .1 to .27.

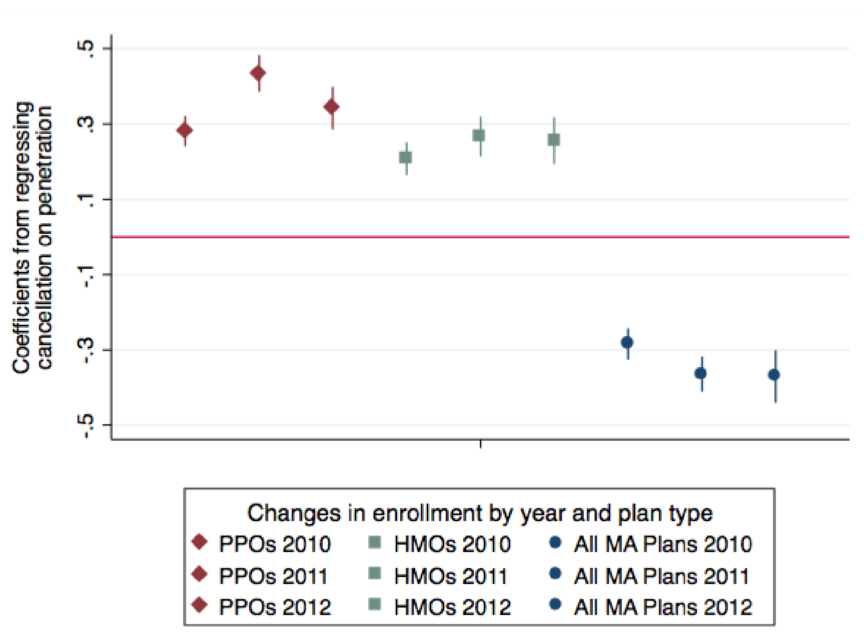
It is not possible to distinguish whether enrollment in HMOs and PPOs increased because they drew cancelled plans' enrollees or because new Medicare beneficiaries were more likely to choose HMOs and PPOs after PFFS plans were removed from the market. However, increased enrollment numbers do not generally support the idea that decreased HMO risk was due to these plans driving away sicker enrollees. The magnitudes of coefficients also generally suggest that enrollees were more likely to choose PPOs than HMOs, consistent with the idea that PPOs were closer substitutes for PFFS plans than HMOs.

Cancellation's overall effect on MA penetration is between -.28 and -.43, suggesting



Coefficients from regressing changes in MA penetration (or HMO/PPO penetration) on all cancelled plans' shares. Regressions are on the county level and unweighted.

Figure A.2: Unweighted effect of cancellation on county enrollment



Coefficients from regressing changes in MA penetration (or HMO/PPO penetration) on all cancelled plans' shares. Regressions are on the county level and weighted by the number of Medicare eligibles in a county.

Figure A.3: Weighted effect of cancellation on county enrollment

Table A.6: Effect of cancellation on county-level enrollment, by plan type and overall

VARIABLES	(1) PPOs 2010	(2) PPOs 2011	(3) PPOs 2012	(4) HMOs 2010	(5) HMOs 2011	(6) HMOs 2012	(7) All Plans 2010	(8) All Plans 2011	(9) All Plans 2012
Total Share	0.10**	0.40**	0.43**	0.11**	0.11**	0.19**	-0.33**	-0.43**	-0.33**
	(0.02)	(0.02)	(0.03)	(0.02)	(0.02)	(0.03)	(0.02)	(0.02)	(0.03)
Observations	2,004	1,366	430	2,004	1,366	430	2,004	1,366	430

Results of regressing changes in enrollment on the share of beneficiaries in cancelled plans. All Plans is the percent of county beneficiaries in all plans. HMO and PPO is the percent of all beneficiaries in HMOs and PPOs, respectively. Share of cancelled plans is the overall share of Medicare beneficiaries in any cancelled plan. Regressions are unweighted and run separately by year. Observations on the county level. Only counties with positive amounts of cancellation are included in each year. $^{**}p \leq .05$

Table A.7: Effect of cancellation on county-level, weighted by number of Medicare eligibles in county

VARIABLES	(1) PPOs 2010	(2) PPOs 2011	(3) PPOs 2012	(4) HMOs 2010	(5) HMOs 2011	(6) HMOs 2012	(7) All Plans 2010	(8) All Plans 2011	(9) All Plans 2012
Total Share	0.28**	0.43**	0.34**	0.21**	0.27**	0.26**	-0.28**	-0.36**	-0.37**
	(0.02)	(0.02)	(0.03)	(0.02)	(0.03)	(0.03)	(0.02)	(0.02)	(0.04)
Observations	2,004	1,366	430	2,004	1,366	430	2,004	1,366	430

Results of regressing changes in enrollment on the share of beneficiaries in cancelled plans. All Plans is the percent of county beneficiaries in all plans. HMO and PPO is the percent of all beneficiaries in HMOs and PPOs, respectively. Share of beneficiaries in MA employer plans captures shifts in enrollees in MA plans sponsored by employers. Regressions are weighted by the average number of Medicare eligibles in a county, averaged across all years. Observations on the county level. Only counties with positive amounts of cancellation are included in each year. $^{**}p \leq .05$

that about one third of cancelled plans' enrollees left MA altogether. Coefficients from weighted regressions are less negative than coefficients from unweighted regressions, (between -.28 and -.37, as opposed to -.33 and -.43), suggesting that PFFS enrollees were more likely to choose another MA plan in larger counties. Differences between weighted and unweighted coefficients using HMO and PPO penetration as an outcome also support this conclusion. Coefficients from weighted regressions are generally more positive than coefficients from unweighted regressions, suggesting that canceled PFFS plans' enrollees switched into other plan types at greater rates in larger counties.

Appendix B

Appendix to Chapter 2

B.1 Supplementary analyses

Table B.1 presents summary statistics for control variables. Of the market characteristics listed, only future benchmark cuts do not differ significantly between markets in which insurers cancelled or kept plans.

Many differences are in intuitive directions. Insurers kept plans in counties with lower service intensity (measured by normalized TM costs), better health among competitors' enrollees (measured by MA county risk), higher per capita income, less of the population in poverty, and lower unemployment. The plans insurers cancelled were much larger (137 enrollees vs. 55), older, and less risky. They also had slightly more generous benefits, as measured by out-of-pocket cost.

Other differences were in less intuitive directions. Insurers were more likely to leave counties with higher benchmark payments, higher risk scores among Traditional Medicare enrollees, and smaller populations over 65. Cancelled plans also had lower premiums. However, with the exception of benchmark payments and Traditional Medicare risk, these variables do not significantly affect the likelihood of exit in regressions with a full set of controls. After controlling for other differences, the sign on benchmark payments is in the expected direction (i.e., insurers were less likely to leave counties with generous payments). Even after adjustments, insurers were more likely to stay in counties with higher Traditional Medicare risk scores. Potential reasons for this are discussed in Section 2.5.

Table B.2 presents results of several robustness checks. As data is missing in provider market variables, regressions are fit in subsamples omitting observations with missing vertical integration data (n=1359, Column 1), missing hospital data (n=599, Column 2), and missing physician data (n=142, Column 3). Column 4 excludes small plans ($n < 25$), as insurers may have cancelled these to reduce administrative burden. Column 5 tests an alternate measure of hospital HHI, constructed based on shares of beds, not admissions. Column 6 tests an alternate version of vertical integration, that captures the percent of doctors billing in the top quartile of vertical integration. Results are broadly similar across models.

Table B.1: Plan-county characteristics

	Stayers	Leavers
Medicare Market		
Medicare benchmark	793.15	801.63
	(60.76)	(73.74)
Future benchmark cuts	22.52	22.27
	(15.57)	(16.76)
Normalized TM costs	626.49	641.21
	(81.38)	(98.99)
Traditional Medicare risk	0.96	0.95
	(0.06)	(0.07)
MA county risk score	0.91	0.92
	(0.09)	(0.09)
Plan Characteristics		
Number of enrollees in plan (100's)	137.33	55.12
	(251.14)	(86.13)
Out-of-pocket cost, no premiums	306.30	310.14
	(23.35)	(20.89)
Premium	24.24	27.23
	(30.46)	(31.78)
Contract age (years)	5.22	4.91
	(1.83)	(1.98)
Plan-level risk	0.92	0.94
	(0.10)	(0.11)
Controls		
Population below poverty (%)	14.75	15.35
	(4.94)	(5.28)
Per capita income (1000's)	33.90	33.56
	(6.41)	(6.74)
County Unemployment	9.14	9.48
	(2.65)	(2.84)
Population 65+ (1000's)	21	23
	(35)	(50)
Observations	2482	3354

Characteristics for plan-county observations where insurer exited (leavers) and did not exit (stayers).

Table B.2: Robustness checks

VARIABLES	(1) Missing Data No Vertical	(2) Missing Data No Hospital	(3) Missing Data No Physician	(4) Enroll 25+	(5) Alt.Measures HHI Bed	(6) Alt.Measures Alt. Vert
Insurer share of all Medicare (%)	-0.28**	-0.33**	-0.27**	-0.23**	-0.26**	-0.26**
	(0.07)	(0.06)	(0.05)	(0.05)	(0.05)	(0.05)
Commerical market share (%)	-0.07**	-0.06**	-0.06**	-0.06**	-0.06**	-0.06**
	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)
Insurer's Share of MA (%)	0.01	0.01	0.01	0.01**	0.01	0.01
	(0.01)	(0.01)	(0.01)	(0.01)	(0.00)	(0.00)
County-level physician HHI (1000's)	0.10**	0.12**	0.10**	0.10**	0.10**	0.10**
	(0.04)	(0.04)	(0.03)	(0.04)	(0.03)	(0.03)
HSA-level admission HHI (1000's)	0.09**	0.10**	0.10**	0.09**		0.10**
	(0.03)	(0.03)	(0.03)	(0.03)		(0.03)
Percent of docs vertically integrated	0.33	-0.05	-0.00	0.31	0.03	
	(0.52)	(0.57)	(0.54)	(0.62)	(0.54)	
Number of enrollees in plan (100's)	-0.16**	-0.17**	-0.17**	-0.19**	-0.18**	-0.18**
	(0.04)	(0.04)	(0.03)	(0.04)	(0.03)	(0.03)
MA penetration (%)	-0.00	-0.00	-0.00	-0.00	-0.00	-0.00
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Medicare benchmark (100's)	-0.60**	-0.52**	-0.50**	-0.48**	-0.48**	-0.49**
	(0.19)	(0.18)	(0.17)	(0.18)	(0.16)	(0.16)
Future Benchmark cuts (+100's)	0.90	0.95	0.82	0.85	0.81	0.83
	(0.65)	(0.50)	(0.49)	(0.55)	(0.46)	(0.46)
Contract age (years)	-0.42**	-0.40**	-0.45**	-0.35	-0.46**	-0.46**
	(0.19)	(0.18)	(0.18)	(0.18)	(0.18)	(0.18)
Normalized FFS costs (100's)	0.58**	0.45**	0.44**	0.30**	0.45**	0.45**
	(0.15)	(0.14)	(0.14)	(0.15)	(0.13)	(0.13)
Standardized FFS risk	-0.46**	-0.38**	-0.36**	-0.29**	-0.36**	-0.36**
	(0.15)	(0.13)	(0.12)	(0.14)	(0.12)	(0.12)
Standardized MA county risk score	0.12	0.13	0.13	0.09	0.13	0.13
	(0.11)	(0.10)	(0.10)	(0.11)	(0.09)	(0.09)
Standardized plan-level risk	0.03	0.04	0.05	0.08	0.06	0.05
	(0.07)	(0.07)	(0.07)	(0.09)	(0.06)	(0.06)
Doctors per 10,000	-0.02**	-0.01**	-0.01**	-0.02**	-0.01**	-0.01**
	(0.01)	(0.00)	(0.00)	(0.01)	(0.00)	(0.00)
County-level Hospital beds per 10,000	-0.00	-0.00	-0.00	0.00	-0.00	-0.00
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
HSA-level hospital bed HHI (1000's)					0.08**	
					(0.03)	
Percent of charges billed HOPD						-0.19
						(0.60)
Observations	4,312	5,072	5,529	3,568	5,671	5,671
Insurer FE	No	Yes	Yes	Yes	Yes	Yes
State FE	No	Yes	Yes	Yes	Yes	Yes
County FE	No	No	No	No	No	No
Log Likelihood	-2042	-2341	-2552	-1701	-2629	-2625
Pseudo R^2	0.297	0.324	0.324	0.312	0.321	0.322

Logit model, standard errors clustered on the county level. Exit=1 if insurers will removed all plans from a market between 2010-2012 and 0 if not. Sample includes PFFS plans offered by insurers who continued to offer PFFS plans in counties where the insurer had no HMO/PPO. Control variables not shown: the number of people in the county over 65, percent of population below poverty, the unemployment rate, per-capita income plan benefits, premiums, plan age, and indicators for missing physician HHI, hospital HHI, or vertical integration data. All variables are 2009 values. Additional stats include log-likelihood and McFadden's R^2. $^{**}p \leq .05$.

B.2 Data appendix

Insurer commercial market share data from Interstudy was matched to Medicare data based on insurer name and county. Names were matched within county and year based on the first “token” word in the insurers’ name. Token words were identified by stripping out common words (e.g., “corporation”, “inc”, “and”, and “the”) and standardizing words that identify many insurers across both data sets (e.g.,“blue cross blue shield” and “healthcare”). Blue Cross Blue Shield plans were recoded based on state names (“Blue Cross Blue Shield of North Carolina” became “BCBS_NC”), and commonly occurring combinations of words were combined into one word (i.e., community health group=community_health_group). Then, names were matched based on county, year, and first token word of the insurers’ name.

This approach is preferred to a more standard distance algorithm in this dataset, because of the frequency of names that create a spurious matches (i.e., “Blue Cross Blue Shield” and “Health Group”.) This words commonly appear in both data sets, and are“closer” in a distance algorithm than words that “should” match.

Tricky cases (i.e., Blue Cross Blue Shield of Kansas and Blue Cross Blue Shield of Kansas City) were resolved using SEC filings and company websites to determine ownership. Acronyms were reviewed by hand to align.[1]

Using this procedure, we were able to match 94% of observations in Medicare data to observations in Interstudy data that are reported as having Medicare enrollment. Match quality iteratively assessed using a string distance algorithm for cleaned variable names. In the final match, only 1% of cleaned names had a string distance greater than 0.

Physician HHI is calculated using office, outpatient, and facility spending from the 20 percent sample of the Medicare Carrier File claims from 2007-2012. Physician practice shares are captured using the percent of allowed charges in a county billed under one

[1]For instance, Tufts HMO plans, a prominent insurer in the Massachusetts area, are coded as "Tufts HMO” in Interstudy and “TAHMO” in Medicare data.

tax identification number. A physician or a physician group must have a tax ID to bill Medicare. Both group and solo physician practices can have tax IDs, and physicians cannot bill under a practice's tax identification number without being part of that practice. However, as a single practice may have multiple tax IDs, this method of calculating physician HHI likely understates the degree of concentration in the market. Practices with fewer than 10 claims in a given year are excluded. Markets with fewer than 10 claims in a year are set to missing, and physician HHI is set to 0. Indicators for missing data are included in all regressions.

Physician-hospital vertical integration measures are constructed using the percent of physician Medicare office, facility, and outpatient charges that are billed in a hospital outpatient department.[2] This measure plausibly captures vertical integration as physicians are only allowed to bill outpatient charges in a hospital setting when that outpatient department is affiliated with the hospital.

Physician-hospital vertical integration measures are defined on the core based statistical area (CBSA).[3] Calculation of vertical integration is limited to CBSAs out of data concerns; there is some noise in the percent of charges billed in a hospital outpatient departments, as physicians often erroneously record the site of service. This noise appears to be a greater proportion of billed charges when the number of physician observations is smaller. When vertical integration is missing, the level of vertical integration is set to 0 and an indicator is constructed to reflect that data are missing.

[2]These relationships are identified using the Medicare carrier and outpatient files, but some ($< 5\%$) of hospital outpatient department claims are misclassified as office-based visits in carrier files. These claims are identified and reclassified by identifying claims with facility fees in outpatient files.

[3]CBSAs are defined by the Office of Management and Budget, and include cities with more than 10,000 people and surrounding counties with significant economic ties.

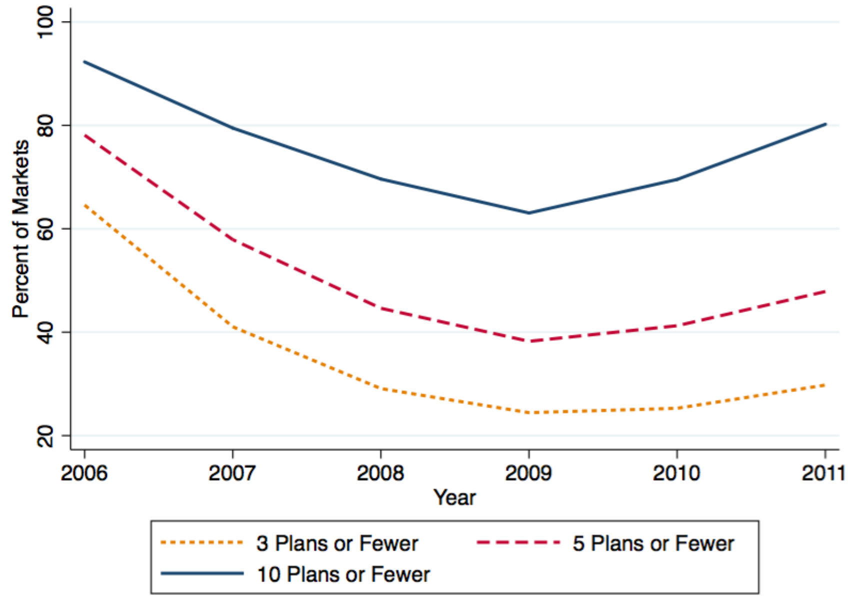
Appendix C

Appendix to Chapter 3

C.1 Supplementary analyses

This section tests several alternate specifications, with a focus on plan availability. One limitation of the approach used in Section 3.6 is that the effect of plan availability is measured using the number of plans. However, the absolute number of plans may be less important to beneficiaries than having a sufficient set of plans from which to choose (i.e., “more than 15”). The percent of markets with very few plans has fallen over time (Figure C.1), and this alone may have increased Medicare Advantage enrollment.

In main results, markets with Medicare Advantage plans were omitted. This restriction might lead to the effects of plan availability being understated, if moving from 0 to 1 plans has a much larger effect than moving from 1 to 2. To address this, a balanced panel of all counties that contain at least one plan in any year is constructed. The number of plans is set to 0 for years where no plans were offered in the county (Table C.1, Cols. 1 and 2). This results in the addition of 821 county-years and makes the upward trend in 65-year-olds joining Medicare Advantage look slightly steeper ($\beta_1 = 2.62$ vs. 2.09 in main regressions). As in prior regressions, controlling for plan availability in this population absorbs some of upward trends ($\beta_1 = 2.25$). However, plan availability has



Percent of markets with more than X plans, over time. Calculated from aggregate enrollment files. Plans with fewer than 11 enrollees are omitted. Markets with 0 plans are included, but a market must have at least one plan sometime during this time period to be in the sample.

Figure C.1: Types of markets, over time

similar marginal effects to those in main results ($\beta_{nPlans} = .15$, $\beta_{nPDPs} = -.07$), and there is still a statistically significant trend break in 2009.

Assuming that the effect of number of plans is linear may also lead to the effects of plan availability being understated. To test whether this is the case, the number of plans and the number of PDPs in the market are transformed by a natural log and added to a regression in the original population of counties (Column 3). Logged variables have much larger marginal effects than linear variables. Each additional percentage point gain in the number of plans increases the percent of 65-year-olds joining Medicare Advantage by a full 2.4 percentage points, while each additional PDP in the county reduces it by 4.4. However, trends in 65-year-olds' Medicare Advantage enrollment look similar, and the trend break is still statistically significant with this specification (F=49.2).

Another limitation of the approach used in Section 3.6 is that plan characteristics (particularly benefit generosity) are measured with error. Increasing benefit generosity might explain more variation in 65-year-olds' enrollment if they were better measured. Alternately, increased enrollment might be tied to unobserved plan characteristics (i.e., service quality, marketing).

To test whether this is the case, county-level benchmark payments to plans are added to regressions to test whether additional payments have any effect on enrollment not captured by benefits, premiums, and number of plans (Col. 4). Adding benchmarks increases the variance on the time trend, making it slightly larger than in past regressions. However, coefficients on benchmarks are insignificant and do not substantially change coefficients on other included variables. This suggests that the additional payments primarily affect enrollment through benefits, premiums, and plan availability, rather than other, unobserved plan characteristics.

Table C.1: Robustness checks

VARIABLES	(1) Balanced Panel Enrollment	(2) Balanced Panel +N Plans	(3) Trends Log Plans	(4) Trends +Benchmark
Trend	2.62**	2.25**	2.28**	1.20**
	(0.12)	(0.17)	(0.19)	(0.42)
Trend 2009 and later	-2.75**	-2.21**	-2.58**	0.03
	(0.15)	(0.24)	(0.28)	(0.52)
Year 2009 or later	1.14**	0.47**	0.93**	-0.52
	(0.12)	(0.17)	(0.15)	(0.45)
Number of Plans		0.15**		0.13**
		(0.02)		(0.02)
Number of PDPs		-0.07**		-0.05**
		(0.01)		(0.01)
Ln(N Medicare Advantage Plans)			2.41**	
			(0.26)	
Ln(N PDPs)			-4.35**	
			(0.53)	
Benchmark				-0.01
				(0.01)
MA OOPC (10's)				-0.18**
				(0.03)
MA Premium (10's)				-0.30**
				(0.05)
MA Part D premium (10's)				-0.40**
				(0.17)
Part D Plan premiums (10's)				-0.24
				(0.28)
Medigap Premium (10's)				0.37**
				(0.08)
Per capita income (1000's)				-0.01
				(0.05)
Unemployment rate				0.12
				(0.07)
Percent of county below poverty				0.12**
				(0.05)
Observations	17,669	17,669	16,848	16,848
County FE	Yes	Yes	Yes	Yes
F-test for time break	177.3	49.20	44.23	1.724

The outcome variable in all regressions is the percent of 65-year-olds enrolling in Medicare Advantage in a county in a year. Regressions with linear time trends, trend breaks, and county fixed effects. F-tests for time breaks in bottom line. OOPC=out-of-pocket cost. Standard errors clustered on the county level. Observations on the county-year level. $^{**}p \leq .05$ *All regressions have county fixed effects and are weighted by number of beneficiaries in a county, averaged across years.*

C.2 Data appendix

Data on the percent of 65-year-olds enrolled in Medicare Advantage, who are dual eligible, who originally qualified for Medicare through disability, or are in only Part A are calculated from beneficiary enrollment records.[1] These data contain complete enrollment records for all beneficiaries in Medicare, including date of birth, county of residence, and whether the enrollee is in Medicare Advantage in a calendar year. A beneficiary is counted as enrolled in Medicare Advantage if they are enrolled in a plan at any point during the year. A beneficiary is treated as dual eligible if they are dual eligible for any month during the year. Beneficiaries are considered to be in only Part A if they are in Part A only or not in Medicare for an entire year.

Out-of-pocket cost (OOPC) data are missing for 2006. To impute missing data, 2006 plan enrollment data is matched to OOPC data for 2007 and 2005. If a plan exists in both 2005 and 2007, and data are missing, then OOPC/premiums are set to the average values for that plan in 2005 and 2007. If a plan exists for only one of those two years, then OOPC/premiums are set to the non-missing value. For the remaining $5-6\%$ of plans with no data, OOPC/premiums were replaced with contract-level averages. After this procedure OOPC/premiums were missing for .18% of plan-county observations.

Data on monthly Medigap premiums are reported in Medicare OOPC estimates. For each year and state, average premiums for plans F and C, which together account for 54% of all Medigap enrollment nationwide, are calculated. Minnesota, Massachusetts, and Wisconsin do not offer the same set of federally standardized plans. For these states, premiums for both basic and extended plans are averaged. State-level data are missing for two states for two years (MO and PA). These were replaced with an average of premiums from the preceding and following years. Underlying data used in Medicare's calculations of Medigap premiums come with a one-year lag. Data are adjusted to be

[1] The Beneficiary Summary File, available under data use agreement from CMS.

contemporaneous.

Medicare also calculates OOPC for each Medigap benefit package. As Medigap plan benefits are standardized, most variation in OOPC comes from changes in Medicare benefit structure. However, as the relative generosity that beneficiaries face across settings is important for choice, Medigap generosity for plans F and C are averaged and added to some regressions. As with Medicare Advantage, data for 2006 Medigap out-of-pocket costs (but not premiums) are missing. These are replaced by an average of 2005 and 2007 out-of-pocket cost estimates for Plans F and C.

County-level average premiums for stand-alone Part D plans are calculated by matching Part D plans from landscape files to enrollment files and averaging the set of plans offered on the county level (unweighted). The number of Part D plans available in a county are calculated from enrollment files. As with Medicare Advantage plans, plans with fewer than 11 enrollees are censored in Medicare data.

Part D out-of-pocket costs were estimated by Medicare starting in 2008. Much of the variation in Part D plans' generosity over time is driven by overall changes in the drug benefit (i.e., changes in donut hole coverage). Estimates are included in regressions alongside Medigap OOPC so that variation in PDP OOPC identifies beneficiary responses to changes in plan generosity, not changes in the overall drug benefit.