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More Insurers Lower Premiums: Evidence from Initial Pricing in the Health Insurance Marketplaces
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ABSTRACT

First-year insurer participation in the Health Insurance Marketplaces (HIMs) established by the Affordable Care Act is limited in many areas of the country. There are 3.9 participants, on (population-weighted) average, in the 395 ratings areas spanning the 34 states with federally facilitated marketplaces (FFMs). Using data on the plans offered in the FFMs, together with predicted market shares for exchange participants (estimated using 2011 insurer-state market shares in the individual insurance market), we study the impact of competition on premiums. We exploit variation in ratings-area-level competition induced by United Healthcare's decision not to participate in any of the FFMs. We estimate that the second-lowest-price silver premium (which is directly linked to federal subsidies) would have decreased by 5.4 percent, on average, had United participated. If all insurers active in each state's individual insurance market in 2011 had participated in all ratings areas in that state's HIM, we estimate this key premium would be 11.1 percent lower and 2014 federal subsidies would be reduced by $1.7 billion.

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1. Introduction

The Patient Protection and Affordable Care Act (hereafter ACA), passed in March 2010 and upheld by the U.S. Supreme Court in June 2012, introduced dramatic reforms to the health insurance industry. A number of benefit designs were banned, premium variation was limited, and online marketplaces for the purchase of insurance were established in every state. Along with Medicaid expansions and mandates for individuals to purchase and large employers to offer coverage, these marketplaces are a key vehicle for expanding insurance coverage. Federal health insurance subsidies are only available to those who purchase a policy through Health Insurance Marketplaces (HIMs).[1] HIMs are intended to promote competition along "beneficial" dimensions (such as price and quality), while at the same time limiting competition along dimensions thought to be socially undesirable (such as the health of enrollees). Whether the federal health reform affordably expands insurance coverage will depend in no small part on the success of HIMs.

The success of HIMs, in turn, will depend on attracting both consumers and insurers. Competition can only have its salutary effects if there are competitors. Prior to the ACA, health insurance markets were very concentrated. The average state HHI for the individual insurance market was 4,100 in 2011, substantially higher than the Department of Justice's threshold of 2,500 for "highly concentrated."[2] HIMs were designed to lower barriers to entry into the insurance industry. By steering a pool of subsidy-eligible consumers to HIMs and mandating that individuals carry insurance, policymakers hoped to create enough new demand to allow entrants to achieve reasonable scale. HIMs also fulfill the role of "certifying" new entrants, whose plans must satisfy federal standards in order to participate in these regulated marketplaces. This federal stamp of approval serves to increase both consumer and supplier confidence in the quality of entrants, a feat that has proved challenging in recent history. And by displaying products online on a centralized website, HIMs reduce marketing, sales, and administrative costs. In addition, the ACA provided subsidized loans to new, nonprofit insurance co-operatives known as Consumer Operated and Oriented Plans, or CO-OPs.

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[1] Until recently, Health Insurance Marketplaces were known as Exchanges. The two phrases are interchangeable.
[2] Calculated using data from the Center for Consumer Information and Insurance Oversight (CCIIO), described in Section 3.

In spite of these policies, there was limited entry into HIMs during 2014, their first year of operation (and the only year for which data are presently available). As our empirical strategy exploits a national decision pertaining to FFMs, we limit attention to these. The FFMs attracted 54% of insurers that ranked among the top three insurers in each state's individual market in 2011.[3] A number of large national insurers, such as Aetna, Cigna, and Humana, participated in only a limited number of HIMs. As we discuss in detail below, the nation's largest insurer, UnitedHealthcare (hereafter United), did not participate in *any* of the FFMs. There were some new entrants, however: 36 insurers participated in FFMs in states in which they did not operate in the individual market in 2011.[4] Of these 36, 13 were CO-OPs.[5]

The combination of concentrated pre-exchange markets, substantial nonparticipation in the exchanges, and limited entry imply highly concentrated marketplaces. **Figure 1** gives the population-weighted distribution of insurers across the 395 federally delineated ratings areas in the 34 FFM states. Seven percent of the population lives in areas with only one insurance option, and about half live in areas with three or fewer options. Across the FFM markets, on a population-weighted basis, there are on average 3.9 insurers per market, with 2.9 incumbents (i.e., insurers who are not new to the individual market), 0.3 CO-OP entrants and 0.7 non-CO-OP entrants.

In this study, we explore the effect of insurer participation in HIMs on 2014 premiums. Prior empirical research finds that insurer consolidation in recent years has increased premiums for large employer-sponsored plans (Dafny, Duggan, and Ramanarayanan 2012). The degree and nature of competition, and hence the quantitative relationship between market structure and price, may be different in the HIMs.[6] On one hand, HIMs standardize some plan features and facilitate plan comparisons, potentially strengthening price competition. These features are intended to create a more Bertrand-like pricing environment, which can result in low markups with as few as two insurers. On the other hand, the transparent display of nonstandardized plan

[3] That is, of the 102 top-three insurers in the 34 FFMs, entry into the relevant state HIM occurred 55 times.
[4] Based on 2011 data, 33 of these 36 had not previously offered individual insurance in any FFM state. Most "entrants" to the individual market are insurers who previously provided Medicaid-managed care in a given state.
[5] There are 13 new, federally sponsored CO-OPs operating in 13 of the 34 FFMs. In 2014, each CO-OP operated in only one state, with three exceptions. . First, "CO-OPortunity Health" operates in both Iowa and Nebraska. Second, "Health CO-OPerative SCW" *and* "Common Ground Healthcare CO-OPerative" operate in Wisconsin.
[6] There are a number of additional reasons why extrapolating from Dafny, Duggan, and Ramanarayanan (2012) to our scenario is difficult. For example, they study the large-group market, and the initial level of concentration in these markets during their study period is significantly lower.

features (e.g., provider networks) and (eventually) plan quality may spur product differentiation, higher markups, and potentially higher average prices. In addition, the existence of subsidies may dampen the price elasticity of some buyers, tempering the relationship between competition and price (and implying more competitors are needed, *ceteris paribus*, to generate competitive outcomes).

Our empirical work focuses on the second-lowest-price silver plan (hereafter *2LPS*) within a market. Federal subsidies are linked to the 2LPS, and past evidence suggests that the lower tail of the premium distribution may be particularly important to consumers (Ericson and Starc 2012a). The *2LPS* exhibits a substantial amount of variation nationally: among FFMs, the 90th percentile of *2LPS* is 45% higher than the 10th percentile.

Existing cross-sectional studies suggest that HIMs with more insurers have lower premiums.[7] **Figure 2** illustrates that exchanges with more participants generally have lower *2LPS.* The graph shows the distribution of 2LPS premiums by the number of exchange participants, along with a fitted line from a univariate regression; while there is substantial variation around the line, the slope is negative (correlation coefficient = -0.35).

This fact admits many interpretations For example, insurers may prefer to participate in geographic markets where medical costs are lower. To mitigate such endogeneity concerns, we exploit United's decision to uniformly avoid all 34 FFMs as a source of quasi-experimental variation in ex-post marketplace concentration. United's nonparticipation differentially affected the competitive environment across markets, owing to its pre-ACA price and product characteristics as well as the participation decisions, prices, and product characteristics of rivals. It is also a policy-relevant source of variation, as insurers similar to United are likely marginal nonparticipants: if expected profits for insurers increase, large national players who shunned the exchanges are likely to enter.

We construct a measure of the change in market concentration resulting from United's decision to avoid FFM markets. We then model the 2LPS premium across rating areas as a function of this measure and find that premiums are highest in markets where United's participation would have most reduced concentration. Our findings are robust to a wide variety of specification checks.

[7]See, for example, http://aspe.hhs.gov/health/reports/2013/MarketplacePremiums/ib_marketplace_premiums.cfm.

We estimate that the population-weighted average *2LPS* premium would have been 5.4% lower had United entered all markets. If all insurers present in a state's individual market in 2011 had entered the exchanges, we estimate FFM premiums would have been 11.1% lower. We also find that markets with CO-OPs have lower premiums, although we caution against a causal interpretation of this association due to the potential endogeneity of CO-OP locations.

These results suggest that additional competitors can have a large impact on premiums and federal subsidies for HIM plans. Spiro and Gruber (2013) estimate that each 1% reduction in 2LPS reduces federal subsidies by 1.25%. Back-of-the-envelope calculations imply that attracting all incumbents to insurance markets would save an estimated $1.7 billion in federal subsidies in 2014, and $105.2 billion over the 2014-2023 ten-year horizon, under the (admittedly strong) assumptions that our findings are generalizable to state-based exchanges and that market structures do not change.

The remainder of the paper proceeds as follows. Section 2 provides background on the health insurance marketplaces, United's nonparticipation decision, and prior research on competition among health insurers. Section 3 describes the construction of our dataset and discusses summary statistics. Section 4 presents the main analysis. Section 5 provides a falsification check of the results by examining the relationship between pre-exchange individual market premiums and the instrument for exchange HHI. We also discuss robustness of the findings to alternative specifications. Section 6 concludes.

2. Background

2.1 Health Insurance Marketplaces

HIMs are regulated online marketplaces for the purchase of health insurance. In this paper, we study HIMs for individual policies.[8] The ACA gave states three options with respect to the development of their exchanges: (1) design and manage their own (so-called "state-based" exchanges)—selected by 16 states and DC; (2) let the federal government design and operate the exchange—selected by 27 states; (3) pursue a hybrid approach ("state–federal partnership" exchange)—selected by 7 states. Options (1) and (2) together comprise the federally-facilitated

[8] HIMs for small-group policies exist—and are known by the acronym SHOP for the Small Health Options Program—but as of this writing they do not yet comply with many requirements included in the ACA. Premium and other data on SHOP plans are not readily available.

marketplaces (FFMs). All HIMs became available as of October 1st, 2013 for individuals to purchase coverage effective in January 2014.

The structure of the HIM in every state is the same. There are five tiers of products available to at least some consumers: a "catastrophic" high-deductible plan offered to those below age 30;[9] a bronze plan with an actuarial value of 0.6; a silver plan with an actuarial value of 0.7; a gold plan with an actuarial value of 0.8; and a platinum plan with an actuarial value of 0.9. All products sold on or off the exchange in the individual and small-group markets must conform to one of these tiers. In addition, all plans in these markets must satisfy federal standards regarding "essential health benefits," which includes coverage of a specified set of services, restrictions on benefits limitations (such as annual spending limits), and a maximum out-of-pocket exposure for enrollees of $6,350 (single)/$12,700 (family).

Subject to this standardization, insurers have wide latitude to design their products in almost all states. In particular, plans can pick any plan design that is within 2% of the actuarial-value target, as long as essential benefits are covered. Plans may therefore adjust features of patient out-of-pocket costs in any way that satisfies that standard. Plans are also free to compete on network design, subject to broad restrictions on network adequacy. The variation across plans is meaningful. Indeed, even on the Massachusetts exchange, which was established prior to the ACA and standardized benefits to a greater degree than required by the ACA, the most expensive plan within a standardized benefits tier (and for a specific zipcode–age combination) was 50% more expensive than the cheapest plan (Ericson and Starc 2013b).

Plans on the HIMs set their own prices. While there is no explicit price regulation, there is regulation on the plan Medical Loss Ratio (MLR), the ratio of medical benefits paid out to premiums collected. MLRs must exceed 80% in the individual market and 85% in the small-group market, which places limits on the ability of firms to make large profits.

Individuals in the exchange will in most cases be purchasing insurance products using a federal tax subsidy. The ACA provides that individuals between 100% and 400% of the federal poverty line may access tax credits to offset some of their premiums. These tax credits offset the difference between premiums and a sliding-scale percentage of income, beginning at 2% of income at 100% of poverty and rising to 9.5% of income from 300–400% of poverty. In some

[9] The catastrophic plan is also available to individuals who do not have the option to purchase insurance below the mandate affordability threshold of 8% of income.

states, a federally funded Medicaid expansion covers all those below 133% of poverty, so exchange participation starts at that higher level; in states without Medicaid expansions, exchange participation begins at 100% of poverty. An estimated 4.8 million individuals have income below the federal poverty line and are ineligible for subsidies and Medicaid.[10]

2.2 United's Nonparticipation

A standard difficulty with any study assessing the impact of market concentration on price is the endogeneity of market participation and market shares. In this setting, one concern with regressing exchange premiums on market concentration arises from the possibility that participation decisions (whether by incumbents or *de novo* entrants) may have been affected by expectations about market prices. Many of the large national insurers, such as Aetna, Humana, and Cigna selectively entered the exchanges. For example, Aetna entered 16 of 34 FFMs.[11]

One exception is United, the nation's largest commercial insurer. Once a midsize regional insurer, United now has a national footprint, achieved largely through acquisitions.[12] Its market share varies widely across states, with no consistent geographic pattern. In the individual insurance market, these shares range from < 1% in Montana, North and South Dakota, New Hampshire, Maine, and Utah to over 20% in South Carolina, Missouri, West Virginia, and Arizona.

The variation in United's pre-exchange market position implies that its blanket nonparticipation decision (discussed below) differentially affected the competitive landscape of individual markets. United's decision not to enter could affect *2LPS* through two mechanisms (1) a "direct effect" arising from the possibility that United could have offered one of the two lowest-priced silver plans in a given market; and/or (2) an "indirect effect" due to rivals' strategically lowering their premiums to compete with United. We expect both effects to be

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[10] Source: "The Coverage Gap: Uninsured Poor Adults in States that Do Not Expand Medicaid ," *Kaiser Family Foundation Issue Brief,* March 2014. http://kaiserfamilyfoundation.files.wordpress.com/2014/04/8505-the-coverage-gap_uninsured-poor-adults-in-states-that-do-not-expand-medicaid.pdf

[11] These entry decisions are nonrandom; for example, Aetna's pre-exchange individual market share (per 2011 CCIO data, described below) was more than twice as high in the markets it entered compared to those it did not. Note that Aetna participates on seven exchanges using the Aetna brand name. In most other exchanges, it offers plans under the brand name of Coventry, which it acquired in 2013.

[12] Major plan acquisitions in the past decade include Oxford and Mid Atlantic Medical Services (MAMSI) in 2004, PacifiCare in 2005, Sierra in 2008, and parts of HealthNet in 2009.

larger in areas where United would have been a more significant competitor on the exchanges. In areas where United had higher pre-exchange market share in the individual insurance market, we can infer that its combination of price and product attributes was relatively attractive. Thus, its decision to stay out of the market ought to have softened competition more considerably in these markets (the indirect effect). If United's price tended to be on the low side in these markets as well, all else equal we would also expect the direct effect to be larger in these areas. Our data on pre-exchange individual market premiums (described in Section 3 below) confirm that United's relative rates are lower in states in which they have a larger presence.[13]

Note that if United's decision not to participate in a market provoked others who would not otherwise have participated to do so—and if this is particularly likely where United had high pre-exchange share because the market opportunity is more substantial—then our estimated effects will be downward-biased. Given the long application process associated with participating in the first-generation FFMs, we believe this bias is likely to be small. For the same reason, the indirect effect may also be low in the first year of the exchange operations, as rivals may not have had ample time to adjust their prices in light of United's nonparticipation decision.

Per the Centers for Medicare and Medicaid Services (CMS), insurers had to submit their plan designs by the end of March 2013 and prices by May 3rd.[14] However, there was likely some flexibility to adjust prices after that deadline, as in late June Kathleen Sebelius, secretary of the Department of Health and Human Services (which oversees CMS), stated that rates were not yet finalized.[15]

The first public proclamation of limited participation by United appeared in January 2013, when the Wall Street Journal reported that United was "expected to participate in 10 to 25 … marketplaces … out of … 100."[16] Given this total incorporates the small-business exchanges (SHOPs), the implication is that United was expected to participate in 5–13 individual market state exchanges. The article further quotes United's CEO as stating, "[United's] level of interest

[13] Specifically, United's relative price position (as measured by where its premium per member falls in the within-state premium distribution) was lower in states where it had greater pre-exchange share.

[14] http://www.cms.gov/CCIIO/Resources/Fact-Sheets-and-FAQs/Downloads/marketplace-timeline-narrative.pdf.

[15] Sebelius asserted, "We will be negotiating rates across the country." While HHS lacks the authority to "actively negotiate" with plans (i.e., exclude plans if their rates are too high), HHS may have had other levers to negotiate with insurers, and insurers would likely have been free to revise premiums downward at this point. http://capsules.kaiserhealthnews.org/index.php/2013/06/sebelius-administration-is-negotiating-rates-in-federal-exchanges/.

[16] http://online.wsj.com/news/articles/SB10001424127887324468104578247332079234240.

in exchanges will be driven by how we assess each local market—how the exchange and rules are set up state by state." This statement foreshadows United's blanket decision to stay out of all the FFMs, which had uniform regulations. It is therefore possible that some insurers accurately predicted United's nonparticipation in at least some states at this time.

On April 18th, 2013, a couple of weeks before HHS's May 3rd Initial Qualified Health Plan Submission Deadline, United's CEO reiterated: "We will be very selective. … [We] do not believe exchanges will be a significant factor … in our 2014 commercial market outlook." Given this statement occurred after the "participation deadline" of March 31 and before prices were finalized, this later announcement could have influenced pricing (the indirect effect). Note, however, that even if rivals did not attempt to predict and incorporate United's decisions into their pricing decisions, the direct effect would still operate as a mechanism to lower price.

2.3 Prior Research

2.3.1 Insurance Market Competition

This study builds on existing research on competition among private insurers. A number of recent studies show that imperfect competition in various U.S. health insurance markets leads to higher premiums. These include Starc (forthcoming) for the Medigap market, Starc and Ericson (2012b) for the Massachusetts health insurance exchange, and Dafny et al. (2012) for the large employer market. Starc predicts that entry of a single additional large insurer would reduce the enrollment-weighted Medigap premium by 21 percent and expand the market by 50 percent. Starc and Ericson build a model of consumer demand using enrollment data from the Massachusetts health insurance exchange and simulate optimal insurer pricing under alternative competitive scenarios. They find that pricing exceeds the levels predicted under perfect competition. Dafny et al. (2012) quantify the impact of market concentration (as measured by HHI) on premium growth in the large group segment, instrumenting for concentration using the predicted change in local market HHI generated by a large, national merger in 1999. This merger had varying impacts on local markets owing to differences in the market shares and geographic overlap of the merging firms. They estimate premiums in the average market were approximately seven percentage points higher by 2007 due to increases in local concentration between 1998 and 2006.

Our instrument is similar in spirit to that used by Dafny et al. We exploit variation in the local impact of United's national nonparticipation decision to identify the effect of exchange market concentration on premiums. Whereas Dafny et al. study the effect of HHI on premium growth, we have only one year of data and hence focus on premium levels. Our point estimates are roughly one-third the size of those reported by Dafny et al. Because their estimate captures the cumulative impact of changes in HHI over time, it is unsurprising to find a smaller single-year effect.

2.3.2 Exchange Research

As HIMs are a recent phenomenon, there is a limited amount of relevant prior research. We briefly discuss the literatures on the two most direct predecessors to HIMs: the Massachussetts Connector exchange and Medicare Part D.

There are a number of recent papers examining the Massachusetts Health Connector, an exchange established by the 2006 healthcare reforms in Massachusetts. In a series of papers, Ericson and Starc study: (1) how changes in the degree of plan standardization required by the exchange affected consumer choice, plans offered, and pricing (Ericson and Starc 2013b), (2) what types of plans consumers choose (20% select the cheapest option; Ericson and Starc 2012a), and (3) the interaction between age-specific consumer price elasticities, imperfect competition, and modified community rating (Ericson and Starc 2012b). Hackmann, Kolstad, and Kowalski (2013) report that average costs and premiums per insured individual in Massachusetts decreased following the imposition of the mandate to carry insurance coverage, confirming adverse selection into the state's individual insurance market prior to 2006.

There is a substantial and growing body of literature on Medicare Part D, a marketplace with many similarities to the HIMs. In both settings, the government subsidizes purchases and creates rules to manage how competition among firms takes place. This literature focuses heavily on whether enrollees make good choices, how limitations in consumer decision-making affect firm behavior, and how alternative choice architecture could improve consumer welfare. (See, for example, Abaluck and Gruber 2011, 2013; Ericson *forthcoming*; Ketcham et al. 2012; Kling et al. 2012; Lucarelli, Prince, and Simon 2012; Zhou and Zhang 2012; and Heiss et al. 2012). Overall, the Medicare Part D literature suggests that even with robust entry, poor optimization by enrollees mitigates the salutary effects of competition.

3. Data and Methodology

We draw on a number of sources to create a dataset of plans offered in the 395 ratings areas (across 34 FFMs), along with measures of ratings-area-level market structure and local health spending. Because United's nonparticipation decision was uniform only across FFMs, we limit attention to these. We also construct a dataset of enrollment and premiums at various units of geography, depending on the source.

3.1 Key Dependent and Independent Variables

Data on plans were downloaded from the healthcare.gov website.[17] The plan data contains insurer identifiers, plan metal tier, ratings areas in which a plan is offered, and premiums for a 27-year-old. Our key dependent variable, *2LPS,* is the premium for the second-lowest-price silver plan in a ratings area. Plan premiums for other ages and family structures are a constant percentage of the 27-year-old single premium.[18]

We focus on the 2LPS for two reasons. First, federal subsidies are linked to the *2LPS* in each market (or "ratings area," the geographic markets utilized on the exchanges and described later). According to CBO estimates, 76% of HIM enrollees will receive subsidies in 2020, accounting for $93 billion of the $197 billion projected cost of ACA's coverage expansions.[19] Thus, *2LPS* is tightly linked to the overall costs of the ACA.

Second, there is evidence that the lower tail of the premium distribution may be particularly important to consumers. As noted above, Ericson and Starc (2012a) report that a substantial number of consumers who purchased insurance on the Massachusetts exchange in 2007–2009 selected the least expensive plan. The results from Massachusetts may not generalize to the nation as a whole. However, given the number of plans, and our inability to judge which of these will prove most popular, a measure like the mean or median is less relevant. For completeness, however, we also report results using such measures.

Our key independent variable for measuring competition is HHI, a predicted Herfindahl–Hirschman Index. Because the market is new, we must predict market shares in order to compute a (predicted) HHI. To do so, we match insurers appearing in the FFM data with state-

[17] Source: https://www.healthcare.gov/health-plan-information/.
[18] States had the opportunity to design their own state-specific age curves for defining how premiums would vary by age. None in our sample did so.
[19] http://www.cbo.gov/sites/default/files/cbofiles/attachments/45231-ACA_Estimates.pdf

insurer enrollment data (in the individual insurance market) for 2011. These data are collected and reported by the Center for Consumer Information and Insurance Oversight (CCIIO) for the purpose of enforcing the Minimum Loss Ratio (MLR) regulations.

For insurer *i* in ratings area *m*, we define $share_{im}$ as its share among those insurers who are active within that ratings area in the exchange, under the assumption that insurers split the market proportionally to their ex ante (i.e., 2011) state shares. Based on the limited empirical evidence available, it appears that pre-exchange shares are highly correlated with exchange shares.[20] This methodology gives new entrants a share of zero. (In Section 5.2, we discuss the robustness of our results to alternative share allocations for entrants.) Denoting the set of insurers in market *m* as I_m, we construct $HHI_m = \sum_{i \in I_m} share_{im}$.

Next, we construct ΔHHI, the change in HHI resulting from United's nonparticipation. The predicted share of each insurer had United entered the market is denoted $share_{im}^{wUHC}$, and the predicted HHI is HHI_m^{wUHC}. United's share had it entered the FFMs is $share_{UHCm}^{wUHC}$. Note that for all insurers other than United, $share_{im} = \frac{share_{im}^{wUHC}}{1 - share_{UHCm}^{wUHC}}$. The increase in HHI from United's nonparticipation can then be expressed as:

$$\Delta HHI_m = HHI_m - HHI_m^{wUHC} \tag{1}$$

$$\Delta HHI_m = \sum_{i \in I_m} \left(\frac{share_{im}^{wUHC}}{1 - share_{UHCm}^{wUHC}} \right)^2 - \left(\sum_{i \in I_m} \left(share_{im}^{wUHC}\right)^2 + (share_{UHCm}^{wUHC})^2 \right) \tag{2}$$

The effect of increasing United's share on ΔHHI_m is:

$$\frac{\partial \Delta HHI_m}{\partial share_{UHCm}^{wUHC}} = 2[HHI_m - share_{UHCm}^{wUHC} HHI_m - share_{UHCm}^{wUHC}] \tag{3}$$

[20] Emerging evidence on exchange enrollment suggests that pre-exchange shares are good indicators of exchange shares. The Huffington Post collected enrollment data for eight states (CA, CT, MA, MN, NV, NY, RI, WA). Using their reported data (which excludes some small players) for states other than MA (which had an exchange prior to 2011), we calculated predicted exchange market shares using 2011 CCIO data (and excluding United). Insurers that entered in 2014 but were not present in 2011 are assigned a share of 0 in 2011. Insurers present in 2011 but not participating in the exchanges are excluded. The correlation between our predicted shares and the actual 2014 shares was 0.63. (Data source: http://www.huffingtonpost.com/2014/01/27/health-insurance-obamacare_n_4661164.html.)

This expression shows that, theoretically, United's nonparticipation has a nonmonotonic effect on ΔHHI. If United is very large and its competitors are all small, ΔHHI will decrease in $share_{UHCm}^{wUHC}$ and can even become negative. As a practical matter, ΔHHI in our data is almost always increasing in United's share, and is only negative for one observation. We censor this observation at zero in our main results; dropping it has little impact on the findings.

An alternative to ΔHHI is United's pre-exchange share. The advantage of using HHI is that it captures the relative importance of United's rivals: a 10% United share matters more in a market with just one rival (ΔHHI = 1800) than in a market with, say, 3 equally-sized rivals (each with pre-exchange market share of 30 percent, yielding ΔHHI =533). As a robustness check, however, we also examine results using United's pre-exchange share in place of ΔHHI.

3.2 Additional Controls

We supplement our dataset with a number of controls that may affect healthcare costs, insurance preferences, or the competitive environment in a ratings area: hospital price, share of nonprofit insurers, whether there is a CO-OP, per-capita income, and % Black and % Hispanic. We construct our measure of acute-care hospital prices for non-Medicare patients using 2007–2009 data from the Centers for Medicare & Medicaid Services' Healthcare Cost Report Information System (HCRIS) dataset, following the methodology in Dafny (2009). We seek a measure that reflects commercial prices (i.e., those paid on behalf of privately insured patients), as exchange plans must negotiate prices with providers (as compared to Medicaid and Medicare plans, which set fixed reimbursement amounts). However, the HCRIS data only permit a reasonably accurate estimate of non-Medicare price, defined as net revenue per non-Medicare case-mix-adjusted admission.[21] Thus, there is nontrivial measurement error in the hospital price variable; in particular, hospital price will be understated for hospitals with high Medicaid or uninsured patients. It is not a priori clear, however, that this measurement error will systematically bias the coefficient on HHI or ΔHHI. Nevertheless, we present results excluding all controls to illustrate their impact on the results.

Per Dafny and Ramanarayanan (2012), nonprofits with significant market share charge lower premiums, ceteris paribus, than for-profits. We control for this by including the expected market

[21] We use each hospital's Medicare Case-Mix Index (CMI) to adjust for admissions severity. Critical Access Hospitals and other hospitals not paid under Medicare's Prospective Payment System are excluded from the sample.

share of nonprofit insurers, *Share NFP*, using the same methodology to assign shares that we used to calculate *HHI*. We account separately for the presence of a nonprofit CO-OP using a dummy variable (which varies at the ratings-area level).[22] Both *share NFP* and *CO-OP* are likely to be endogenous. However, the similarity of the results with and without controls tempers concerns that their inclusion biases the effect of interest (i.e., how competition affects price).

We add demographic controls from Census data, specifically the share of the population that is black (*% Black)* and Hispanic (*% Hispanic*). These shares are highly correlated with other covariates we considered (and which are included in robustness tests discussed in Section 5.2), such as percent diabetic and percent uninsured. We report results from weighted regressions, using 2011 ratings-area population estimates as weights.[23] Finally, we control for *per capita income* for 2011 using estimates from the Bureau of Economic Analysis.

3.3 Addressing Limitations with Instrumental Variable

This empirical approach potentially addresses the endogeneity of insurer entry into FFMs, but concerns remain about the correlation between our instrument and the error term in the pricing regression. This correlation can arise in two ways. First, the share of the market that is controlled by United may itself capture underlying market conditions in a way that is reflected in premiums. For example, United may be more able to compete effectively in high-cost insurance markets where its ability to negotiate tough deals with providers can be most valuable.

Second, the variation in ΔHHI comes not only from variation across states in United's individual insurance market share, but also from the decisions of other insurers to participate on the exchanges in each given ratings area. This arises from the fact that United's predicted share for each ratings area is defined as the ratio of its state-level share to the sum of state-level shares of all insurers participating on the exchange in that ratings area. The advantage of this definition (over using United's state-level share) is that it provides a more accurate estimate of United's likely market share in a ratings area. Some insurers are not active in all areas of a state, and this is likely a principal driver of their decision not to participate on the exchanges in these areas. The disadvantage is that participation may also depend on unobserved factors correlated with

[22] Our methodology assigns zero share to entrants, hence the need for a separate variable. In addition, nonprofit CO-OPs are of independent interest given they are new entrants partially funded by government loans.
[23] Unweighted models yield similar results.

exchange premiums. For example, more insurers may wish to participate in ratings areas where exchanges are likely to attract the healthiest enrollees, generating a spurious negative correlation between premiums and concentration.

We address these concerns with the instrument in two ways. To deal with the first concern, we use pre-HIM data on insurance prices to show that there is no pre-existing correlation between our measure and insurance pricing. Ideally, we would like to have a measure of pre-exchange prices for individual policies at the ratings-area level. Unfortunately, these data do not exist. Therefore, we consider four distinct alternatives, each with strengths and limitations. We construct two measures of prices from the 2011 Medical Expenditure Panel Survey Insurance Component (MEPS-IC). The first is the average estimated single enrollee (as opposed to family) premium for private-sector establishments. The MEPS-IC publishes this data for large MSAs and state "residuals" (i.e., non-MSA areas). Therefore, the strength of this measure is that it is available at a relatively fine level of geography; our 34 states contain 79 MEPS-IC markets.[24] However, employer premiums are likely imperfectly correlated with individual-market premiums—in spite of the fact that both reflect local market cost and utilization trends—limiting the value of evidence that employer premiums are uncorrelated with ΔHHI (the falsification exercise). Hence, we also present results using the average estimated single enrollee premium for small employers only, which is more closely linked to the individual market. The limitation of this second measure is that it is only available at the state level, owing to MEPS-IC confidentiality restrictions.

Our third measure of pre-exchange premiums is the average 2011 individual market premium by state, as reported by CCIIO. (This source is also used to calculate our pre-exchange market shares, as described above.) This CCIIO average premium is available for the most relevant market segment (the individual market), but only at the state level. The fourth and final measure of premiums comes from the Large Employer Health Insurance Dataset (LEHID) for 2009, the most recent year for which we have this data. LEHID is a proprietary dataset containing details on the health insurance plans (and associated premiums) offered by a sample of very large employers. The details of this data, as well as its comparability with other sources, are discussed in Dafny (2010). LEHID's main strengths are that it is available at a relatively disaggregated

[24] MSAs and ratings areas do not perfectly match. We assign each ratings area to the MSA with the highest share of the ratings area's population. We follow the same procedure for assigning ratings areas to LEHID markets.

level of geography (our 34 states contain 98 LEHID markets), and that it includes a rich set of variables we can use to control for plan and employee characteristics. However, the data are older and reflect an even more distant market segment from the individual market than the MEPS-IC all-employer sample.[25]

To deal with the second concern, we estimate models using an instrument that is not contingent on HIM participation decisions. We do so by relying on pre-HIM insurer market shares available at a geographic unit finer than the state, which is the level at which the CCIIO data are reported. The only sources of such data are LEHID and InterStudy, a market intelligence company which reports insurer enrollment by MSA.[26] For both sources, we compute insurer shares using purely ex-ante data (i.e., not conditioning upon who actually entered the exchange) for fully insured private insurance plans.[27]

Although the InterStudy and LEHID datasets provide market shares at a finer level of geography than the CCIIO data, we do not rely on either source for our primary analysis for two reasons. First, neither pertains to the individual market, which can have very different insurer market shares than those observed in group markets. Second, the accuracy of the InterStudy data has been questioned in a number of studies, including Capps (2009) and Dafny et al. (2011). The data in both of these studies are now dated, and the newer InterStudy data may be more accurate. However, in light of the historical limitations, the results using InterStudy data should be interpreted with caution.

3.4 Summary Statistics

Table 1 presents population-weighted summary statistics for the 395 ratings areas ("exchange markets") in FFM states. Exchange markets are highly concentrated: the average number of insurers per market is only 3.9. Predicted HHIs are correspondingly very high, with

[25] To improve the precision of our estimates for ΔHHI, we pool LEHID data from 2007–2009 and include both self and fully-insured enrollees when constructing insurer shares. When constructing LEHID premiums, we use only 2009 and only fully-insured enrollees, as the fully insured segment is more similar to the individual market than the self-insured segment. The falsification results are not sensitive to this decision.

[26] We attempted to create additional observations for "state residuals." However, most states have MSAs that cross state boundaries, making it impossible to infer market shares for state residuals. Adding in the state residuals for which this is not a problem does not substantively change the results.

[27] The InterStudy data also contain enrollment for self-insured plans and commercial Medicaid. We examined whether the results are robust to (1) including the self-insured lives, and (2) including Medicaid lives and adding a separate control for Medicaid's share of covered lives. In both cases, the main results remain qualitatively similar, but the coefficient on ΔHHI ceases to be significant at conventional levels.

an average of 7,323, much greater than the DOJ/FTC threshold of 2,500 for "very concentrated." We caution that these HHIs are overstated because our methodology does not allocate share to entrants (who do not appear in the CCIIO data). Nearly 30% of people live in markets with one to two insurers, and half live in markets with three or fewer insurers.[28] Despite the relatively small number of insurers, most ratings areas feature a large number of plans: the mean is 50.9 (including all metal tiers), and 17.2 for silver plans only. The predicted share of nonprofit insurers averages 61%. One in three markets contains a CO-OP.

Figure 3 is a histogram depicting the number of ratings areas with different ranges of ΔHHI. The figure reveals that the predicted impact of United on market concentration is large and varies significantly across markets. The population-weighted mean of ΔHHI is 1,644, which is similar in magnitude to the change in HHI that would result from a transition from three to two evenly sized firms.

Figure 4 presents information on the identities of the firms that offer one of the two lowest silver premiums in exchange markets. The Blues offer the plurality of low-premium exchange plans (57%), which is unsurprising given their high market shares in pre-exchange individual insurance, low prices, and near-universal participation in exchanges.[29] As a first hint that CO-OPs are associated with lower *2LPS*, we find they are often represented in the bottom two. Significantly, for-profit incumbents (i.e., firms like United) offered 20% of these low-priced plans.

4. The Relationship between Market Structure and Prices

4.1 Are Prices Correlated with Market Structure?

We begin by examining whether $2LCS$ is correlated with our endogenous measure of competition. More specifically, we estimate the following equation using data at the ratings-area level:

(4) $$ln(2LCS)_m = \beta HHI_m\ [+X_m\lambda] + \varepsilon_m.$$

HHI_m is our estimate of market competition and X_m is a vector of optional controls, specifically ln(*Hospital Price),* ln(*Per Capita Income), Share NFP, CO-OP, Percent Black* and *Percent*

[28] Because more populous markets tend to have more competitors, the average market is less competitive than the population-weighted numbers suggest. The unweighted average number of insurers and HHI are 2.8 and 8,320, respectively.

[29] Dafny and Ramanarayanan (2012) find evidence suggesting the largest nonprofit Blues have lower prices than comparably sized for-profit Blues.

Hispanic. All observations are weighted by the 2011 ratings-area population. Results are presented in the first two columns of **Table 2**. The first column excludes the control variables, while the second column includes them. In both specifications, greater concentration is positively and significantly correlated with 2LPS. The results imply a one-standard-deviation decrease in HHI (equal to 0.2, per Table 1, which is slightly larger than the mean decrease in HHI that would result if United entered all ratings areas) is associated with a reduction in 2LPS of 5.6–7.2 percent. Of course, given the endogeneity concerns raised above, we are hesitant to place a causal interpretation on the findings.

4.2 Does Competition Have a Causal Effect on Premiums?

Next, we investigate whether competition has a causal effect on premiums. We posit that United's decision not to participate in any of the FFMs is a source of plausibly exogenous variation in exchange market structure. We use ΔHHI_m, as defined in Section 3.1, to instrument for HHI_m. In the following three subsections, we (1) confirm that ΔHHI_m is correlated with HHI_m; (2) show that ΔHHI_m is correlated with 2LPS; and (3) estimate equation 4 using ΔHHI_m as an instrument for HHI_m.

4.2.1 First Stage Model

To evaluate whether ΔHHI_m is indeed predictive of changes in HHI_m, we estimate the following model:

(5) $$HHI_m = \beta \Delta HHI_m \, [+X_m \boldsymbol{\lambda}] + \varepsilon_m$$

Results are presented in the third and fourth columns of **Table 2**, first excluding and then including the controls described above. Across both specifications, changes in ΔHHI_m translate into HHI_m nearly one for one, and the coefficient estimates are highly statistically significant. A number of the controls, such as income, racial composition, and *share NFP*, are significant predictors of *HHI*.

4.2.2 Reduced Form

The reduced-form model relates exchange premiums to the instrument, i.e.,

(6) $$ln(2LCS_m) = \beta \Delta HHI_m \, [+X_m \boldsymbol{\lambda}] + \varepsilon_{m.}$$

The results, presented in the fifth and sixth columns of **Table 2**, imply that prices are higher in markets where United's nonparticipation has a larger effect on predicted market competition. For example, in a market with the median weighted ΔHHI_m, we predict 2LPS would have been

3.6 percent lower.[30] The remaining variables enter with the expected signs. We discuss them further in the following section.

4.2.3 Instrumental Variables

Finally, we estimate the IV regression

(7) $$ln(2LCS_m) = \beta HHI_m\ [+X_m\boldsymbol{\lambda}] + \varepsilon_m,$$

instrumenting for HHI_m with ΔHHI_m. The results, presented in the final two columns of **Table 2**, suggest a meaningful impact of United's nonparticipation on premiums. Given the first-stage coefficient estimates are close to 1, the coefficients are very similar in magnitude to the reduced-form estimates in the adjacent columns. The results are also fairly similar to the OLS results from the first two columns, potentially mitigating endogeneity concerns with the OLS results.

To gauge the magnitude of the results, we examine how premiums would change under two scenarios: (1) United enters all FFM ratings areas; and (2) all incumbent insurers enter all FFM ratings areas in the states in which they offered individual insurance in 2011. Using the coefficient in column (8) (the specification with controls) as our central estimate, we calculate that population-weighted 2LPS would have been 5.4% lower under scenario (1) and 11.1 percent lower under scenario (2).

The estimate for the effect of *CO-OP* on prices is of independent policy interest. *2LPS* is 8.1% lower in markets with CO-OPs; however, as we discuss in Section 5.2 below, CO-OP location may be endogenous. The coefficients on the remaining controls enter with plausible signs and magnitudes. A 1% increase in inpatient hospital prices is associated with a ~0.2% in insurance premiums. This is the same proportion of private healthcare expenditures attributable to inpatient care for privately insured, nonelderly patients.[31]

5. Robustness

5.1 Falsification Exercise

As noted earlier, there are potential concerns about the endogeneity of our instrument. In this section, we present a series of falsification tests designed to examine those concerns.

Our first test documents that ΔHHI is uncorrelated with pre-period premiums, which should allay concerns that the share of the market held by United is correlated with omitted

[30] Using the estimate of 0.293 from column 6 (the specification with controls), together with the median weighted ΔHHI_m of 0.12, yields exp(0.293 *0.12)= 1.036

[31] Figure is from the *2012 Health Care Cost and Utilization Report*, Health Care Cost Institute, September 2013.

determinants of exchange premiums. Pre-period premium data does not exist at the ratings-area level. We therefore use several sources of premium data, some available at the state level, some at roughly the MSA (and MSA residual) level, and one at the LEHID market level. Given the higher level of aggregation (relative to the ratings area), our statistical tests will have lower power, making it harder to reject the null of no correlation between ΔHHI and pre-period premiums. Hence, we compare the results from these regressions with those obtained from estimating our primary reduced-form regression (equation 6) using the same geographical market definitions. **Table 3** presents these results. To increase the comparability of estimates across different dependent and independent variables, we standardize both the dependent and independent variables (by subtracting the mean and dividing by the standard deviation). All specifications are weighted and include the set of controls from prior models.

Column (1) presents results using MEPS MSA-level data on premiums for employer-sponsored plans. Specification 1 (i.e., the top specification) examines whether our reduced-form relationship between $2LCS$ and ΔHHI from Column (6) of Table 2 is present when the data are aggregated to the MSA level. The point estimate is smaller than our estimate from the ratings-area data (i.e., 0.194 vs. 0.336), and statistically significant at $p<0.10$. In contrast, specification 2 (i.e., the bottom specification) contains no evidence of a statistically significant relationship between MEPS employer premiums in the pre-period and ΔHHI. The point estimate is near zero, albeit with large standard errors. The difference between the coefficient estimates in specifications 1 and 2 is not statistically significant at conventional levels.

Columns (2) and (3) repeat the same analysis using state-level pre-period premium data; as discussed above, both the MEPS small-employer premium data and the CCIIO individual insurance data are only available at the state level. Given the high level of aggregation, it is unsurprising that the coefficients from specification 1 (while very similar in magnitude to that in column 1) are not statistically significant at conventional levels in either column (2) or column (3). For both dependent variables, the coefficient estimates from specification 2 are near zero, although with 34 observations our standard errors are quite large and two-sided tests of coefficient equality easily accept the null.

Column (4) repeats the analysis again using LEHID market definitions and premiums. The LEHID specifications include controls for the underlying plan and enrollee characteristics. These include *plan design factor*, which reflects the actuarial value of observed plans in the relevant market, and *demographic factor,* a summary measure capturing characteristics of the insured LEHID population (e.g. family size and gender).[32]

The point estimate in specification 1 is again two-thirds as large as our central estimate and remains significant at $p<0.10$. The coefficient estimate on ΔHHI in specification 2 is again near zero; there is no evidence that ΔHHI is significantly correlated with pre-period LEHID premiums. Here, the coefficients from specifications 1 and 2 are distinguishable at $p=0.16$.

The other major concern raised above was that our instrument conditions on insurers who choose to participate in state exchanges. To address this point, we turn to the LEHID and InterStudy datasets, which allow us to construct measures of ΔHHI at finer levels of geography using purely ex-ante estimates of market share. (Note that the InterStudy data we have do not include premiums.) Column (5) uses LEHID data to construct not only premiums but also a measure of ΔHHI. The point estimate in Column (5) for specification 1 is similar in magnitude (0.18) and statistically indistinguishable from the point estimates in Columns 1–4. However, it is noisily estimated, with a standard error of 0.17. In specification 2, there is no evidence that ΔHHI is correlated with pre-period LEHID premiums.

Finally, the last column of Table 3 uses ΔHHI constructed from InterStudy MSA market shares for all employers. We are only able to estimate specification (1), as we lack InterStudy premium data. The coefficient estimate is similar in magnitude to the other columns and statistically significant at $p<0.10$.

In summary, there is some evidence (albeit weaker and noisier) that ΔHHI is correlated with $2LCS$ even when the data are aggregated to higher levels of geography. By contrast, there is no evidence that ΔHHI is correlated with pre-exchange prices: the point estimates in the falsification exercises in specification 2 are always near 0. However, the coefficients from the exchange and pre-exchange periods are not statistically distinguishable from one another. We also find that substituting our version of ΔHHI with a measure that does not depend on the exchange participation decisions of other insurers has little impact on the reduced-form point estimates.

[32] We also include the market-level shares of plan types (Indemnity, Preferred Provider Organization, Health Maintenance Organization, and Point of Service), as well as the share of plans denoted as "consumer-directed" (i.e., high-deductible plans).

5.2 Robustness Checks

Table 4 presents results of our reduced-form equation using other measures of price: the mean premium across all silver plans offered in a ratings area; the median premium across the silver plans; and the mean of within-insurer mean silver premium (i.e., a mean calculated using one observation per insurer, so as to avoid overweighting insurers with many plans). For this analysis, we exclude the state of Virginia, which has some extreme price outliers.[33] The first column presents the results obtained using this sample and our primary dependent variable, 2LPS. Our conclusions are robust to using these other dependent variables. The point estimates are somewhat smaller, but the differences across specifications are not statistically significant. There are a number of possible causes for the smaller estimated magnitudes obtained using these alternative price measures.

First, the alternative price measures have smaller standard deviations than *2LPS*, suggesting that there is less variation to explain. Second, some of the variation in *2LPS* is related to the sheer number of plans offered in a ratings area. Even if plan prices are in expectation the same (e.g., drawn at random from the same distribution of prices), adding more plans will lower *2LPS* without affecting many other measures of premiums. Third, the coefficients could be interpreted as evidence that the effect of stronger competition is particularly great for plans in the low-priced silver segment. Finally, one may also infer that United's larger impact on 2LPS implies it has a greater direct than indirect impact on exchange premiums.

Table 4 also shows that CO-OPs are significantly related to *2LPS*, but not to other measures of premiums. We take this, along with the relatively large share (in Figure 4) of markets in which CO-OPs are among the two lowest-price firms, as suggestive evidence that CO-OPs are decreasing *2LPS* more through the direct effect (i.e., by being one of the two lowest-price firms in the market) than indirect effect (i.e., they may not—in the first year—have inspired competitors to reduce their premiums.) Due to the potential endogeneity of CO-OP locations, and the lack of an instrument for their presence, the CO-OP results are merely suggestive. Additional research on the impact of CO-OPs would be valuable, as the budget compromise of

[33] Three Virginia insurers (Optima, Aetna, and Innovation Health) have premiums that are extreme outliers. The mean silver premium (for a 27-year-old) across these three insurers is 885. This compares to a mean of 256 for the rest of the country.

January 2013 eliminated funding to support prospective CO-OPs and slashed funds for current CO-OPs.[34]

Our results are robust to a series of other specification choices. In **Appendix Table 1,** we present reduced-form results from models including several additional controls: share of the population located in an urban area, share obese, share diabetic, whether a state is expanding Medicaid, share aged less than 19, share uninsured, and Medicare fee-for-service spending per capita (to capture variation in utilization of healthcare services). These have a minimal impact on the coefficients of interest.[35] In **Appendix Table** 2, we present results from a series of other specification choices: (1) excluding HIMs w/ 5+ insurers; (2) adding dummies for # of exchange insurers; (3) excluding the top and bottom 5% of ΔHHI; (4) excluding the top and bottom 5% of 2LPS; (4) allocating entrants 5% share; (6) including state fixed effects.. The point estimates for the effect of ΔHHI on 2LPS range from 0.151 to 0.669, with most remaining near 0.3 and all statistically significant at the 5% level. Finally, we also estimate a reduced-form model replacing ΔHHI with United's pre-exchange market share. This alternative instrument is also a significant predictor of 2LPS (with $p<0.01$), but the implied effect of United's presence on 2LPS is smaller. The smaller estimated effect is expected given that share is a less accurate indicator of United's effect on market competition than predicted change in HHI.

6. Conclusion

In this study, we evaluate the impact of insurer participation and competition in the FFMs on premiums. We find that exchange premiums are responsive to competition. To contend with the endogeneity of exchange market structure, we exploit the decision by United to forgo participation in the FFMs. This decision differentially impacted markets due to United's pre-exchange market position.

We estimate that the population-weighted average second-lowest silver premium would have been reduced by 5.4% had United entered all markets. If all insurers present in a state had

[34] By December 2012, the federal government had awarded $2 billion in loans, out of $6 billion initially set aside by ACA. In January 2013, Congress eliminated all but ~ $200 million of the remaining funds, and this sum was designated to support the 24 CO-OPs already existing at that time. Thirteen of these CO-OPs offered plans in 2014. *Source:* "Health Policy Brief: The CO-OP Health Insurance Program," Health Affairs, February 28, 2013. Available from: http://healthaffairs.org/healthpolicybriefs/brief_pdfs/healthpolicybrief_87.pdf.

[35] We do not include these controls in our preferred specifications because of multicollinearity issues, and because they absorb degrees of freedom that are particularly scarce in models using state and MSA-level data.

entered all ratings areas in that state's exchange, we predict FFM premiums would have been 11.1% lower. We also find that markets with CO-OPs have lower premiums, and some of this relationship is causal because CO-OPs are often among the two lowest-price silver plans in a market.

The magnitude of the relationship between HHI and exchange premiums is roughly one-third that obtained by Dafny et al. (2012) for the large-employer group market. Although the estimates are not perfectly comparable (in particular, the Dafny et al. estimate reflects the cumulative effect of changes in HHI on premium growth over a few years' time), the similar order of magnitude suggests that the competitive dynamic characterizing early exchange markets is akin to that of the mature, but imperfectly competitive, large-group market. This suggests that exchanges have not (to date) produced a Bertrand-like outcome in which a small number of players can drive price down to cost. Of course, future entry and greater plan standardization may change this assessment.

Given the incipiency of these markets, this study is but a first step in what will surely become a deeper and broader literature on insurance exchanges and the nature and significance of competition among exchange participants. There is substantial room for further research on how competition affects pricing and other outcomes in this market. Future studies will be easier to execute once information about consumer enrollment decisions has been released, and once the market is in longer-term equilibrium. These conditions will allow researchers to apply well-established supply-side methodologies to studying competition on the exchanges. Such research will permit more-nuanced conclusions and recommendations regarding the impact of competition and competition-related policies on various outcomes of interest. Given the large federal role in developing and regulating the exchanges, and in subsidizing the purchase of plans offered on the exchanges, research on how competition affects consumer choice and insurer behavior is of critical importance.

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Table 1: Summary Statistics for Federally Facilitated Marketplaces

Variable	**Mean**	**St Dev**	**Min**	**Max**
Number of Insurers	3.9	2.0	1	9
Number of Plans	50.9	29.6	7	169
Number of Silver Plans	17.2	10.3	2	48
Price of 2nd Lowest Price Silver Plan (2LPS)	214	37	138	395
Under 65 Population	443,830	759,394	7,391	7,612,795
Income per Capita ($)	39,519	7,176	19,049	65,173
Hospital Price ($)	6,597	1,392	3,447	11,906
COOP Present on Exchange	0.33	0.47	0.00	1.00
Share Non-Profit	0.61	0.38	0.00	1.00
% Black	0.16	0.12	0.00	0.75
% Hispanic	0.15	0.15	0.01	0.96
United Market Share (if Participating in Exchange)	0.16	0.12	0.00	0.98
Predicted Exchange HHI (United is not participating)				
HHI	0.73	0.20	0.32	1.00
Predicted Exchange HHI (if United were participating)				
$HHI_{\text{plus United}}$	0.57	0.17	0.24	0.99
Implied ΔHHI				
ΔHHI	0.16	0.11	0.00	0.49

Notes: N=395. The unit of observation is the ratings area. There are 395 ratings areas in the 34 states with federally facilitated marketplaces. Price for the 2nd Lowest Price Silver Plan is the individual premium for a 27 year old. Premiums move proportionally with age. Hospital Price is defined as net revenue per case-mix adjusted discharge, excluding Medicare revenues and discharges, per Dafny (2009). It is constructed using Medicare's HCRIS database. For each ratings area, we use the discharge-weighted average of prices for hospitals located in the area. Share Non-Profit is constructed using the 2011 individual insurance market shares of non-profit insurers participating in the exchange, as reported by CMS' Center for Consumer Information and Insurance Oversight (CCIIO). Summary statistics for variables other than population are reported on a population-weighted basis.

Table 2: Main Results

	Endogenous Regression		First Stage		Reduced Form		Instrumental Variables	
	Dep Var = ln(2LCS)		Dep Var = HHI		Dep Var = ln(2LCS)		Dep Var = ln(2LCS)	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
HHI	0.274***	0.348***					0.260***	0.336***
	(0.040)	(0.041)					(0.079)	(0.083)
ΔHHI			0.954***	0.871***	0.248***	0.293***		
			(0.081)	(0.079)	(0.079)	(0.078)		
ln(Per Capita Income)		0.058		-0.138***		0.009		0.055
		(0.045)		(0.049)		(0.048)		(0.048)
ln(Hospital Price)		0.183***		-0.011		0.179***		0.183***
		(0.038)		(0.041)		(0.041)		(0.038)
COOP in Market		-0.086***		0.022		-0.078***		-0.085***
		(0.017)		(0.019)		(0.019)		(0.018)
Share Non-Profit		-0.077***		0.157***		-0.023		-0.075***
		(0.022)		(0.023)		(0.023)		(0.024)
% Black		0.156**		0.241***		0.240***		0.159**
		(0.064)		(0.069)		(0.068)		(0.066)
% Hispanic		-0.087*		-0.281***		-0.186***		-0.091
		(0.052)		(0.055)		(0.054)		(0.056)
R-sq	0.105	0.290	0.259	0.398	0.025	0.185	0.105	0.289

Notes: N=395. All regressions are weighted by the ratings-area population under 65, as reported by the U.S. Census. The instrument for HHI is ΔHHI.

Standard errors in parentheses. * $p<0.10$, ** $p<0.05$, *** $p<.01$

Table 3: Reduced Form Falsification Exercise

	(1)	(2)	(3)	(4)	(5)	(6)
Source of Pre Period Premiums	MEPS, all firms	MEPS, small firms	CCIIO	LEHID	LEHID	N/A
Source of ΔHHI	CCIIO	CCIIO	CCIIO	CCIIO	LEHID	InterStudy
Specification 1 (Confirmation that main results persist):						
Dep Var = ln(2LCS), studentized						
ΔHHI (Studentized)	0.194*	0.174	0.174	0.188*	0.179	0.149*
	(0.113)	(0.180)	(0.180)	(0.101)	(0.170)	(0.089)
Specification 2 (Falsification):						
Dep Var = ln(Pre Period Premiums), studentized						
ΔHHI (Studentized)	0.029	0.011	0.039	0.018	-0.033	
	(0.113)	(0.147)	(0.159)	(0.061)	(0.104)	
Number of Observations	79	34	34	98	98	248
Unit of Observation	MSA	State	State	LEHID Market	LEHID Market	Ratings Area
P-value for H_0: identical effect of independent variable on both dependent variables	0.30	0.48	0.58	0.16	0.30	

Notes: All regressions are weighted by the ratings-area population under 65, as reported by the U.S. Census.

MEPS MSA definitions break states into MSAs and state residuals (i.e., areas outside the MSAs).

All specifications include the controls in the even columns in Table 2. Regressions with a LEHID dependent variable also control for plan type shares, plan design factor, and demographic factor. The standard errors in column 6 are clustered at the MSA level (196 clusters).

Standard errors in parentheses. * p<0.10, ** p<0.05, *** p<0.01

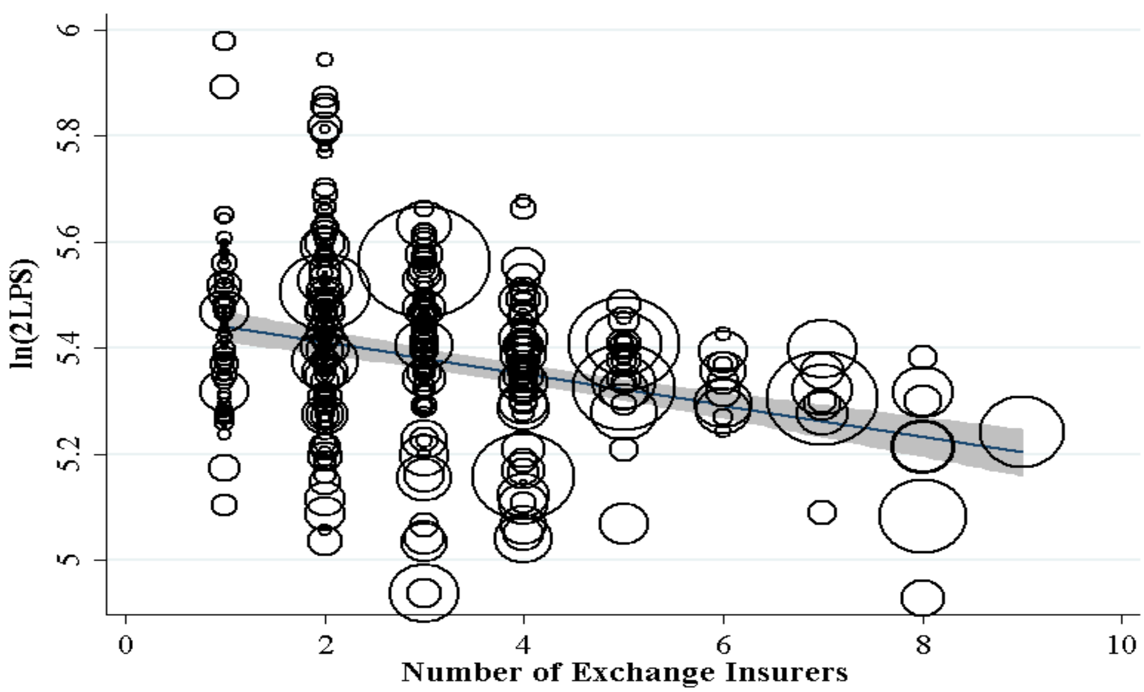
Table 4: Reduced Form
Effect of ΔHHI on ln(Prices)
(Robustness to Alternative Measures of Prices)

	(1)	(2)	(3)	(4)
	Dep Var = ln(2LCS)	Dep Var = ln(Mean Premium)	Dep Var = ln(Median Premium)	Dep Var = ln(Mean of Within-Insurer Mean Premiums)
ΔHHI	0.293***	0.162***	0.182***	0.175***
	(0.079)	(0.057)	(0.062)	(0.055)
ln(Per Capita Income)	-0.013	0.058	0.097**	0.061*
	(0.053)	(0.038)	(0.041)	(0.037)
ln(Hospital Price)	0.197***	0.169***	0.151***	0.161***
	(0.043)	(0.031)	(0.033)	(0.030)
COOP in Market	-0.079***	-0.009	-0.002	-0.000
	(0.019)	(0.014)	(0.015)	(0.013)
Share Non-Profit	-0.018	-0.032*	-0.051***	-0.018
	(0.024)	(0.017)	(0.019)	(0.017)
% Black	0.260***	0.175***	0.218***	0.105**
	(0.070)	(0.050)	(0.055)	(0.049)
% Hispanic	-0.187***	0.070*	0.131***	0.049
	(0.054)	(0.039)	(0.042)	(0.038)
R-sq	0.190	0.173	0.185	0.160

Notes: N=383. All regressions are weighted by the ratings-area population under 65, as reported by the U.S. Census. Samples exclude Virginia, which has very large pricing outliers. When Virginia is included, specifications utilizing a mean premium (i.e., columns 2 and 4) yield statistically insignificant coefficients on ΔHHI.
Standard errors in parentheses * $p<0.10$, ** $p<0.05$, *** $p<.01$

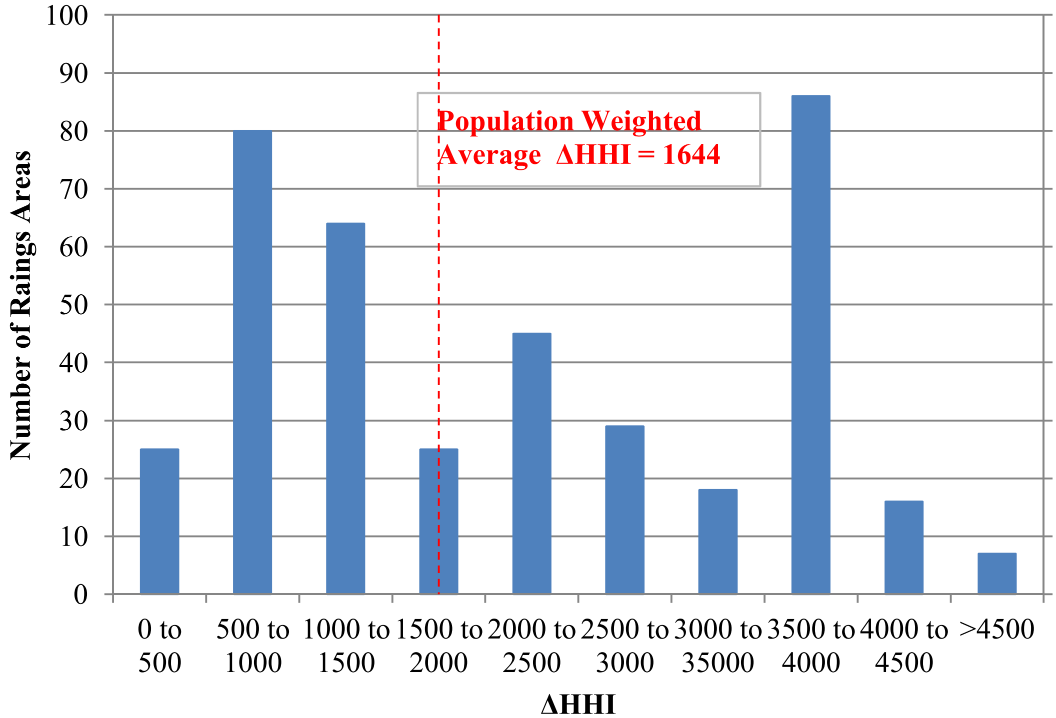
Figure 1: Few Insurers in Many Markets

Figure 2: More Insurers Means Lower Premiums



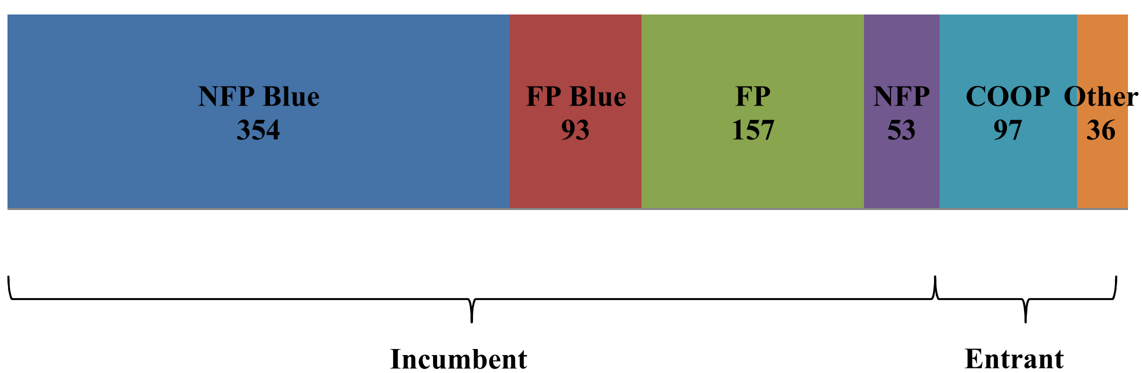
Notes: Scatter plot reflects 395 ratings areas, with circle sizes corresponding to population. Figure also contains weighted regression line and and 95 percent shaded confidence interval.

Figure 3: Predicted Impact of United's Decision



Notes: N=395

Figure 4: Identity of 1st and 2nd Lowest-Priced Silver Insurers, by Category



Notes: N=790

Appendix Table 1: Robustness To Inclusion of Extra Controls

	Endogenous Regression		First Stage		Reduced Form		Instrumental Variables	
	Dep Var = ln(2LCS)		Dep Var = HHI		Dep Var = ln(2LCS)		Dep Var = ln(2LCS)	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
HHI	0.274***	0.330***					0.260***	0.350***
	(0.040)	(0.043)					(0.079)	(0.079)
ΔHHI			0.954***	0.935***	0.248***	0.327***		
			(0.081)	(0.077)	(0.079)	(0.079)		
ln(Per Capita Income)		0.086		0.168***		0.141**		0.082
		(0.064)		(0.064)		(0.067)		(0.064)
ln(Hospital Price)		0.190***		0.036		0.201***		0.188***
		(0.038)		(0.038)		(0.040)		(0.037)
COOP in Market		-0.087***		0.042**		-0.075***		-0.089***
		(0.020)		(0.021)		(0.021)		(0.021)
Share Non-Profit		-0.083***		0.182***		-0.022		-0.085***
		(0.022)		(0.022)		(0.023)		(0.023)
% Black		0.289***		0.312***		0.396***		0.287***
		(0.100)		(0.102)		(0.106)		(0.098)
% Hispanic		-0.323***		-0.033		-0.333***		-0.322***
		(0.097)		(0.099)		(0.103)		(0.096)
% Urban		-0.273***		-0.409***		-0.409***		-0.266***
		(0.065)		(0.064)		(0.066)		(0.068)
% Obese		-1.169**		-1.603***		-1.722***		-1.161**
		(0.486)		(0.502)		(0.521)		(0.478)
% Diabetic		-1.360		4.230***		0.025		-1.454
		(1.269)		(1.272)		(1.318)		(1.288)
Medicaid Expansion		0.018		-0.018		0.011		0.017
		(0.020)		(0.020)		(0.021)		(0.019)
% Less than 19		0.968**		1.103**		1.345***		0.959**
		(0.418)		(0.428)		(0.443)		(0.411)
% Uninsured		0.481*		-0.035		0.458		0.471*
		(0.266)		(0.273)		(0.283)		(0.264)
ln(Medicare FFS)		0.225***		-0.312***		0.122*		0.231***
		(0.072)		(0.071)		(0.074)		(0.073)
R-sq	0.105	0.290	0.259	0.398	0.025	0.185	0.105	0.289

Notes: N=395. All regressions are weighted by the ratings-area population under 65, as reported by the U.S. Census. The instrument for HHI is ΔHHI. The odd columns are reproduced from Table 2. Additional controls are from the most recent year of data available. % Urban is derived from the 2010 Census Urban and Rural Classification data. % Obese and % Diabetic are from the Centers for Disease Control and Prevention for 2010 and 2009, respectively. The state Medicaid expansion indicator is from the Kaiser Family Foundation. % Less than 19 refers to the under-65 population in 2011, and is obtainted from the U.S. Census. % Uninsured is from the Census's 2010 Small Area Health Insurance Estimates and refers to the population under 65. Combined Part A and B spending per Medicare Fee For Service enrollee is from 2011 data reported by the Centers for Medicare and Medicaid Services. Standard errors in parentheses. * p<0.10, ** p<0.05, *** p<.01

Appendix Table 2: Robustness Checks of Reduced Form Model

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
ΔHHI	0.293***	0.296***	0.187**	0.358***	0.151**	0.422***	0.659***
	(0.078)	(0.087)	(0.078)	(0.099)	(0.061)	(0.089)	(0.141)
Explanation	Original	Excludes ratings areas with 5+ firms	Adds dummies for # of firms	Excludes 5% tails of ΔHHI	Excludes 5% tails of ln(2LPS)	Give entrants 5% share	State FEs
N	395	343	395	367	357	395	395
R-sq	0.185	0.166	0.290	0.162	0.120	0.202	0.706

Notes: All regressions are weighted by the ratings-area population under 65, as reported by the U.S. Census.
Standard errors in parentheses. * p<0.10, ** p<0.05, *** p<.01