

# **Insurers' Negotiating Leverage and the External Effects of Medicare Part D\***

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## **Abstract**

By influencing the size and bargaining power of private insurers, public subsidization of private health insurance may project effects beyond the subsidized population. We test for such spillovers by analyzing how increases in insurer size resulting from the implementation of Medicare Part D affected drug prices negotiated in the non-Medicare commercial market. On average, Part D lowered prices for commercial enrollees by 3.7%. The external commercial market savings amount to \$1.5 billion per year, which, if passed to consumers, approximates the internal cost-savings of newly-insured subsidized beneficiaries. If retained by insurers, it corresponds to a 5% average increase in profitability.

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## 1. Introduction

Recent expansions of health insurance coverage rely heavily on public financing of privately provided insurance. The Medicare Modernization Act (MMA) of 2003 established prescription drug benefits for Medicare beneficiaries through Part D, the largest expansion of Medicare in its history. While the government subsidizes roughly 75-percent of premiums under Part D, drug procurement is the domain of private insurers who compete to design, price, and administer insurance policies. Likewise, the landmark Patient Protection and Affordable Care Act (PPACA) extends premium subsidies to many uninsured Americans, but tasks the private-sector with administering health insurance policies for the newly insured. The deep involvement of the private-sector creates a possible linkage between government health insurance subsidies and the private price negotiations typical in the market for medical goods and services. This could have important implications for total health care costs and the distributional effects of such policies.

This study focuses on how publicly financed coverage expansions—by increasing the size and bargaining power of private insurers—may affect prices in the broader commercial marketplace. Such price changes affect those newly insured by the coverage expansions, as well as individuals covered by the commercial insurance market external to the program. We consider the case of the MMA, which may have affected millions of individuals enrolled in the commercial, non-Part D, insurance plans of Part D-participating insurers.

Much of the existing literature on Part D prices has focused on quantifying the success or failure of private firms in efficiently disseminating drug insurance to previously uninsured Medicare beneficiaries. There exists a general consensus that the private sector has improved seniors' access to drug coverage while lowering the drug prices faced by previously uninsured individuals (Lichtenberg & Sun, 2007; Ketcham & Simon, 2008; Yin *et al.*, 2008; Duggan & Scott Morton, 2010). Duggan & Scott Morton (2010) observe that Part D lowers average manufacturer revenues per prescription, because previously uninsured seniors gain access to the discounted drug prices obtained by large private insurers.

There seems to be little doubt that insurance leads to lower prices for consumers who take up insurance. Less well-studied, however, is the possibility that Part D's effects spilled over into the commercially insured population by influencing *unit* drug prices negotiated by insurers. MMA premium subsidies dramatically increased the number of Medicare beneficiaries with

prescription drug coverage and injected millions of new customers into the insurance market. This increase was absorbed primarily by existing insurance firms—not new entrants—so that the MMA generally increased the enrolled population in existing private insurers. In turn, growth in Part D enrollments may affect the bargaining power of private payers in their general pricing negotiations with suppliers. According to standard bargaining theory, if larger buyers generate more surplus per unit for their suppliers, they possess more leverage and will be able to negotiate lower unit prices, and retain a larger share of the total surplus; the reverse holds true when larger buyers generate less surplus per unit (Stole and Zwiebel, 1996; Brooks et al, 1997; Chipty and Snyder, 1999; and Raskovich, 2003).<sup>1</sup>

When insurer size lowers pharmacy prices, for example, we should observe price declines for insured consumers *external* to Part D. Specifically, we should observe larger declines in unit profits earned by pharmacies for the commercial (non-Medicare) claims of Part D-participating insurers that experience larger enrollment increases. Finally, theory predicts the biggest effects on unit profits when pharmacies and insurers—rather than manufacturers—hold more of the total producer surplus. To see why, consider the extreme case where manufacturers possess all the bargaining power, and perfectly competitive pharmacies and insurers never earn any profit. In this case, negotiations between insurers and pharmacies cannot alter the social division of profits, because there are no profits to share between these two parties. This prediction implies stronger effects of insurer size on the retail prices of generic drugs, compared to branded drugs.

The existing literature suggests the potential importance of these effects in health care. Sorensen (2003) finds buyer size effects for hospital services. He finds that large insurers obtain discounts for hospital services, but the magnitude of the effects are small relative to the price effects of insurers' ability to steer patients. Ellison and Snyder (2010) examine the effects of buyer size in the purchase of antibiotics, a large therapeutic class of drugs. Like Sorensen, they find buyer size effects; unlike Sorensen, they show that buyer size effects are most relevant where there are substitution opportunities, highlighting an important interaction between buyer

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<sup>1</sup> In the simplest case, where buyer and seller's profit functions are linear in buyer size, and where the outside option for each side in the Nash game is zero, non-cooperation penalizes the buyer and seller identically. Hence, the solution to the game is invariant to buyer size. A large theoretical literature offers a variety of explanations for why buyer size has an ambiguous effect on upstream price negotiations. The recent studies cited in the text specify concavity conditions of the supplier's surplus function in order for larger buyers to extract larger rents in bilateral negotiations. Conditions in dynamic setting in which bargaining take place over repeated negotiations are studied by Snyder (1996).

size and upstream competition. Ellison and Snyder compare prices obtained by chains to those of independent retail pharmacies. Hence identification of buyer size effects rests on the assumption that, besides size, there are no other differences between chains and independent pharmacies that affect the prices of antibiotics negotiated with manufacturers.<sup>2</sup> Differences in pharmacy cost structure, distribution networks, and demand of its consumers, may also affect negotiated manufacturer prices.

Both these earlier papers demonstrated the existence and importance of buyer-size discounts in healthcare, and provided insight into their sources. We build on the existing literature in at least three ways. First, we provide a natural experiment in buyer size. The earlier studies did not have access to a natural experiment. Second, we develop the policy implications of buyer-size effects for public health insurance schemes. Specifically, we demonstrate that buyer-size effects create spillover effects from such schemes, because subsidies for insurance purchase expand private health insurers. Finally, we provide a very simple theoretical illustration of why buyer-size effects are ambiguous *a priori*, a point that is not always appreciated within the empirical literature.

We measure the effect of Part D enrollment increases—buyer size—on negotiated retail prices and profits earned on prescriptions for individuals *outside* the Part D program. The empirical analysis relies on disaggregated claims data from one large national retail pharmacy chain that reports the drug prices negotiated between the pharmacy and every insurer with whom it contracts.

An attractive feature of our approach is the absence of ex post rebates in agreements between pharmacies and insurers, making negotiated retail pharmacy prices readily observable and transparent. In addition, because the pharmacy's acquisition costs are constant across insurers for any given drug, between-insurer variation in negotiated prices reveals the marginal effect of insurer size on the pharmacy's unit *profits*, even without direct data on pharmacy costs. Finally, economic theory allows us to draw qualitative inferences about manufacturer profits from information about pharmacy profits. This strategy for indirect inference is valuable,

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<sup>2</sup> For clarity, note that buyers in our study are Part D insurers, who negotiate with upstream retail pharmacies over the *retail* price of drugs.

because net prices paid to by manufacturers and paid by insurers are almost never observed by researchers.<sup>3</sup>

We directly test the theoretical predictions that relate insurer enrollment size to retailer unit profits. We address possible endogeneity of Part D enrollment by exploiting variation in insurers' *potential* Part D enrollment, a plausibly exogenous measure of insurers' geographic exposure to Medicare beneficiaries without drug coverage prior to Part D. We find that increases in insurer enrollment lead to lower unit profits earned by the pharmacy. Consistent with theory, buyer size generates larger profit reductions for generic drugs than for branded drugs. We then estimate the enrollment effect on retail *prices*, which is of independent interest as a measure of insurer and consumer expenditures. We find that enrolling an additional 100,000 Part D beneficiaries enables an insurer to negotiate 2-percent lower unit prices for individuals enrolled in commercial plans external to Part D. We also test and reject alternative explanations for the observed buyer size effect, including the possibility that on average new Part D enrollees have lower demand for drugs than existing commercial enrollees, thereby lowering average demand (and negotiated prices) across insurers' total book of business.

Our results suggest that the “external effect” of Part D is not simply theoretical. Given the observed change in the enrollment of insurers participating in Part D, and the estimated enrollment-size elasticity, the program lowered overall retail prices for insurers' non-Part D enrollees by 3.7%. We estimate the market-wide external savings to be \$1.5 billion per year—nearly as large as the total “internal” cost-savings for Part D enrollees that lacked any previous drug coverage. If insurers retain the additional rents—as suggested by the research demonstrating limited competition in the commercial insurance market (Dafny, 2010)—then these price reductions represent a greater than 5% increase in average profitability for insurers writing both commercial prescription drug coverage and Part D policies.

The paper proceeds as follows: Section 2 discusses the MMA and the features of the drug market, and then describes the predictions of the standard Nash-bargaining framework applied to

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<sup>3</sup> Negotiations between insurers and pharmaceutical firms offer a second setting in which to test how insurer bargaining power affects bargaining outcomes. However, insurer-manufacturer negotiations typically involve complex pricing arrangements that include upfront pricing terms, or ex post rebates contingent on volume and other factors (Levy, 1999). And despite the policy importance of evaluating the rebates negotiated by manufacturers, the Centers for Medicare and Medicaid Services (CMS) proscribes the release of this data, which are similarly unavailable from private data vendors. It is thus difficult to measure directly the effect of bargaining power on price negotiations with manufacturers.

multilateral bargaining. Section 3 lays out the empirical strategy for estimating how enrollment affects pharmacy profits and prices as a consequence of insurer enrollment increases associated with Part D. Section 4 reports results of the empirical analyses of pharmacy profits. In Section 5, we report results of the price analyses; then decompose the effect of Part D on total prescription drug expenditure reductions into internal and external effects. Section 6 concludes.

## **2. The Economics of the Medicare Part D Prescription Drug Market**

### **2.1 Background on the Pharmaceutical Market and the MMA**

The 2003 Medicare Modernization Act (MMA) established Medicare outpatient prescription-drug coverage through the creation of the Part D drug benefit. The federal subsidies required to finance the program are significant, and have led to recent work examining how the program has affected pharmaceutical profitability (Frank and Newhouse, 2008; Friedman 2009). The high public cost of the MMA, coupled with concerns over its impact on Medicare's long-term sustainability, has focused attention on the Part D drug-purchasing model in particular.

Under the MMA, the government contracts with private insurers to administer drug plans. Hence, individual private insurers must negotiate retail drug prices and rebates directly with pharmacies and manufacturers. Figure 1 summarizes the purchasing model. A drug manufacturer earns revenues by selling drugs to wholesalers or directly to retail pharmacies at a price negotiated directly with each buyer. The manufacturer also negotiates rebates with individual insurers (private insurers, government agencies, and prescription benefits managers) that are tied to the insurers' purchase of its drug and inclusion, or preferential tiering, in their formularies.<sup>4</sup>

Pharmacies similarly negotiate with individual insurers over the retail price they are paid when they dispense prescriptions to an insurer's enrollees. These bilateral negotiations generate retail prices reported in pharmacy claims, such as the claims data used in this analysis. How the negotiated payment to the pharmacy is then split between enrollee and insurer depends on the specific premium, copayment and deductible architecture of the enrollee's insurance plan.<sup>5</sup>

Part D insurers are restricted by statute to negotiate manufacturer rebates for Part D enrollees in a separate and "firewalled" manner. In principle, this separation limits the

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<sup>4</sup> In contrast, rebates are rarely paid to or by pharmacies.

<sup>5</sup> For a comprehensive discussion of price negotiations and trends in reimbursements in the pharmaceutical industry, see Berndt and Newhouse (2010).

relationship between Part D and commercial lines of business within the same insurer. In practice, however, an insurer with more Part D enrollees may possess more *de facto* negotiating leverage in all transactions.

Note that no firewall exists for pharmacy retail price negotiations, allowing Part D enrollment to impact price negotiations for an insurer's commercial business directly. Moreover, changes in retail pharmacy prices impact both insurer and pharmacy profits. Any change in buyer profits may also indirectly impact manufacturer profits, because they determine the quantity of surplus available for pharmacies and manufacturers to share. This indirect effect on manufacturer profits will obtain even if the rebate firewall is respected.

## 2.2 Theoretical Relationships

Arguments for the possible effect of firm size on negotiations with suppliers have been posited since Galbraith (1952), and have been studied more formally in recent theoretical and empirical work.<sup>6</sup> In the pharmaceutical industry, where the distribution of rents among manufacturers, retail pharmacies, insurers, and enrollees has implications for health care costs, insurance coverage, and incentives to innovate, changes to the bargaining power of insurers can have a variety of effects that have not been widely studied.<sup>7</sup>

Models of two-way bargaining between retailers and buyers exist in the literature. A well-known example is that of Chipty and Snyder (1999), who implement a straightforward Nash-bargaining approach to the problem. The principal difference in our context is the existence of a third "player" in the negotiations, namely the upstream manufacturer. In the Theoretical Appendix, we demonstrate that Chipty and Snyder's results generalize quite

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<sup>6</sup> Among many theoretical studies on the topic, recent work includes Stole and Zwiebel (1996), Brooks et al. (1997), Chipty and Snyder (1999), and Raskovich (2003), who specify concavity conditions that the supplier's surplus function must satisfy in order for large buyers to extract rents. Snyder (1996) studies this issue in dynamic settings. In the health literature, Sorenson (2003) studies the extent to which insurers' ability to exclude hospitals affect negotiated hospital (supplier) prices. In this paper, we estimate how bargaining power changes with the *size* of the buyer, holding constant its ability to steer its market. Consistent with Sorenson, our model suggests that buyer size affects bargaining power only when the buyer has some ability to steer its share across suppliers. We then explicitly test a model in which buyer size can either augment or diminish the impact of threats of network exclusion on supplier profits and negotiated prices in the pharmaceutical industry.

<sup>7</sup> See the discussion of Sorenson (2003) and Ellison and Snyder (2010) in the introduction. Other studies in the health care literature have primarily focused on how characteristics of *upstream* providers affect negotiations with *downstream* payers (Town and Vistnes 1999). More recently, Ho (2009) studies how hospital performance and provider network structure affect bargaining outcomes with downstream payers. In the pharmaceutical industry, the complex market structure and paucity of negotiated price data makes these issues difficult to study.

naturally to this three-way bargaining context. Following the theoretical findings of Chipty and Snyder, the analysis of bargaining among these players generates three empirical implications.

**Implication 1: Insurer size has an ambiguous effect on the profits of both the pharmacy and the insurer.** The negotiating leverage of one side depends on the surplus it generates for its trading partners. If larger buyers generate more surplus per unit for their partners, they will receive better prices, and vice-versa. Typical of standard bilateral bargaining models, the effect of buyer size on unit surplus is ambiguous and depends on the curvature of the supplier's surplus function (Stole and Zwiebel, 1996; Brooks et al., 1997; Chipty and Snyder, 1999; Raskovich, 2003). Intuitively, if the seller's surplus function is convex in quantity, a larger marginal buyer generates more surplus per unit sold than would be generated by a smaller marginal buyer. In such a case, the larger buyer is more valuable to the seller on the margin and will thus extract more favorable terms. The reverse is also true: a concave surplus function means a larger buyer generates less surplus per unit sold and receives less favorable terms. Figure 2 illustrates the basic intuition. In the empirical analysis that follows, we directly estimate the direction and magnitude of the buyer-size effect on unit profits in the context of the retailer-buyer price negotiations.

**Implication 2: When drug manufacturers hold more of the bargaining power, insurer size causes smaller changes in insurer or pharmacy profits.** Intuitively, consider the extreme case where manufacturers extract all the rents. This would be true if the manufacturer held a patent monopoly on a product with perfectly inelastic demand. Thus, all downstream parties earn zero profits, and there is nothing for the downstream firms to divide. As a result, increases in insurer size cannot change the prices paid by the insurer to the pharmacy. Branded drugs with no therapeutic substitutes (and thus, inelastic demand) may exemplify such cases. Generic drugs represent the opposite extreme. For these drugs, bargaining power is held entirely by pharmacies and insurers, in which case the potential impact of changes in insurer size on the share of surplus accruing to pharmacies and insurers is maximized.

**Implication 3: When all sides have some degree of bargaining power, changes in profits of manufacturers and pharmacies correlate positively with external shocks to the insurer's size.** Among other things, this means that increases in the size of an insurer will have qualitatively similar effects on the profits of pharmacies and manufacturers. Because manufacturer profits are not observed, we cannot test this implication directly. Nevertheless, if

the Nash-bargaining model is valid, we can draw qualitative inferences about changes in manufacturer profits, using the sign of the enrollment-size elasticity we estimate on retail pharmacy profits.<sup>8</sup> Specifically, if larger insurers receive lower prices on Drug X from pharmacies, the Nash-bargaining framework implies that larger insurers drive down the profits earned on Drug X by both the manufacturer and the pharmacy.

### **3. Empirical Strategy**

To test the theoretical predictions about profits, we empirically examine the impact of insurer enrollment on changes in pharmacy profits per unit (i.e., per pill) and investigate how these vary with the competitiveness of drug classes. These findings, coupled with the theoretical predictions above, are used to draw quantitative and qualitative inferences about the distribution of rents among manufacturers, pharmacies, insurers, Part D enrollees, and commercial enrollees. We then measure the marginal impact of insurer enrollment on retail prices, which we use to quantify the absolute external impact of Part D-related enrollment on retail drug expenditures in the non-Part D commercial market.

#### **3.1 Data**

Data on prescription drug utilization and expenditures come from a national retail pharmacy chain. As of January 1, 2006, when Medicare Part D was implemented, the pharmacy chain had retail presence in 45 states, and prescriptions filled at its pharmacies accounted for approximately one-fourth of the US prescription market.

We obtained all pharmacy claims for a five percent random sample of unique pharmacy customers over the age of 60. For these individuals, we obtained data on claims for every prescription filled at the chain between September 1, 2004 and April 31, 2007. Each claim reports the National Drug Code (NDC) of the prescription filled; its therapeutic class; the pill quantity; the number of treatment days; the date dispensed; the identification of the third-party payer; whether the insurance plan is a commercial, Medicare Part D, or Medicare Advantage

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<sup>8</sup> This inference is valid even if the Part D firewall necessitates parallel and independent sets of rebate negotiations between insurers and manufacturers. Again, regardless of how insurers negotiate with manufacturers, a change in insurer size will impact the surplus flowing to pharmacies; this surplus is ultimately shared with manufacturers. This creates a positive relationship between pharmacy and manufacturer profits that goes beyond the rebate negotiation. For this reason, it is simpler to develop the bargaining model without specifying particular assumptions about the firewall between the Part D and non-Part D markets.

plan; the out-of-pocket and third-party payer expenditures; and the address of the pharmacy where the claim took place. The claims data also contain information on subjects' demographic characteristics (date of birth, sex, language preference, and zip code of residence). With these data, we are able to determine the drug prices negotiated between the pharmacy and each insurer for every drug appearing in the claims.

The pharmacy claims data report drug utilization largely consistent with that reported in the Medical Expenditure Panel Survey (MEPS) for the same period. Table 1 lists the top 25 drugs utilized by seniors between 2004 and 2006 by pharmacy revenues, which correspond closely in rank with drugs observed in the MEPS over the same period. Notable exceptions are physician-administered drugs that are under-represented in out-patient retail pharmacy claims.

Negotiated pharmacy prices for the same drug vary considerably across insurers. Figures 2a and 2b show the distribution of normalized retail prices of drugs at the NDC-insurer level, separately for branded and generic drugs. In Figure 2a, prices are measured as the percentage difference between the retail price an NDC-level drug negotiated by an insurer and the average price of that drug across all insurers; in Figure 2b, prices are measured as the absolute difference between these two quantities.

In both figures, variation for a given drug exists across insurers. The distribution of relative prices of generic drugs is noticeably wider than the distribution of prices of branded drugs, whether measured in ratios or in absolute differences. These facts are consistent with relatively less surplus accruing to pharmacies when manufacturers have more bargaining power; when there is less surplus to be bargained over in negotiations between the pharmacy and downstream buyers, less price dispersion results. The fact that the absolute price differences for generic drugs exceed that of branded drugs is particularly striking given the much lower price levels for generic drugs. This suggests the strength of the economic forces that eliminate downstream surplus for branded drugs.

Our database contains most large insurers that participate in Medicare Part D. In general, two reasons explain why a Part D insurer would not appear in our claims database: (1) the pharmacy did not contract with the insurer; or (2) claims from the insurer are not sampled from the full pharmacy claims. Both these reasons suggest that smaller insurers are less likely to appear in the claims data. Table 2 shows the distribution of Part D insurers represented in our sample of pharmacy claims according to their 2007 Part D enrollment. In total, 86 Part D

insurers appear in the claims data. This list includes insurers that offer Part D Plans (PDPs), Medicare Advantage plans, or demonstration plans. Our analysis is eventually restricted to a set of 33 insurers that offer at least one PDP.<sup>9</sup> Table 2's columns parse the insurer universe by Part D enrollment. Note that the distribution of Part D enrollment by insurer is highly skewed. For instance, the median Part D insurer enrolls fewer than 6,400 Part D seniors, while the 90<sup>th</sup> percentile Part D insurer has more than 20-times greater Part D enrollment.

Data on enrollment, premiums and benefit design for Part D and Medicare Advantage plans come from the Centers for Medicare and Medicaid Services (CMS). Plan-level information also identifies the sponsoring insurance firm, so we can aggregate enrollment to the insurer. Premium information corresponds to end-of-year open-enrollment premium pricing for coverage beginning the following year. The CMS website makes publically available both the enrollment and Part D Landscape files.

## **3.2 Basic Empirical Framework and Threats to Identification**

### **3.2.1 *Supplier Profit Equation***

To test how enrollment affects profits, we exploit the introduction of Part D, which brought millions of newly insured individuals onto the rolls of existing insurers. We test whether insurers that experienced greater enrollment increases due to Part D negotiated lower unit profits earned by the pharmacy in their non-Part D commercial market.

At its core, the theory suggests that pharmacy profits earned on drug  $d$  and insurer  $i$  depend on insurer  $i$ 's size. We take two steps to make this conceptual framework empirically tractable. First, since the functional form is unknown, and there is no obvious functional form satisfying all the relevant conditions of the theory, we take a linear approximation to the profit function around the pre-Part D equilibrium point. The linear approximation approach will capture the local average treatment effect of Part D, but it does not allow us to conduct counterfactual simulations, which we avoid. The result is a regression equation that relates the change in profits due to Part D, as a linear function of the change in enrollment, and a vector of

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<sup>9</sup> The first-difference estimation framework outlined below requires repeated claims for each insurer-NDC cell. 74 of the 89 Part D insurers observed in the claims data have repeated claims for at least one insurer-drug cell. 33 of 74 of the observed insurers offer at least one stand alone Part D plan. This sample eliminates insurers that offer only Medicare Advantage plans (whose premiums also cover medical care) and Medicare prescription drug demonstration plans. To maintain consistency across specification, the analyses are restricted to the 33 private insurers that offer stand alone Part D plans.

other drug-insurer characteristics and controls,  $X_{d,i}$ . Second, we use the change in an insurer's Part D enrollment as a proxy for the total change in enrollment – the empirical implications of this assumption are discussed in Section 3.2.3, and our approach for dealing with the resulting bias is discussed in Section 3.3.1. The estimating equation reads as follows:

$$(1) \quad \Delta profits_{d,i} = \alpha + \beta \Delta Part\ D\ Enrollment_i + \gamma \Delta X_{d,i} + \varepsilon_{d,i}$$

The dependent variable is the change in the profits per pill earned by the pharmacy over prescriptions of drug  $d$  filled by enrollees of insurer  $i$  from before and after the implementation of Part D.<sup>10</sup> Here and throughout the balance of our analysis, a “drug”  $d$  is an NDC number.<sup>11</sup> Pharmacy profits are defined as  $profits_{d,i} \equiv price_{d,i} - cost_{d,t}$ .  $Price_{d,i}$  is the negotiated retail price, which varies across drug and insurer;  $cost_{d,t}$  is the acquisition cost of drug  $d$  to the pharmacy, which is constant across insurers. It follows that  $\Delta profits_{d,i} = \Delta price_{d,i} - \Delta cost_{d,t}$ , where differences in retail profits across insurers is driven solely by differences in negotiated retail prices, for a given drug.

Note that the pharmacy claims data do not report acquisition costs or profits —only negotiated prices are reported. To estimate equation (1), therefore, we rewrite it as:

$$(2) \quad \Delta price_{d,i} = \alpha + \beta \Delta Part\ D\ Enrollment_i + \gamma \Delta X_{d,i} + \delta_d + \varepsilon_{d,i}.$$

The drug-level intercept,  $\delta_d$ , captures drug-specific unobserved characteristics, including  $\Delta cost_{d,t}$ , which varies across drugs but is constant across insurers for a given drug. Regressing first-differenced price levels on enrollment and NDC-level drug fixed effects allows  $\beta$  in equation (2) to be interpreted as the effect of enrollment on pharmacy *profits*. This interpretation is justified rigorously in the theoretical appendix. All estimates of the semi-elasticity  $\beta$  reported below come from the estimation of equation (2).

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<sup>10</sup> Price negotiations for drugs between insurers and pharmacies are conducted at the national level; hence, this study is conducted at the insurer-drug level. Some insurers negotiate through prescription benefits managers (PBMs). In cases when the pharmacy claims specifically identifies both an insurer and its PBM as the payer, the insurer in the sample is defined at the level of its PBM. In a previous version of this paper, we conducted the analysis at the insurer-drug-state level in order to capture slight differences in factor costs (and hence, unit profits) across states. These differences are negligible and are averaged-out in the insurer-drug analysis. Conducting the analysis at the disaggregated level yields nearly identical results to those reported here.

<sup>11</sup> Technically, an NDC number is below the aggregation level of a drug, or molecule, because it identifies different labelers and packages for a given molecule. The 10-digit NDC code captures package numbers (for example, drugs sold in blister packs), and are the most disaggregated level of analysis possible. Note that for many NDCs, pill counts are at the discretion of the prescribing physician, and are unlikely to be systematically related to insurer size. Not surprisingly, inclusion of the average pill count control variable has negligible effects on the estimated buyer size effects. Its inclusion serves mainly to absorb residual variation.

We calculate the dependent variable in equation (2) as the difference in unit drug prices averaged over the second half of 2005 and the first half of 2006; these are the six-month periods prior and subsequent to the implementation of Part D. To assess whether there are any pre-existing trends that might contaminate our estimation, we use price changes between the second half of 2004 and the second half of 2005 in falsification tests. The key independent variable is the change in each insurer's Part D enrollment due to the implementation of Medicare Part D in 2006.

The vector of covariates,  $X$ , includes a measure of each insurer's exposure to the pharmacy,<sup>12</sup> the average wholesale price of the drug, and the average number of pills sold per prescription for a given drug  $d$  and insurer  $i$ .<sup>13</sup> In specifications that include drug fixed effects, we necessarily drop the wholesale price change variable. The first-difference specification also eliminates time-invariant drug, insurer, and market-level characteristics. The coefficient on enrollment changes captures the average linear effect of enrollment increases on negotiated retail prices across insurers of all sizes.

The coefficient  $\beta$  in equation (2) is also relevant for understanding the effect of insurer enrollment increases on drug manufacturer profits. The Nash-bargaining framework implies a positive correlation between changes in pharmacy and manufacturer profits that result from changes in insurer size. Therefore, under Nash-bargaining, the coefficient  $\beta$  in equation (2) has the same sign as the effect of insurer enrollment on manufacturer profits.

The analysis is conducted at the insurer-NDC level, which is equivalent to an NDC-weighted insurer-level regression of average profits earned by the pharmacy on insurer-level enrollment increases. All analyses are clustered at the insurer level, an acknowledgment that identification comes from cross-insurer variation in enrollment changes. This clustering allows for cross-NDC intra-insurer correlation in the error structure, allowing us to remain agnostic on

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<sup>12</sup> Theory suggests that greater exposure to the pharmacy may affect bargaining power in bilateral negotiations. *Exposure* is an insurer-level measure of the market share of the pharmacy with respect to an insurer. It is calculated as the average market share of the pharmacy across states, where state-level market shares are weighted by an insurer's total business in each state. Data on pharmacy market share was obtained from the Chain Store Guide, which reports annual sales and store counts of all pharmacies (total and by-chain) for local geographies in the US, for 2004 through 2008.

<sup>13</sup> Recall that the dependent variable is the price per pill averaged over all prescriptions observed in each cell. Given that prescriptions contain any number of pills, this measure controls for changes in the average number of pills per prescription in each cell over time, which may affect average price per pill through slight bulk-rate pricing.

how discounts are distributed across NDCs within an insurer. As discussed in Section 2, our model does make one theoretical prediction related to differential buyer size effects across NDCs—that there should be smaller size effects when more surplus is retained by the manufacturer (e.g. branded drugs with few substitutes). This motivates specification where we limit the sample to only branded drugs, or to only generic drugs.

### 3.2.2 *Retail Pricing Equation*

While economic theory provides more powerful predictions for changes in profits than changes in pharmacy prices, the latter are of independent empirical interest as a measure of changes in external costs to commercial enrollees and insurers. Therefore, we estimate a second set of models relating log price changes to Part D enrollment changes:

$$(3) \quad \Delta \ln(\text{prices}_{d,i}) = \alpha + \theta \Delta \text{Part D Enrollment}_i + \gamma \Delta X_{d,i} + \varepsilon_{d,i}$$

The log-linear specification in equation (3) generates an estimate of the semi-elasticity,  $\theta$ , that captures the average effect of enrollment increases on log negotiated retail prices across insurers of all sizes at baseline. We then use our estimate of  $\theta$  in equation (3) and the observed distribution of enrollment increases to quantify the change in insurer costs in the commercial market, in levels and percentages, associated with Part D-related insurer size increases.

Note that in estimating  $\beta$  in equations (2) and (3), we are estimating enrollment semi-elasticities, not strict elasticities. If we observed total commercial enrollment, a more structural approach to estimating the effect of log changes in total enrollment levels would be possible. As noted earlier, absent data on non-Part D commercial enrollment, we adopt linear approximations, which will be sufficient as long as we limit ourselves to estimating effects of observed enrollment increases, rather than attempting to forecast counterfactuals far off the equilibrium path. To the extent that a given enrollment increase generates heterogeneous effects at different initial levels, the local average treatment effects in the data will faithfully estimate the actual historical effects of Part D on profits and prices. This is how we interpret our results, and do not attempt to engage in simulations of what might happen in the future with enrollment increases that are far out of sample.

### 3.2.3 *Threats to Identification*

In equations (2) and (3), we estimate the impact of changes in enrollment due to Part D on supplier unit profits and prices. However, the model in Section 2 implies that changes in *total*

enrollment, not just changes in Part D enrollment, affect unit profits of suppliers. Data on insurers' total prescription drug enrollment are not available,<sup>14</sup> but Part D enrollment data is publicly available from CMS. Therefore, we model the impact of new Part D enrollment on changes in negotiated prices. Unobserved changes in commercial enrollment appear in the error term. Potential correlation with Part D enrollment creates an omitted variables problem.

Related to this, Part D enrollment reported by CMS represents the total number of new Part D plan members. This figure includes those seniors who were previously enrolled in commercial retiree plans and then switched to a Part D plan. Crowd-out of this sort appears to be sizeable in aggregate (Englehardt and Gruber 2010). Lack of data on total drug insurance enrollment at the firm level prevents our calculating *net* increases in Part D enrollment for each insurer. As a result, crowd-out will generally lead to our *underestimating* the true impact of enrollment on negotiated prices. To see how, note that:

$$(4) \quad \Delta \text{Part D Enrollment}_i^{\text{Net}} = \rho \Delta \text{Part D Enrollment}_i^{\text{Observed}} + \xi_i.$$

We model *observed* changes in Part D enrollment as mismeasurement of the *net* change in Part D enrollment, where  $0 < \rho < 1$  is the market-average proportional extent of crowd-out, and  $\xi$  is an insurer-level mean-zero error term. Also note the accounting identity:

$$(5) \quad \Delta \text{Total Enrollment}_{i,t} = \Delta \text{Part D Enrollment}_{i,t}^{\text{Net}} + \Delta \text{Commercial Enrollment}_{i,t}.$$

The object of interest is the relationship between total insurer size and pharmacy profits, or the coefficient  $\psi$  in:

$$(6) \quad \Delta \text{profits}_{d,i} = \alpha + \psi \Delta \text{Total Enrollment}_i + \gamma \Delta X_{d,i} + v_{d,i}.$$

Substituting equations (4) and (5) into equation (6) yields:

$$(7) \quad \Delta \text{profits}_{d,i} = \alpha + \psi \rho \Delta \text{Part D Enrollment}_i^{\text{Observed}} + \gamma \Delta X_{d,i} + \varepsilon_d,$$

where  $\varepsilon_{d,i} \equiv \psi \Delta \text{Commercial Enrollment}_i + \psi \xi_i + v_{d,i}$ .<sup>15</sup> Our estimate of  $\beta$ , the Part D enrollment effect on profits in equations (1) and (2) is, therefore, an *underestimate* of the

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<sup>14</sup> Data on insurers' *total medical* insurance enrollment are available from several sources. For instance, TheStreet.com (and previously Weiss Ratings), report enrollment, network size, assets and income for every medical underwriter annually. However, enrollment in drug insurance—the relevant measure of buyer size for drug price negotiations—is not reported separately. Further, enrollment of prescription benefits managers (PBMs) is not reported by these publishers.

<sup>15</sup> Analogous to equations (1) and (2), recall that the coefficient on enrollment in equation (7) is equivalent to the coefficient on enrollment in  $\Delta \text{prices}_{d,i} = \alpha + \psi \rho \Delta \text{Part D Enrollment}_i^{\text{Observed}} + \gamma \Delta X_{d,i} + \delta_d + \varepsilon_d$ —i.e. the profit equation in which changes in level prices are regressed against Part D enrollment and NDC-level drug fixed effects.

structural enrollment effect,  $\psi$ , by a proportionality factor of  $0 < \rho < 1$ ; as a result, one should view our estimates as lower bounds on the true enrollment effect.<sup>16</sup> Likewise, our estimate of  $\theta$ , the enrollment semi-elasticity of retail prices from equation (3), is an underestimate of the structural elasticity by the same proportional factor,  $\rho$ .

Equation (7) makes clear how potential sources of endogeneity may bias our estimate of  $\beta \equiv \psi\rho$ . Most obviously, changes in Part D enrollment may be correlated with changes in commercial enrollment (i.e., correlation between Part D enrollment and  $\xi_{i,t}$ ). This could happen if a more aggressive pursuit of enrollment in the Part D market is associated with a similar pursuit of commercial market share, leading to an upward bias in our estimate of  $\beta$ ; alternatively, pursuit of Part D enrollment may be associated with less aggressive growth in commercial enrollment, due to insurers specializing in the new Part D market, leading to downward bias. To the extent that this correlation exists, instruments are needed to generate unbiased estimates of  $\beta$ .

There may also be additional sources of bias operating through a correlation between changes in enrollment and  $v_{d,i,t}$ . There is evidence that Part D enrollees place excessive weight on premiums over drug prices and cost-sharing when selecting plans (Abaluck and Gruber, 2011; Kling *et al*, 2012). The sickest patients, however, may be more sensitive to out-of-pocket costs relative to the average enrollee, and therefore more likely to select into plans with greater expected future price declines, leading to higher premiums for those plans. Consequently, plans with larger increases in out-of-pocket costs may have lower initial premiums and greater enrollment; this would bias estimates of  $\beta$  downward. Institutional rules may drive dual-eligible Medicaid beneficiaries (individuals enrolled in both Medicaid and Medicare) into plans with below-median premiums that may enjoy systematically slower (or faster) growth in drug prices. Either self-selection, or automated enrollment, into plans with lower future drug price growth would also generate biased estimates of the enrollment elasticity of price changes.<sup>17</sup> These kinds of issues are also addressed by an instrumental variables strategy, discussed in detail below.

Finally, the timing of and nature of contracts could cause OLS estimates of  $\beta$  to be downwardly biased. Contracts typically last one to two years, and are volume-based. Volume-

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<sup>16</sup> The market-wide extent of crowd-out,  $\rho$ , can be calculated in the aggregate data, implying that  $\psi$  can be recovered by scaling up our estimated  $\beta$  by a factor of  $1/\rho$ .

<sup>17</sup> Another source of bias may come from a correlation between enrollment (due to low premiums) and the propensity of a plan to drive generic drug utilization. This would bias the estimate of the enrollment elasticity in the cross section, but not in a first difference analysis of unit price changes for a given insurer-drug.

based contracts account for ongoing changes in enrollment and claims volume. In instances where contracts are not volume-based, the timing of Part D is such that our data may measure prices between contracts. Forward looking parties would nevertheless factor in *anticipated* changes in enrollment, but such contracts would not capture *unanticipated* enrollment changes. Unanticipated enrollment changes function statistically as measurement error, biasing downward OLS estimates of  $\beta$ . The prevalence of volume-based contracts suggests that this bias would be small. Nevertheless, this bias is addressed by the instrumental variables strategy discussed below.

### **3.3 Identification Strategy**

#### **3.3.1 Instrumental Variables and Validity**

To address the sources of endogeneity identified above, we implement an instrumental variables strategy that exploits a key predictor of insurers' new 2006 Part D enrollment: insurer's geographic exposure to total potential Part D enrollment. Intuitively, insurers may find themselves to be in stronger or weaker positions to capture Part D enrollees, purely as a function of their geographic presence several years prior to the implementation of Part D. For example, insurers located in states with many Medicare beneficiaries without private health insurance are well-positioned to attract new Part D enrollees, and vice-versa.

Conceptually, insurance is regulated at the state-level, creating regulatory barriers and costly entry by an insurer into a new state. Labor and other capacity constraints may also compound the regulatory barriers. Hence, commercial underwriting presence in a state prior to Part D facilitates entry into that state's Part D market. Indeed, among Part D insurers, little difference exists between their 2006 Part D state penetration and their 2004 or 2005 commercial presence. Therefore, we use geographic variation in insurer location prior to the implementation of Part D to generate plausibly exogenous variation in their subsequent exposure to Part D enrollees.

A number of possible concerns with this instrument are alleviated by the unique form of our estimation problem. Since we are analyzing the first year of Part D implementation, the *level* of Part D enrollment is equal to the *change* in Part D enrollment. Therefore, we are estimating the relationship between price changes and Part D enrollment *levels*. Bias will occur if unobserved heterogeneity in predictors for enrollment levels is correlated with unobserved

heterogeneity in price changes. This would be an unusual type of correlation. In particular, our instrument presumes that the level of an insurer's exposure to previously uninsured Medicare beneficiaries is uncorrelated with anticipated changes in an insurer's future negotiated prices. Later, we present evidence supporting this assumption.

The instrument implies the following first-stage equation, which precedes the second-stage profit-enrollment equation:

$$(8) \quad \Delta \text{Part D Enrollment}_i = \gamma_0 + \gamma_1 \text{Potential Part D Enrollment}_i + \Gamma X_i + \eta_i$$

An insurer's *Potential Part D Enrollment* is the sum of seniors without private health insurance in 2005 (and hence, without prescription drug coverage), across all states in which the insurer was present in 2005.<sup>18</sup> We formally define *Potential Part D Enrollment* as:

$$(9) \quad \text{Potential Part D Enrollment}_i = \sum_m (\text{Seniors without PHI}_{m,t=2005} \cdot 1[\text{insurer}_i \text{ in } m]_{t=2005})$$

Naturally, insurers with potential for a large enrollment in their Part D market are on average likely to be large insurers with national commercial presence. A potential validity issue arises if larger insurers are both more likely to be present in more markets and more likely to experience systematically different price changes, through channels other than Part D enrollment. For example, due simply to geographic coverage, larger insurers may have greater exposure to uninsured seniors. If the commercial market size of insurers simply grew at the same rate, larger insurers would naturally have greater increases in commercial enrollment. Buyer size effects from increased commercial enrollment would then be correlated with the potential Part D enrollment variable.

We examine this potential validity issue in two ways. First, we estimated the relationship between firm size, as measured by the size of its observed commercial market in the pharmacy claims, and pre-Part D changes in drug prices. Importantly, we find no relationship between the two quantities across various specifications and definitions of insurer size. Results from these tests are reported in Appendix Table 1. Second, we directly estimate the relationship between potential Part D enrollment levels and changes in commercial enrollment. While we are limited

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<sup>18</sup> Data come from the CPS. This count includes seniors enrolled in Medicaid prior to Part D implementation. While seniors eligible for both Medicaid and Medicare received their drug coverage through state Medicaid prior to Part D, they are covered by private Part D insurers under the MMA, and thus constitute a part of the increase in enrollment in private insurance rolls as a result of Part D implementation.

by the lack of market data on insurers' commercial enrollment, we do have pharmacy claims data for a limited set of commercial enrollees—those ages 60-64. Using these data, we are able to regress the potential Part D enrollment variable on 2005 to 2006 changes in commercial market size, as measured by total commercial claims for those aged 60-64. This provides a scaled measure of commercial enrollment size. We find that the coefficient on potential Part D enrollment is negative and statistically insignificant. This finding is inconsistent with the hypothesis that exposure to greater levels of uninsured seniors is systematically related to increases in commercial market size.<sup>19</sup>

To our knowledge, there are no other obvious theoretical reasons why pre-Part D exposure to more uninsured Medicare beneficiaries should predict differential future post-Part D growth in the unit prices paid by an insurer. This is especially true in light of our finding that there is no relationship between enrollment and prices changes prior to the implementation of Part D, or between Part D enrollment and changes in commercial market size.

### **3.3.2 *Understanding the Explanatory Power of the Enrollment and Pricing Instruments***

Table 3 illustrates the operation of the instrument. The table presents data on four insurers—similar in the sizes of their total commercial enrollment—ordered from smallest to largest in terms of commercial claims expenditures, reported in column (1). Exposure to the Part D marketplace, reported in column (2), is not just a simple function of size. Insurer A (the smallest) has the greatest potential Part D exposure, due to its heavy market penetration into states with high elderly population shares (e.g., Florida). Column (3), which reports actual Part D enrollment, shows that the greatest Part D enrollment ends up accruing to Insurer A, as predicted by the potential enrollment variable. Note that the ranking of actual enrollment values tracks that of potential enrollment, except that potential enrollment fails to distinguish between Insurers B and C, which have very similar potential enrollment values.

Our estimation sample contains the 33 insurers in our data for which we can calculate all the necessary covariates. Figure 4 reports actual versus potential enrollment for 32 of these 33. While potential enrollment is not perfectly correlated with actual enrollment, there is a visual upward slope in the relationship. The “noise” in the relationship appears to come from some insurers that stay out of the Part D market, rather than insurers who secure far greater Part D

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<sup>19</sup> We thank an anonymous referee for raising this possible source of endogeneity in the potential Part D enrollment instrument. Results of this test are available upon request.

enrollment than predicted. This suggests that potential enrollment creates option-value for insurers, which many (but not all) exercise. Figure 3 excludes one insurer, Humana, whose actual enrollment of 4.5 million would skew the figure so much as to render the other points indistinguishable.

### 3.3.3 First-Stage Power

The enrollment instrument appears to successfully “treat” 32 of the 33 insurers, with the one exception of Humana. To show this formally, we calculated the first-stage F-statistics after sequentially dropping each of the 33 insurers. We found that the first-stage F statistic is no larger than 2.01 across the 32 insurer combinations that include Humana; and 16.00 in the one specification that excludes Humana. Next, we repeated the exercise among the 32 “treated” insurers, excluding Humana. The F-statistics are consistently larger than 10.65, suggesting that no one insurer is driving the first stage explanatory power of the potential enrollment measure.

This finding is consistent with known features of Humana’s business strategy, which aimed to under-price premiums to gain Part D market share, and then to switch enrollees into their highly profitable Medicare managed care plans (Krasner, 2006; Business Week, 2006). To see Humana’s pricing strategy more clearly, we regress the log premium of plan  $p$  offered by insurer  $i$  in market  $m$  on plan characteristics  $D^{20}$  and insurer fixed-effects,  $\sigma_i$ , as in:

$$(10) \quad \ln(\text{premiums}_{p,i,m}) = \alpha + D_{p,i,m}\beta + \sigma_i + \theta_m + \varepsilon_{p,i,m}$$

The sample of plans is restricted to standard and actuarially equivalent plans sold in the same CMS-defined Part D market, thereby allowing us to purge observed premiums of benefit design and plan generosity characteristics, and to isolate pure premium pricing variation at the insurer-level. Figure 5 depicts the distribution of insurer fixed effects for the 33 insurers in our sample. The figure reveals a fairly tight “bell” shape in the distribution of the fixed effects. Humana is clearly an outlier, on average pricing its plan at a 70 percent discount relative to identical plans sold in the same market. While the potential enrollment measure predicts Humana to be among the larger Part D insurers, it nonetheless vastly understates the actual enrollment increase for Humana, likely due to its aggressive premium discounting strategy.

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<sup>20</sup> The vector  $D$  contains the following plan-level characteristics: plan annual deductible; whether the plan is a low income subsidy plan; whether the plan covers generics and branded drugs in the “coverage gap;” whether the plan covers generics and some branded drugs in the coverage gap; and whether the plan covers no drugs in the coverage gap.

Including Humana in the analysis necessitates accounting for its exceptional premium pricing strategy, and the nonlinear form of its enrollment effect. Analyzing all the non-Humana insurers allows for a simpler, linear design that serves as the baseline specification reported in the text. The Appendix reports more complex, nonlinear models that include Humana as well. Significantly, the estimates are highly similar across the two samples.

To reflect the sample of insurers, we interpret our baseline estimates as valid local average treatment effects of enrollment on profits and prices for all insurers other than Humana. Our estimates of aggregate cost savings also conservatively assume that Humana makes no contribution to the market-level effects of enrollment on aggregate expenditures.

The Empirical Appendix introduces a nonlinear effect of enrollment on prices and profits, along with a more complex first-stage estimation strategy that better reflects Humana's pricing behavior. Specifically, the Appendix exploits both geographic variation in insurers and variation in each firm's quality-adjusted premium (premium purged of benefits generosity and drug costs), as an additional instrument that bears on Humana's premium-setting strategy. This approach generates estimates that fit Humana's behavior better and should in principle apply to Humana as well. Estimated enrollment elasticities and predicted price declines are nearly identical for the 32 common insurers across the two strategies. The only notable difference is in the estimated aggregate cost savings: including Humana leads to slightly larger savings.

### **3.4 Market Structure and General Equilibrium Considerations**

The coefficient on new Part D enrollment in equations (2) and (7) can be interpreted as the effect of insurer size on retail unit profits and log prices. Yet, in general equilibrium, the estimate might be confounded by the responses of suppliers. For example, when insurers gain more leverage, pharmacies have greater incentives to grow, consolidate, or otherwise strengthen their own bargaining positions. If this occurs, our coefficient will underestimate the true effect of insurer size on negotiated retail prices and unit profits of the pharmacy.

We account for the general equilibrium possibility in two ways. First, as mentioned in section 3.2.1, we control for the leverage of the pharmacy with a measure of changes in insurers' exposure to the pharmacy. *Exposure* is an insurer-level measure of the state-level market share of the pharmacy, weighted by the fraction of the insurer's total retail expenditures in each state.

The control variable strategy works, so long as the pharmacy's bargaining leverage enters the estimating equation linearly.

As a more general approach, we explicitly test for the presence of general equilibrium effects, over the time-frame of our data. In particular, we estimate whether the pharmacy increases its market share when insurers increase enrollment as a result of Part D, as in:

$$(11) \quad \Delta \log(\text{exposure}_i) = \rho_1 + \rho_2 \Delta \text{Part D Enrollment}_i + \omega_i$$

We estimate this equation for all insurers in the sample, where *exposure* is the insurer-level measure of the pharmacy's market share described above. We estimate this equation using OLS and IV, making use of the instrument described in Section 3.3.

Results are presented in Appendix Table 2. Columns (3)–(6) indicate that changes in pharmacy bargaining power—between 2005 and 2006, and between 2005 and 2007—are uncorrelated with corresponding increases in insurers' Part D enrollment over the same periods. Columns (1) and (2) show that changes in pharmacy market share between 2004 and 2005 are uncorrelated with changes in insurer enrollment between 2005 and 2006, implying that pharmacies did not respond to Part D-related enrollment increases. These results suggest that, at least in the short-run two-year period of our analysis, there is no evidence of general equilibrium responses that would contaminate the estimation of the structural enrollment elasticity of interest.

#### **4. Identifying the Effects of Part D on Pharmacy Profits**

We examine the effects of insurer enrollment size on pharmacy profits per pill, and price per pill, in the commercial market external to Medicare Part D.

##### **4.1 Correlation between Unit Prices and Enrollment**

To examine the correlation between changes in unit prices and enrollment, we regress change in log drug prices on all covariates in equation (3) except the key explanatory variable, insurers' Part D enrollment. We then plot residuals from this regression against insurers' Part D enrollment. Figure 6a shows the scatter plot of residuals against first year Part D enrollment. The correlation is clearly negative. We test for a causal relationship in the analyses below.

##### **4.2 Enrollment and Retail Profits in the Commercial, Non-Part D, Market**

Table 4 reports enrollment effects on changes in retail profits-per-pill from the estimation of equation (1) on all commercially insured, non-Part D claims. As described in Section 3, the

effect on profits-per-pill can be recovered by estimating equation (2) using price changes as the dependent variable, with drug-level fixed effects. The dependent variable reflects changes in average profits per pill between the second half of 2005 and the first half of 2006 at the insurer-drug level. In all specifications, insurer-drug level observations are weighted by the number of claims for the drug observed for that insurer.

Columns (1) and (2) report our results. Among the 32 “treated” insurers, the first-stage explanatory power of the potential enrollment instrument is sufficiently strong. In the second stage, among all other insurers, enrollment lowers pharmacy profits. 100,000 additional Part D enrollees lower profits per pill in the commercial market by 1.3 cents in the OLS specification (column 1), and 2 cents in the IV specification (column 2).

Columns (3) and (4) report results of validity tests using the pre-Part D period. We test whether insurers’ Part D enrollment is correlated with retail profit changes from 2004 to 2005, prior to the implementation of the program. If the effects reported in columns (1) and (2) are causal, then there should be no such relationship evident. However, if there are differential trends in profits per pill that are systematically correlated with firms’ geographic distributions, these placebo regressions would turn up significant effects. We obtain fairly precise zeroes. The standard errors on the coefficients are even smaller than in our benchmark models, and the estimates are insignificant. This provides evidence against the concern that our main results are driven by long-term trends in price negotiations that happen to be correlated with changes in Part D enrollment.

Results of analogous specifications that include Humana and the additional quality-adjusted premium instrument are discussed in the Empirical Appendix. In these specifications, the estimated enrollment elasticity and predicted effects are similar, if slightly larger.

### **4.3 Enrollment Effects and Retail Profits, by Drug Type**

The theory predicts that results might differ by the degree of bargaining power held by the manufacturer. Recall that if manufacturers hold all the bargaining power, enrollment will have no impact on profits for pharmacies or insurers, as all the rents remain with the manufacturer. To test this hypothesis, we repeat the analysis stratifying the sample by branded and generic drugs. The operative assumption is that manufacturers of generic drugs possess less bargaining power against insurers and pharmacies than manufacturers of branded drugs.

Results are reported in Table 5. Column headings specify the comparison period (relative to the second half of 2005) used in the first difference profit analysis. Columns (1) through (3) report the enrollment effects separately for branded and generic drugs. An additional 100,000 enrollees leads to an approximate 2 cent decline in retail profits per pill observed in the non-Part D commercial market between the second half of 2005 to the first half of 2006. This effect is driven by a 3 cent decline in unit profits earned by the pharmacy on generic drugs. The less-than-one cent decline in profits on branded drugs is statistically insignificant.

Columns (4) through (6) report the falsification test relating enrollment increases to pre-Part D profit changes. We find no evidence of pre-existing trends in the profits earned on either generic or branded drugs.

Theory suggests that pharmacy profits on a particular product are positively correlated with the profits of the corresponding manufacturer. These results suggest, therefore, that Part D health insurers experienced gains at the expense of pharmacies, which lose profits they were previously earning on drugs in competitive categories. We show evidence for generic drugs, which account for over 60% of all claims, and roughly 25% of all total outpatient pharmaceutical expenditures (Berndt and Newhouse, 2010). These results are also consistent with theory that suggests gains in insurer bargaining power versus manufacturers of those drugs.<sup>21</sup>

Earlier research by Duggan and Scott Morton (2010) shows that Part D lowers average revenue per prescription for both branded and generic drugs. Our results indicate that the branded drug effects they estimate are likely to be driven by cash-paying patients switching into insurance and receiving the lower retail prices enjoyed by insurers. There are no additional *unit* price reductions within the set of insured patients which could extend beyond the Part D population. For generic drugs, on the other hand, we find this spillover effect.

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<sup>21</sup> Technically, the theory provides one other alternative explanation for this result, although it is much less direct and requires a number of conditions to hold. The first condition is that insurers with larger Part D enrollment increases also experience larger changes in generic quantity than in branded quantity. The second is that the pharmacy's surplus function is convex. In this case, pharmacy profits will rise by more for generic drugs than for branded drugs. We know of no evidence suggesting that bigger enrollment increases were (or should be) associated with bigger increases in the share of generic drugs purchased by an insurer, but there is no direct evidence to rule this explanation out either.

#### 4.4 Alternative Mechanisms

The buyer size effects reported above may be explained by heterogeneous demand effects. Specifically, new Part D enrollees may have lower demand for drugs than existing commercial enrollees. If so, increases in enrollment due to Part D would have lowered average demand for drugs across an insurer's total book of business, potentially leading to lower negotiated prices and pharmacy unit profits.<sup>22</sup>

We test this hypothesis empirically. We test whether demand for prescription drugs—as measured by total prescription drug spending in 2006—for newly insured Part D enrollees is smaller than for individuals with commercial drug coverage in both 2005 and 2006. Using prescription drug utilization data reported in the MEPS, we find that total spending on drugs in 2006 among newly insured Part D enrollees (i.e. Part D enrollees who did not have drug coverage in 2005) was \$2200 on average, as compared to \$690 for individuals with commercial coverage in 2005 and 2006 ( $p$  value for differences  $< 0.01$ ), and \$840 for adults with commercial coverage in 2005 and 2006 ( $p < 0.01$ ).<sup>23</sup> Hence, contrary to the alternative mechanism, the marginal insured who obtained coverage in response to Medicare Part D on average had significantly greater demand for drugs than the commercially insured population.

The finding is quite intuitive given the channels of enrollment that account for commercial drug coverage. Relatively healthy individuals may choose to hold prescription drug coverage because plans in the commercial market are partially risk-rated, and are often bundled with employer-sponsored health insurance. Moreover, on the margin, group prescription drug coverage is low cost due to the tax exclusion and the employer contribution to premiums. These factors strengthen the incentives for commercial employer-sponsored insurance take-up above and beyond the “unsubsidized” willingness to pay.

### 5. Identifying the Effects of Part D on Drug Expenditures

We estimate the effect of enrollment on negotiated prices paid by insurers. We then use these elasticity estimates to identify aggregate expenditure effects of Part D, and decompose it into effects on consumers inside and outside the program.

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<sup>22</sup> We thank an anonymous referee for pointing out this alternative explanation for the buyer size effects.

<sup>23</sup> Detailed results and additional specifications are available upon request.

## 5.1 Enrollment Effects and Retail Prices

While the marginal enrollment effects on unit profit levels are the most theoretically relevant, the effect of insurer size on prices drives changes in retail expenditures, which are of independent interest to policymakers. Given the available data on prices, an analysis of retail prices permits us to estimate enrollment effects as either elasticities, or predicted changes in expenditures as a percent of baseline given the observed increases in insurer enrollment.

Table 6 reports results from estimating equation (3), the log price equation. Column (1) report results specification that includes all drugs. 100,000 additional Part D enrollees lower prices per pill in the commercial market by 2%. For the mean prescription, given the actual distribution of insurer enrollment increases, Part D is predicted to have lowered overall prices by 7%. The median effect is 8.5%.

Consistent with the profit analysis, the enrollment effect on prices is concentrated in the price negotiations for generic drugs. The enrollment elasticity of 2% decomposes into a 4.6% decline in generic prices (column 3), but only a 0.2% decline in branded prices (column 2). On average, given the distribution of observed enrollment increases, Part D lowered retail prices by 15% for generic drugs, and by less than 1.0% for branded drugs.

Appendix Table 3 reports specifications that define alternative comparison periods. As in the profit analysis, the falsification test using price changes from 2004 to 2005 reveals no evidence of pre-existing price trends correlated with eventual enrollment increases. The table also reports tests of whether the initial enrollment effects on prices are sustained into later periods. The effects appear to persist through April of 2007, the latest month for which data is available. Consistent with the theory and previous empirical estimates, the buyer size effects are almost exclusively contained within the market for generic drugs. The effect of additional enrollment on prices for generics is stable, and even increases slightly over time.

## 5.2 Quantifying the Internal and External Effects of Part D

The first-order effect of Part D implementation on welfare is equal to the change in expenditures, holding quantity fixed. We can decompose this expenditure change into components for enrollees inside and outside the Part D program, as shown in the following equation:

$$(12) \quad \Delta Expend =$$

$$\sum_i(\Delta p_i^{Cash/PartD} \cdot q_{Pre,i}^{Cash}) + \sum_i(\Delta p_i^{\geq 60,Com} \cdot q_{Pre,i}^{\geq 60,Com}) + \sum_i(\Delta p_i^{<60,Com} \cdot q_{Pre,i}^{<60,Com}).$$

The first-order change in expenditures equals the change for Part D seniors who were cash-paying at baseline, plus the change for individuals aged 60 and older who were commercially enrolled at baseline (on whom the analysis in Section 4 is based), and the change for commercially enrolled people under age 60. The first component is the direct or “internal” effect of Part D on the prices paid by Part D enrollees; the last two capture the first-order “external” spillover effects on the commercial market.

In equation (12),  $q_{Pre,i}^j$  represents the total quantity of drug  $i$  purchased by individuals in group  $j$ , prior to Part D implementation.  $\Delta p_i^j$  represents the average change in the price of drug  $i$  due to Part D experienced by individuals in group  $j$ . We operationalize the decomposition in equation (12) by computing the percent reduction in price for each group, multiplied by baseline expenditures, as in:

$$(13) \quad \Delta Expend = Expend_{Pre}^{Cash} \gamma_1 + Expend_{Pre}^{\geq 60,Com} \gamma_2 + Expend_{Pre}^{<60,Com} \gamma_3 .$$

$Expend_{Pre}^j$  represents total drug expenditures among individuals in group  $j$  prior to the implementation in Part D, and  $\gamma_j$  represents the average Part D-related decline in log prices for individuals in group  $j$ . For all groups  $j$ , we estimate  $Expend_{Pre}^j$  from the 2005 Medical Expenditure Panel Survey (MEPS).<sup>24</sup> Our estimate of  $\Delta Expend$  therefore reflects changes in prices, holding quantity consumed at pre-Part D levels, for each group  $j$ .

To estimate  $\gamma_1$ , we use our claims data to calculate that, on average, uninsured cash-paying seniors who enroll in Part D experience an expenditure-weighted 30% reduction in annual drug prices between 2005 and 2006. This number must be scaled down to account for the fact that not every previously uninsured senior enrolled in Part D. According to the MEPS, uninsured seniors who chose to enroll in Part D accounted for 94% of expenditures.<sup>25</sup> We thus estimate the first-

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<sup>24</sup>  $Expend_{Pre}^{Cash}$  is the average outpatient prescription drug expenditure among seniors without health insurance coverage at any period during 2005.  $Expend_{Pre}^{j,Com}$  is the average outpatient prescription drug expenditure among group  $j$  who report having private commercial insurance (or private Medigap) as the usual third-party payer for drugs purchased during 2005.

<sup>25</sup> From the MEPS, we estimate that the fraction of uninsured seniors dropped from 24-percent to 8-percent of seniors between 2005 and 2006. This is virtually identical to estimates based on the Health and Retirement Survey (Levy, et al 1999). We assume that the two-thirds of individuals who enrolled in Part D have the highest expenditures among the baseline uninsured group. This generates an upper bound estimate for the internal effect of Part D on previously uninsured seniors, resulting in lower bound estimate of the relative size of the external effect.

order change in expenditures for this group as  $\gamma_1 = (0.94) * (-0.30) = -28.2\%$ . We apply this reduction to the MEPS baseline expenditures for this group, of  $Expend_{pre}^{Cash} = \$12.6B$ .<sup>26</sup>

Next,  $\gamma_2$  is the average percentage change in price for commercially enrolled individuals aged 60 and over, for whom we observe retail prices in the pharmacy claims. According to column (1) of Table 6, this group experiences a 7.0% price reduction. This estimate must be scaled down to account for the fact that our sample is restricted to Part D insurers, rather than the entire commercially enrolled population. In our data, 40% of claims from this commercial enrollee group are covered by a Part D participating insurer.<sup>27</sup> This results in  $\gamma_2 = 40\%(-0.07) = -2.1\%$ . This percentage reduction is applied to the quantity,  $Expend_{pre}^{\geq 60, Com} = \$42.3B$ . This figure does not include the estimated  $\$5.8B$  attributed to Humana's commercial market. Last,  $\gamma_3$  is the corresponding quantity for commercially enrolled individuals under age 60. Since we do not have claims data to estimate this quantity directly, we assume that—for a given drug—the price reduction for commercially enrolled insureds varies linearly with the share of Medicare beneficiaries consuming that drug. Our own validity tests as well as prior literature support this assumption.<sup>28</sup> From the MEPS we calculate that the Medicare share of drugs consumed by commercial enrollees under age 60 is 0.168. The same figure for all commercial enrollees aged 60 and over is 0.467. Linearity implies a price reduction for the under-60 commercial enrollees of  $\frac{0.168}{0.467}(-0.07) \approx -2.5\%$ . Approximately 40% of the under-60 commercial enrollees are covered by a Part D-participating insurer, according to the pharmacy claims data. This implies an aggregate price decline for the under-60 age-group of  $-(2.5\%) *$

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<sup>26</sup> This figure includes spending by Humana that is internal to Medicare Part D. We compare this to external cost savings associated with Part D-related insurer-size increases. Preferred estimates for the external cost savings do not include the savings contributions from Humana, resulting in a lower bound estimate of the relative size of the external effect.

<sup>27</sup> Note that we make two assumptions when calculating the aggregate cost savings among the commercially insured: 1) that the fraction of claims covered by a Part D participating insurer observed in our data is representative of the national retail market; and 2) that the estimated enrollment elasticity and aggregate price reductions is representative of all commercial retail claims.

<sup>28</sup> From the 2005 MEPS, we calculate that the Medicare share of drugs consumed by commercial insureds ages 60 and over is 0.467. For all commercially insureds aged 60-64, this figure is 0.319—30 percent smaller. Linearity implies that the enrollment elasticity on prices for the 60–64 year old commercial population will be 30 percent smaller than for the 60 and over commercial insureds. This prediction lines-up with the Table 6, where columns (4)–(6) are restricted to ages 60–64. Both the enrollment elasticity of prices and implied decreases in overall drug prices are 30-percent lower than corresponding price effects reported in columns (1)–(3). Linearity is also supported by aggregate data showing that the introduction of Part D was associated with declines in manufacturer revenues as a linear function of each drug's Medicare share (Duggan and Scott-Morton, 2010).

(40%) = -1%. We apply this percentage reduction to  $Expend_{pre}^{<60,Com} = \$56.1B$ . This figure does not include the estimated \$7.6B attributed to Humana's commercial market. Combining these parameter estimates according to our decomposition formula, we find a first-order reduction in drug expenditures of \$5.00 billion, of which we can attribute \$1.45 billion (30%) to the external effect of greater insurer bargaining power from Part D enrollment increases. Note that these savings, equivalent to 3.7% of total drug costs among individuals insured by commercial plans of Part D-participating insurers, are *annual* savings, accruing to the insurer and enrollees in each year after the implementation of Medicare Part D.<sup>29</sup>

Humana's contribution to external cost savings are based on estimates reported in Appendix Table 6, column (1). These estimates imply an annual external cost savings of 2.63B. This is equivalent to a 5.7% decline in drug costs for enrollees Part D participating insurers.<sup>30</sup>

It is not clear how much of this expenditure reduction is retained by insurers and how much flows to the commercially insured. At a minimum, there is evidence of noncompetitive behavior in the group insurance market (Dafny, 2010). Regardless of where the rents end up, they are likely to have significant effects on the distribution of welfare in the market.

On the one hand, if they are passed through to commercial enrollees, the total savings of \$1.45B to \$2.63B would imply that commercial enrollees of insurers that participate in Part D accrued total savings nearly equal to the savings experienced by the newly insured Part D beneficiaries. If retained entirely by insurers, the savings would imply that prescription drug costs for Part D insurers would have fallen by 3.7 to 5.7%, and that the average profitability of their commercial prescription drug insurance operations would have risen by more than 5%.

We arrive at this back-of-the-envelope calculation using loading factors in the individual health insurance market, estimated to range between 25% and 40% (Newhouse, 2002).<sup>31</sup> We assume that average loads for stand-alone prescription drug insurance lie in this (wide) range. If

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<sup>29</sup> The 5.3% cost reduction is an expenditure-weighted average cost reductions estimated for 60-and-over and under-60 commercially enrolled populations of Part D participating insurers.

<sup>30</sup> We estimate  $\gamma_2 = 40\%(-0.083) = -3.32\%$ . This is applied to the quantity  $Expend_{pre}^{\geq 60,Com} = \$48.1B$ . For the under 60 commercial population,  $\gamma_3 = \text{of } \frac{0.168}{0.467}(-0.083) * (40\%) = -1.2\%$ , which is applied to  $Expend_{pre}^{<60,Com} = \$63.7B$ . Combining these parameter estimates according to our decomposition formula, we find a first-order annual reduction in drug expenditures of \$6.18 billion, of which we can attribute \$2.63 billion (43%) to the annual external cost savings to the market.

<sup>31</sup> Loading factors in the large group market are estimated to range between 6% and 10% (Newhouse, 2002); using these numbers would yield much larger calculated effects on profitability.

so, we can calculate the minimum impact on profitability by taking 40% as the loading factor and assuming that the entire load goes to profits. Under these conservative assumptions, a 3.7% reduction in prescription drug costs translates into a 5.2% boost to the average profitability of prescription drug insurance provided by Part D-participating insurers.<sup>32</sup> If, instead, profits were equal to 25% of premium revenue, a 3.7% reduction in drug costs would translate into a 10.4% increase in drug insurance profitability for Part D-participating insurers.

## 6. Conclusions

We present a simple and stylized theoretical model that demonstrates the complexity of the relationship between buyer-size and prices, with an application to the pharmaceutical industry. Even a simple model generates an ambiguous relationship between buyer-size and prices. However, it does demonstrate that greater bargaining power by upstream manufacturers will mute whatever relationship exists. These simple results are meant to build on prior literature exploring the link between buyer-size and pricing. While a number of different mechanisms may be operating to produce this relationship, all have similar and policy-relevant implications for the effects of public health insurance subsidy schemes, which are growing in importance.

In the case of Medicare Part D, publicly subsidized health insurance enrollment tilted bargaining power in favor of participating health insurers. Gains in their negotiating leverage came at the expense of pharmacies and generic drug manufacturers, both of which saw their profits erode. Branded drug manufacturers with more *ex ante* bargaining power seem to have escaped this erosion. The total size of the price reduction in the commercially enrolled marketplace was quite significant in relation to health insurer profitability and in terms of its aggregate value to the commercially enrolled population. Therefore, Part D may have transferred resources from competitive firms to firms with existing bargaining power. It also created substantial spillover effects onto the non-Part D marketplace. It is significant to note that these external effects are present in spite of the theoretical separation between commercial price negotiations and Part D price negotiations; these effects can only exist when Part D is

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<sup>32</sup> The change in profits is given by  $\frac{Premiums - (94.7\%)Drug\ Costs}{Premiums - Drug\ Costs}$ . We then exploit the relationship that  $Drug\ Costs = 60\% * Premiums$ , from the loading factor estimates.

administered through the private insurers with large commercial enrollment external to Medicare.

Our results illustrate the interaction between insurer bargaining power and the competitive pressure faced by manufacturers. For molecules with less competition, insurer enrollment growth is unlikely to make significant price inroads, as manufacturers appear much more of the bargaining power. However, for drugs that have identical molecular equivalents, price negotiation by insurers can have significant benefits for consumers. The optimal degree of competitiveness faced by manufacturers depends on both efficient drug pricing and the provision of sufficient incentives to innovate. Therefore, it is not clear whether policies to reduce manufacturer revenues would harm future welfare by more than they enhance current welfare.

More generally, our findings suggest an important external effect of public subsidies for private health insurance. Direct and indirect subsidies are becoming more prevalent in the US health care system, whether in the form of tax exemption for employer-based health insurance premiums or direct subsidies for insuring the poor. The welfare analysis of such policies must consider the spillover effects created by providing insurers with bargaining power. In our context, those external effects were quite significant relative to the internal price effects of the program. Of note, the landmark Patient Protection and Affordable Care Act extends private health insurance premium subsidies to millions of uninsured. As with Part D, it may represent another context in which spillover effects of public financing of private insurance on provider prices are important.

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## Theoretical Appendix

This appendix generalizes the main results of Chipty and Snyder in the context of 3-way Nash bargaining among a manufacturer, an insurer, and a pharmacy.

### TA-1) Correlated Profits and Mark-ups

A monopolistic manufacturer with varying degrees of bargaining power bargains with a monopolistic pharmacy to set the upstream price of drugs. Downstream, the pharmacy bargains with a set of insurers. For a given drug, pharmacy profit consists of payments received from  $n$  insurers,  $\sum_{i=1}^n \tau_i$ , net of the lump-sum transfer  $T$ , payable by the pharmacy to the manufacturer, for sale of  $Q$  units of a given drug. In general, the payments will depend on the total quantity provided. In addition, the pharmacy may derive other benefits from selling  $Q$  units of drugs. For instance, drug sales may drive traffic to stores and produce sales of other merchandise. The net return to such activity is represented by  $G(Q)$ . In sum, pharmacy profits are given by  $G(Q) + \sum_{i=1}^n \tau_i(Q) - T$ .

Profits of the manufacturer are given by  $T - \sum_{i=1}^n r_i(Q) - C(Q)$ , where  $T$  is the lump-sum pharmacy transfer,  $\sum_{i=1}^n r_i(Q)$  is the total lump-sum rebates paid to insurers as a function of aggregate quantity, and  $C(Q)$  represents the cost of manufacturing and selling the drug.

The outcome of the bilateral negotiation between the manufacturer and the pharmacy maximizes the Nash product:

$$(1) \quad \max_{Q,T} (T - \sum_{i=1}^n r_i(Q) - C(Q))^\gamma (G(Q) + \sum_{i=1}^n \tau_i(Q) - T)^{1-\gamma}$$

The exponent  $\gamma$  captures the bargaining power of the manufacturer, vis-à-vis the pharmacy, in negotiations over lump-sum transfers for a particular drug. It can be interpreted as the share of incremental surplus appropriated by the manufacturer.<sup>33</sup> The polar case where  $\gamma = 1$  is one of

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<sup>33</sup> This parameter is the focus of Ellison and Snyder (2008) who show empirically that the wholesale price of an antibiotic negotiated by manufacturers and pharmacies depends on the substitutability of that antibiotic. Another way to capture bargaining power in this negotiation is to explicitly model the pharmacy's threat point in the expression of its surplus. The threat of non-cooperation with the manufacturer comes from the pharmacy's ability to choose which drugs to stock, and its ability to steer demand among therapeutically substitutable drugs. The pharmacy's bargaining power vis a vis one manufacturer is tied to its outside option—profits earned when steering demand towards a therapeutic substitute. Modeling bargaining power in this way generates the same qualitative results for correlated mark-ups and the impact of increased insurer size as when modeled by Nash exponents in appendix equation (1). Similarly, exponents in the Nash product could be used to capture bargaining power in the rebate negotiations between manufacture and insurers. For the purposes of this model, we can capture bargaining power of a manufacturer through the manufacturer-pharmacy negotiation, although it is trivial to add Nash exponents in the manufacturer-insurer negotiation as well.

complete manufacturer bargaining power, perhaps because it sells a drug that faces no competition and faces perfectly inelastic demand. The opposite case, where  $\gamma = 0$ , obtains perhaps when the manufacturer produces a drug (e.g., a generic) that faces competition from perfect substitutes. This problem has the first-order conditions:

$$(2) \quad C'(Q) + \sum_{i=1}^n r'_i(Q) = G'(Q) + \sum_{i=1}^n \tau'_i(Q)$$

$$T = \gamma G(Q) + (1 - \gamma)C(Q) + (1 - \gamma) \sum_{i=1}^n r_i + \gamma \sum_{i=1}^n \tau_i(Q)$$

Substituting the expression for the equilibrium pharmacy transfer into the two surplus functions gives expressions for the profits of the manufacturer and pharmacy, as a function of aggregate quantity:

$$(3) \quad \Pi_M(Q) = \gamma(G(Q) - C(Q) - \sum_{i=1}^n r_i(Q) + \sum_{i=1}^n \tau_i(Q))$$

$$\Pi_P(Q) = (1 - \gamma) \left( G(Q) - C(Q) - \sum_{i=1}^n r_i(Q) + \sum_{i=1}^n \tau_i(Q) \right)$$

Equation (3) states that the share of total supplier surplus retained by the pharmacy and the manufacturer depends on  $\gamma$ . The equation also implies that any *change* in the total surplus to suppliers – i.e., pharmacies and manufacturers – will lead to positively correlated changes in pharmacy profits ( $\Pi_P$ ) and manufacturer profits  $\Pi_M$ . Changes in unit profits – i.e., profits per unit of quantity – are also positively correlated with changes in unit surplus. A literal interpretation of the model would suggest that any change in the log of profits would be identical for both the manufacturer and the pharmacy.

#### TA-2) Downstream Negotiation and the Impact of Insurer Enrollment Size

Based on the solution to the upstream bargaining problem, the pharmacy bargains simultaneously downstream with each insurer  $i$ . The outcome of each negotiation is a quantity and lump-sum transfer,  $(q_i, \tau_i)$ . Under the Nash framework, each insurer believes that all other insurers are playing optimally, and that it is the marginal insurer in the negotiations with the

pharmacy.<sup>34</sup> The solution to the negotiation maximizes the product of insurer surplus and the incremental profits to the pharmacy of contracting with the insurer. Based on the expression for  $\Pi_P(Q)$  from above, this can be written as:

(4)

$$\max_{q_i, \tau_i} \left(1 - \gamma\right) \left[ \left( G(q_i + \sum_{j \neq i} q_j^*) - C(q_i + \sum_{j \neq i} q_j^*) - (r_i + \sum_{j \neq i} r_j) + (\tau_i + \sum_{j \neq i} \tau_j^*) \right) - \left( G(\sum_{j \neq i} q_j^*) - C(\sum_{j \neq i} q_j^*) - \sum_{j \neq i} r_j + \sum_{j \neq i} \tau_j^* \right) \right] (u(q_i) - \tau_i + r_i)$$

This problem has the following first-order conditions:

$$(5) \quad \begin{aligned} u'(q_i) &= -G'(q_i + \sum_{j \neq i} q_j^*) + C'(q_i + \sum_{j \neq i} q_j^*) \\ (\tau_i - r_i) &= \frac{1}{2} \left[ u(q_i) - \left( G(q_i + \sum_{j \neq i} q_j^*) - G(\sum_{j \neq i} q_j^*) \right) + \left( C(q_i + \sum_{j \neq i} q_j^*) - C(\sum_{j \neq i} q_j^*) \right) \right] \end{aligned}$$

The manufacturer bargains separately but simultaneously with each insurer  $i$ . The outcome of each negotiation is a quantity and lump-sum rebate,  $(q_i, r_i)$ . This problem has first-order conditions identical to those in the pharmacy-insurer negotiation, implying that separate expressions for equilibrium  $\tau_i$  and  $r_i$  cannot be derived. Instead, the solution consists of a unique set of net prices and profits.

There are a number of ways to conceptualize an increase in enrollment for an insurer. To economize on notation, we implement it as an amalgamation of two existing insurers,  $h$  and  $i$ .<sup>35</sup> The total gross surplus earned by this combined insurer is equal to  $v(q_i^m + q_h^m) = u(q_i^m) + u(q_h^m)$ , while the total tariff paid by the merged insurer is denoted as  $\tau_{hi}$ , and the rebate received by the merged insurer denoted as  $r_{hi}$ . The linearity in the combined insurer's gross surplus function implies that enrolling in a larger insurer confers no benefit to an insured, holding quantity constant. The combined insurer bargains with the pharmacy according to:

<sup>34</sup> One could enrich this model by specifying it as an extensive-form game in which there is a set of probabilities that other players' negotiations break down. Chipty and Snyder (1999) note that the Nash-bargaining approach leads to a limiting perfect Bayesian equilibrium of the extensive-form game, in which the probability of breakdown approaches zero. Practically speaking, the Nash framework is both simple and likely relevant to the pharmaceutical context, where negotiations rarely break down entirely between the players.

<sup>35</sup> The analysis can easily be adapted to the case of uninsured consumers joining an insurer, but at the cost of some additional notation.

$$(6) \quad \max_{q_h, q_i, \tau_{ih}} \{G(q_i^m + q_h^m + \sum_{j \neq i, h} q_j^{m*}) - C(q_i^m + q_h^m + \sum_{j \neq i, h} q_j^{m*}) - (r_{ih} + \sum_{j \neq i, h} r_j^m) + (\tau_i h + \sum_{j \neq i, h} \tau_j^{m*})\} - \{G(\sum_{j \neq i, h} q_j^{m*}) - C(\sum_{j \neq i, h} q_j^{m*}) - \sum_{j \neq i, h} r_j^m + \sum_{j \neq i, h} \tau_j^{m*}\} * \{u(q_i^m) + u(q_h^m) - \tau_{ih} + r_{ih}\}$$

This problem has the first-order conditions:

$$(7) \quad \begin{aligned} u'(q_i^m) = u'(q_h^m) = C'(q_i^m + q_h^m + \sum_{j \neq i, h} q_j^m) - G'(q_i^m + q_h^m + \sum_{j \neq i, h} q_j^m) \\ (\tau_{ih} - r_{ih}) = \frac{1}{2} \left[ u(q_i^m) + u(q_h^m) - \left( G(q_i^m + q_h^m + \sum_{j \neq i} q_j^{m*}) - G(\sum_{j \neq i} q_j^{m*}) \right) + \left( C(q_i^m + q_h^m + \sum_{j \neq i} q_j^{m*}) - C(\sum_{j \neq i} q_j^{m*}) \right) \right] \end{aligned}$$

For all insurers  $k$ , the first-order conditions for  $q_k^m$  are identical to the corresponding conditions for  $q_k^s$ , the equilibrium quantity for firm  $k$  in the unmerged, separated insurer environment. Therefore, it follows that  $q_k^m = q_k^s$ , for all  $k$ . This allows us to suppress the superscripts on the quantity variables for the rest of this section. Defining the net surplus function,  $J(Q) \equiv C(Q) - G(Q)$ , we can now write the difference in the net price paid by insurers in the merged and unmerged settings as:

$$(8) \quad \begin{aligned} (\tau_{ih} - r_{ih}) - [(\tau_i - r_i) + (\tau_h - r_h)] = \\ \frac{1}{2} \{ [J(Q) - J(Q - q_i - q_h)] - [(J(Q) - J(Q - q_i)) + (J(Q) - J(Q - q_h))] \} \end{aligned}$$

This expression demonstrates “Implication 1,” that the impact of insurer size on the pharmacy profit function is ambiguous. The net price paid by the merged insurers lies strictly below that paid by the disintegrated firms if  $J(Q)$  is strictly convex in  $Q$ , but not otherwise. Next, equation (3) demonstrates that – for a fixed aggregate quantity  $Q$  – the pharmacy profit function rises if and only if the sum of net prices paid by insurers rises. Therefore, the profit function of the pharmacy strictly falls if  $J(Q)$  is strictly convex in  $Q$ . The relationship between changes in pharmacy profits and changes in insurer net prices is important empirically, because changes in pharmacy profits are more easily observed than changes in insurer net prices. The meaning and importance of this point is explained later, when we discuss the link between observable variation in pharmacy prices and variation in pharmacy profits.

Note finally that the results are identical for the unit profit function of the pharmacy. To see this, observe that dividing both sides of equation (8) by  $q_i + q_h$  reveals that insurer size has an

ambiguous impact on net unit prices paid by insurers; moreover, equation (3) implies that – for a given level of aggregate quantity – the unit profit function of the pharmacy rises if and only if the average net unit prices paid by insurers rise as well.

This result has a number of corollaries, which make clear the theoretical ambiguity of this prediction. First, increased insurer size lowers insurer net prices and pharmacy profits if  $G$  is strictly concave, and  $C$  is weakly convex. Alternatively, if  $G$  and  $C$  are both linear, size increase has no impact on net prices or pharmacy profits. If  $J$  is strictly concave—e.g., due to increasing returns in the manufacture of pharmaceuticals—increased payer size actually leads to higher net prices and pharmacy profits. The effects of insurer size on net prices and profits may be non-monotonic and depend on the curvature of the surplus functions at the margin. These results are analogous to the conditions derived by Chipty and Snyder (1999) for the single-seller model.

Now recall “Implication 2,” which states that the effect of insurer size on the pharmacy profit function (and unit profit function) is smaller in absolute value when the manufacturer has more bargaining power. This result follows from the expressions in equation (3), coupled with the results we have shown for insurer size. The logic proceeds as follows:

- (1) Insurer size has some effect on the net prices paid by insurers to the pharmacy.
- (2) Equation (3) shows that the resulting change in supplier surplus – i.e., the surplus accruing jointly to the manufacturer and pharmacy – is exactly equal to the total change in net prices paid.
- (3) Equation (3) also shows that, when manufacturers hold more of the bargaining power, the pharmacy receives a smaller share of any change in the supplier surplus.
- (4) Therefore, the effect of insurer size on pharmacy profits (i.e., the pharmacy’s share of supplier surplus) is smaller in absolute value, when manufacturers hold more of the bargaining power.

Finally, note that, since every step in this argument also holds true for the unit profit function, “Implication 2” holds for both profits and unit profits. Intuitively, when manufacturers have a large degree of bargaining power—e.g., for branded drugs with few substitutes and highly inelastic demand—the pharmacy surplus is small, as is the effect of insurer size on the

pharmacy's surplus and price markup. In these cases, a given change in insurer size will have little effect on pharmacy prices or profits. Implications 1 and 2 motivate the empirical analysis.

As discussed above, equation (3) also generates Implication 3. The profits of the pharmacy and the manufacturer are both proportional to total supplier surplus and will thus co-vary positively with buyer size. The same is true for unit profits, because unit profits of both the pharmacy and manufacturer are proportional to the unit supplier surplus (i.e., supplier surplus per unit of quantity). While we do not test this implication directly, we can use this result to infer the qualitative direction of changes in manufacturer profits from measured changes in pharmacy profits in response to increase in insurer size.

#### TA-3) Measurement of pharmacy unit profits

The final point to be demonstrated is that changes in the pharmacy's unit profits can be measured using data on the unit prices insurers pay to the pharmacy. First, note that pharmacy profits are defined as:  $G(Q) + \sum_{i=1}^n \tau_i(Q) - T$ . In words, profits consist of: the indirect "foot traffic" value of selling  $Q$  drugs,  $G(Q)$ ; plus the prices paid by insurers to the pharmacy; net of the lump-sum transfer paid by the pharmacy to the manufacturer. Similarly, the pharmacy's unit profits are defined as:  $\frac{G(Q) + \sum_{i=1}^n \tau_i(Q) - T}{Q}$ .

Now suppose we observe the prices paid by each insurer  $i$  for drug  $d$ . The unit profits earned by the pharmacy from each unit sold by insurer  $i$  are given by:  $\frac{G(Q) + \tau_i(Q) - T}{Q}$ . The difference in unit profits earned by the pharmacy across insurers  $i$  and  $j$  is thus given by:  $\frac{\tau_i(Q)}{Q} - \frac{\tau_j(Q)}{Q}$ . Note that this is simply the difference in unit prices paid from the two insurers to the pharmacy. The latter quantity is observable in pharmacy claims data, even though net insurer prices and overall pharmacy profits are not directly observable.

Note this argument rests in part on the institutional detail that rebates to or from pharmacies are rarely, if ever, observed in the marketplace. Instead, rebate arrangements are negotiated between insurers and manufacturers.<sup>36</sup> Pharmacy rebates would add another unobservable term to the expression for the difference in unit profits earned by the pharmacy across insurers  $i$  and  $j$ .

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<sup>36</sup> Department of Health and Human Services (2000) provides a useful primer on drug pricing institutions.

## Empirical Appendix

### EA-1) Specifying the First-Stage Equation with Humana

The empirical appendix incorporates Humana into the analysis. Humana's enormous enrollment in response to its extreme loss-leader pricing strategy implies that the potential instrument does not locally predict Humana enrollment. Humana's inclusion in the analysis requires a second instrument to predict Part D enrollment. We propose an instrument that is directly motivated by Humana's under-pricing strategy: the quality-adjusted plan premium. This variable can be thought of as insurers' load price of Part D coverage, controlling for plan characteristics and generosity.

The unique regulatory features of Part D allow us to restrict the sample to identical and actuarially equivalent plans sold in the same CMS-defined Part D market, thereby allowing us to purge observed premiums of benefit design and plan generosity, so as to isolate pure premium pricing variation.<sup>37</sup> We estimate equation (10), as discussed in Section 3. The estimated insurer fixed-effects,  $\sigma_i$ , represent the generosity- or quality-adjusted premium charged by insurer  $i$  relative to premiums for identical and actuarially plans sold in the same market.

The issue of whether the quality-adjusted premium measure is uncorrelated with expected changes in retail drug prices, before and after the implementation of Part D, is a key issue for validity. In theory, premiums reflect plan design and utilization, drug prices, administrative insurer costs, and premium pricing strategies related to market share. The most obvious threat to the premium instrument is the potential for premiums to reflect changes in prices anticipated by the insurer. Recall that insurers set premiums for their 2006 plans by open enrollment in mid-2005. It is not problematic for the premium instrument to reflect contemporaneous 2005 drug price levels, or even realized prices in 2006. Rather, violation of the exclusion restriction requires a more nuanced correlation, namely that the premium load instrument is correlated with *changes* in drug prices between 2005 and 2006 through channels other than their Part D enrollment. One possible "cost-based" mechanism may be that premiums not only reflect 2005

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<sup>37</sup> CMS defines a standard minimum benefit coverage plan. The majority of Part D plans offered by private insurers are either standard plans, or plans that are actuarially equivalent to the standard plan. Equation (10) is estimated using region fixed effects, and is restricted to standard and actuarially equivalent plans, greatly reducing biases due to unobserved plan characteristics. "Premium plans" that offer greater levels of coverage are omitted from estimation of equation (10). Furthermore, reimbursements to Part D insurers are tied to plan type and benefit design, which necessitates CMS collecting and reporting detailed information about benefits design and generosity.

prices, but also capture negotiated drug prices that insurers expect to pay in 2006. Another possibility is that insurers that aggressively price plans in 2005 to gain market share also more aggressively negotiate unit drug prices in 2006 *relative to their negotiations in 2005*. While we cannot test this latter mechanism directly, we provide direct evidence against anticipatory cost-based premium pricing.

Appendix Table 4 reports regressions where premiums for basic and actuarially equivalent Part D plans are regressed on plan and insurer characteristics, as well as an insurer-level index of negotiated drug prices. Plan characteristics are strong predictors of plan premiums in every specification. Column (2) controls for 2005 drug prices. The 2005 drug price index is marginally significant ( $p = 0.17$ ), and its inclusion increases the R-squared by 10 percent. Column (3) controls for the 2006 drug price index. The coefficient is economically small and statistically insignificant, and controlling for it leaves the R-squared unchanged. Similarly, the change in drug prices between 2005 and 2006 (column 4) is unrelated to premiums. Taken together, the results imply that premiums are largely explained by plan generosity, and to a lesser extent, contemporaneous drug prices. We reject the class of validity threats that imply a direct relationship between premiums and future prices or price changes.

#### EA-2) Specifying the Second-Stage Equation with Humana

The theoretical model does not specify the functional form that obtains between pharmacy profits or prices and enrollment. The data offer guidance on this issue. We calculate residuals from the change in log drug price (equation (3)), including all covariates except the key explanatory variable, insurers' Part D enrollment. Figure 6a and 6b plots the residuals from this regression against insurers' Part D enrollment.

Figure 6a shows this relationship for all insurers excluding Humana. The relationship is clearly negative, an apparently linear relationship that bears out in the regression analysis. Figure 6b shows the relationship for all insurers including Humana. A downward-sloping relationship is visually evident. However, this linear relationship diminishes—nearly to zero—given the enormous enrollment increase achieved by Humana. Humana's presence may be identifying a diminishing effect of enrollment on negotiated prices. If so, failing to account for the potential non-linearity will produce misleading linear estimates of the enrollment effect. Therefore, for specifications that include Humana, we allow for a quadratic in Part D enrollment.

Our estimated enrollment elasticity, and predicted price changes in response to enrollment increases, are nearly identical across the linear and non-linear model specifications. Appendix Table 5 reports the non-linear effect of enrollment on retail profits per pill, analogous to the linear results in Table 5. Appendix Table 6 reports the non-linear effect of enrollment on retail prices, analogous to the linear results in Table 6. In both all cases, estimated effects are quantitatively similar, if not larger, in the non-linear specifications that include Humana.

Figure 1. Transfers and Payments in the Prescription Drug Market

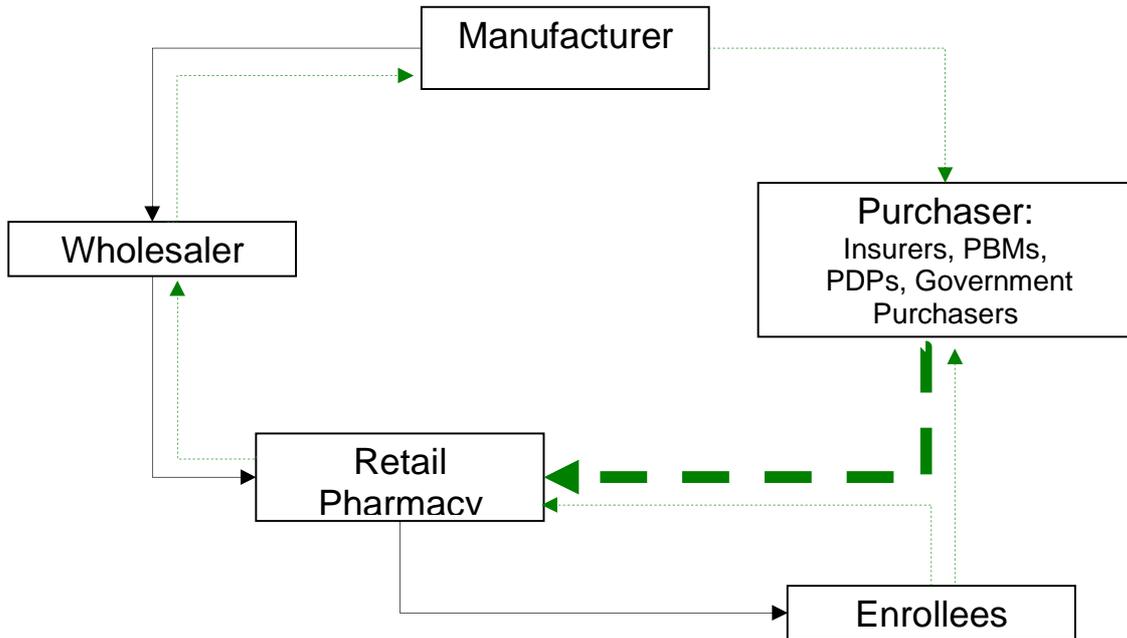


Figure 1 summarizes the flow of product and payments in the prescription drug market. The solid lines represent the flow of drugs; the dotted lines represent the flow of payments. The bold dotted line represents the retail drug prices negotiated between retail pharmacies and insurers.

Figure 2. Buyer size and seller surplus functions

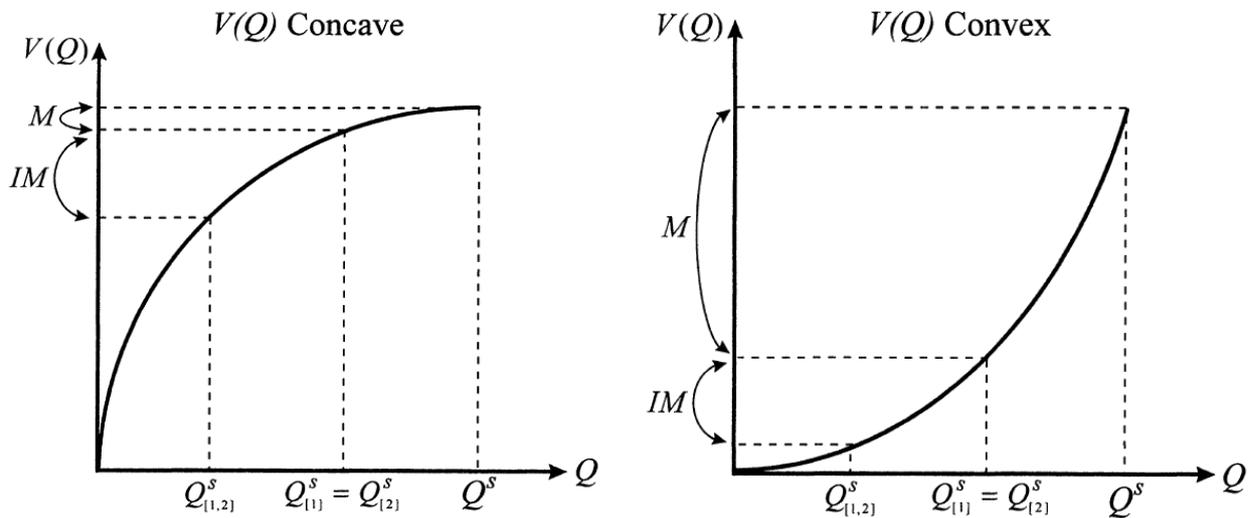
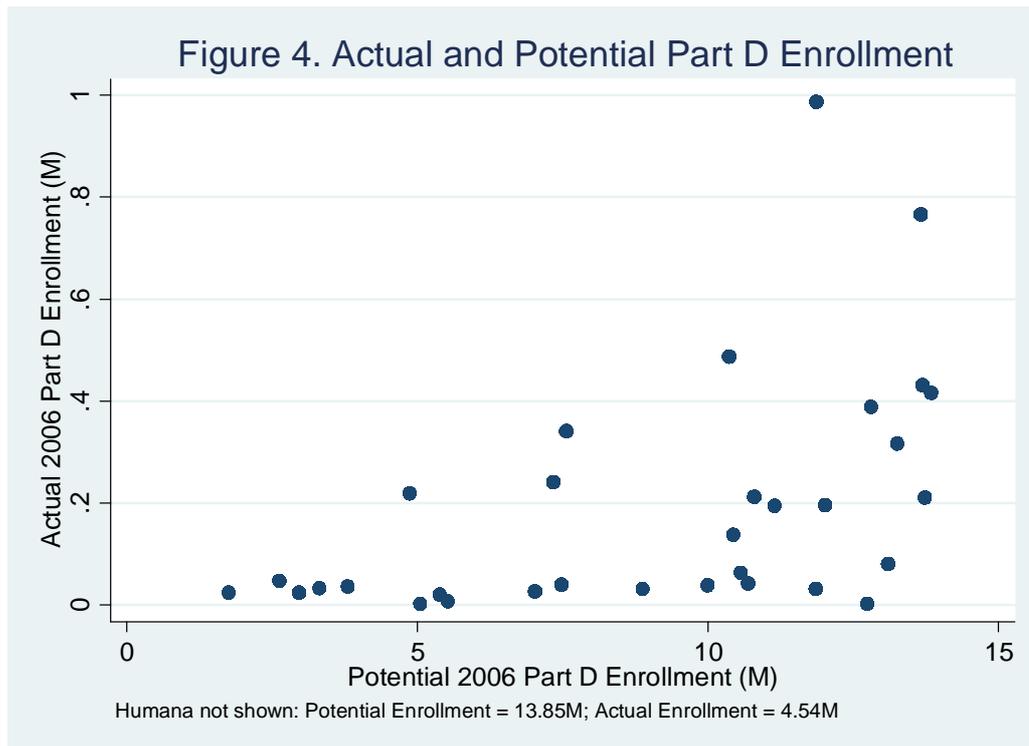


Figure 2, reproduced from Chipty and Snyder (1999), shows the surplus value,  $V(Q)$ , created by equal-sized inframarginal and marginal buyers, as a function of total quantity sold. All buyers in this framework bargain as if they are the marginal firm. When surplus is concave, a larger buyer (accounting for  $Q^s - Q_{[1,2]}^s$  of volume) generates larger unit surplus,  $(M + IM) / (Q^s - Q_{[1,2]}^s)$  than unit surplus generated by a smaller unmerged marginal buyer,  $M / (Q^s - Q_{[2]}^s)$ . The opposite is true for convex surplus function. That all buyers bargain as the marginal buyer is the limiting case of a more general extensive form game, where the limit is taken as the probability of negotiating breakdowns approaches zero.



Note: Figure 3a shows the distribution of retail prices of drugs across insurers, where price is measured as the percentage difference between the retail price of an NDC-level drug negotiated by an insurer and the average retail price of that drug across all insurers. In Figure 3b, price is measured as the absolute difference between the retail price of a given NDC-level drug and the average price of that drug across all insurers. Data come from one point in time (September 2005) so that variation in both figures, for a given NDC, comes from variation in negotiated prices across insurers.



Note: Figure 4 plots insurers' Part D enrollment against their *Potential Part D enrollment*, a measure of each insurer's exposure to the previously uninsured Part D-eligible population according to its geographic presence in the commercial insurance market prior to Part D.

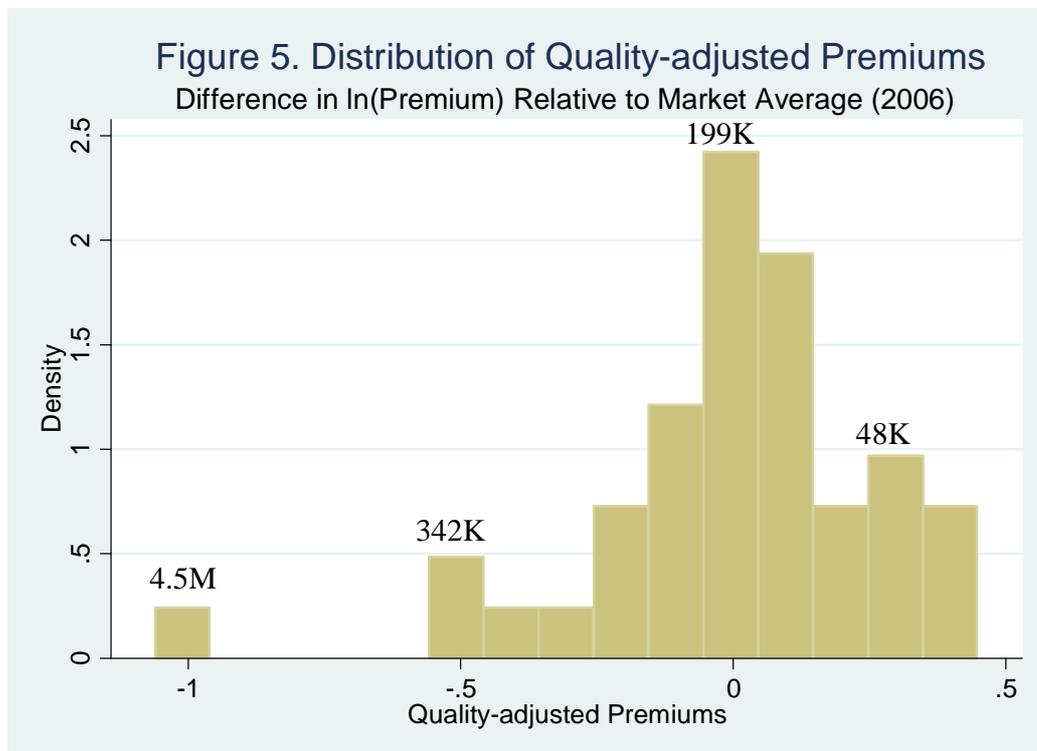


Figure 5 shows the distribution of insurers' *Quality-adjusted premium*, calculated as the difference in log premiums between across standardized Part D plans sold in the same market, controlling for plan generosity and drug costs. Average Part D enrollment across insurers within a bin is labeled above key points in the distribution.

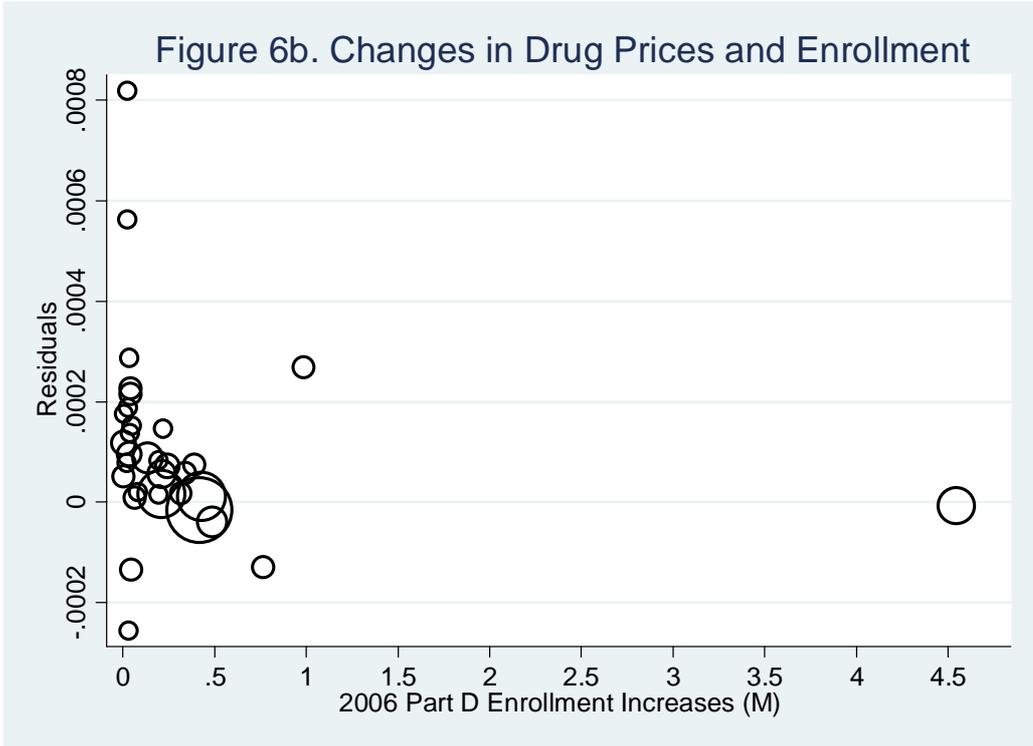
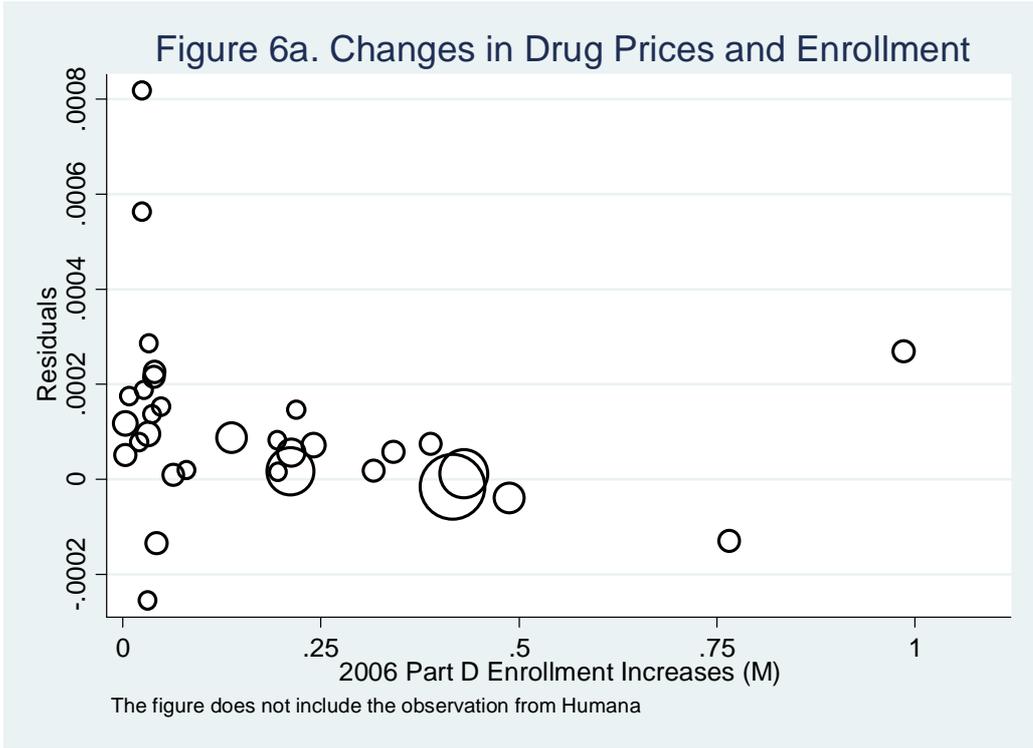


Figure 6a and 6b plot residual changes in log drug prices against insurer Part D enrollment increases. We regress changes in log drug prices between 2005 and 2006 against all covariates except for Part D enrollment increases, then averaging the residuals of the regression across all drugs at the insurer level. Circle size reflects the size of insurers, measured by the number of claims observed in the data. Figure 6a plots these residuals for all insurers excluding Humana. Figure 6b displays the same figure as in Figure 6a including Humana.

Table 1. Sales Rank of Top Selling Drugs in the Pharmacy Claims Data and MEPS

Rank	Drug	MEPS Rank
1	LIPITOR	1
2	PLAVIX	4
3	ZOCOR	2
4	NORVASC	6
5	PREVACID	10
6	NEXIUM	5
7	FOSAMAX	8
8	ADVAIR	16
9	PROTONIX	15
10	PRAVACHOL	11
11	DIOVAN	14
12	ACTOS	17
13	CELEBREX	12
14	TOPROL XL	7
15	AVANDIA	18
16	COREG	21
17	AMBIEN	58
18	ARICEPT	20
19	ACTONEL	31
20	LEVAQUIN	98
21	ZETIA	19
22	ZOLOFT	22
23	FLOMAX	33
24	ACIPHEX	34
25	COSAAR	40

Table 1 lists the top 25 drugs, ranked by expenditures in the 2004 and 2005 pharmacy claims for individuals aged 60 and higher, and their corresponding rank in the 2004 and 2005 Medical Expenditure Panel Survey (MEPS) Prescription Medicines modules. Expenditures are measured as the sum of payments to the pharmacy made by the customer plus third party payers. High ranking drugs in the MEPS that do not appear in the pharmacy claims data are physician administered, and are therefore less likely to appear in out-patient pharmacy claims. For example, the two highest ranking omissions are Procrit (rank #3 in the MEPS, rank #79 in the pharmacy claims) and Atenolol (#9 in MEPS, #62 in pharmacy claims), are mainly physician administered.

Table 2. Distribution of Insurers by 2007 Part D Enrollment

	Below Median	Above Median	50-75th Percentile	75-90th Percentile	90-95th Percentile	Above 95th Percentile
	(1)	(2)	(3)	(4)	(5)	(6)
Insurer's Part D Enrollment	< 6,400	> 6,400	6,400-28,000	28,000- 126,000	126,000- 354,000	> 354,000
Number of Part D Insurers	124	124	62	36	13	13
Number of Insurers Appearing in Claims	15	71	29	21	9	12
Fraction of Insurers Appearing in Claims	0.22	0.87	0.46	0.54	0.70	0.95

Table 2 shows the distribution of insurers by their Part D enrollment. For each enrollment bin, the table reports the number insurers participating in Part D reported by CMS; the number of these insurers that appear in the pharmacy claims; and the fraction of Part D participating insurers that appear in the claims, where each insurer is weighted by their Part D enrollment.

Table 3. Illustration of the Potential Part D Enrollment Instrument

	Total Expenditures in 2005	Potential Part D Enrollment	Actual 2006 Part D Enrollment
	(1)	(2)	(3)
Insurer A	\$4.6M	12.6M	986,108
Insurer B	\$5.4M	3.4M	20,735
Insurer C	\$6.8M	3.3M	37,388
Insurer D	\$7.0M	8.2M	221,359

Table 3 illustrates the explanatory power of the *Potential Part D enrollment* instrument. *Potential Part D enrollment* is an insurer-level variable defined as the number of seniors in 2005 without private drug insurance (including Medicaid recipients who receive coverage under Part D) residing in states in which the insurer is present in the commercial market, weighted by the insurer's commercial market share in those states. Data on insurance coverage come from the 2005 Current Population Survey. Column (1) reports the commercial market size, as measured by the total reimbursements to the pharmacy, of four similarly-sized insurers. Column (2) reports the *Potential Part D enrollment* for these four insurers. Actual Part D enrollment in 2006 is reported in column (3).

Table 4. Linear Enrollment Effect on Retail Drug Profits-per-Pill

Dependent Variable: $\Delta(\text{Drug Price-per-Pill})$				
Period	Second Half 2005 vs First Half 2006		Second Half 2004 vs First Half 2005	
Model	OLS	IV	OLS	IV
	(1)	(2)	(3)	(4)
$\Delta$ Firm's PartD Enrollment (1M)	-0.126** (0.052)	-0.191** (0.072)	0.034 (0.020)	0.024 (0.027)
$\Delta$ Log Exposure to Pharmacy	-0.301 (0.205)	-0.338 (0.218)	-0.013 (0.027)	-0.016 (0.028)
$\Delta$ Avg Quantity per Rx/100	-0.030*** (0.005)	-0.031*** (0.005)	-0.066*** (0.014)	-0.066*** (0.013)
	First Stage			
<i>Excluded Instruments</i>		$\Delta$ Enrollment		$\Delta$ Enrollment
Potential Enrollment (1M)		0.035*** (0.008)		0.037*** (0.008)
F-stat for Excluded Variables		16.00		16.89
Average Baseline Unit Price	4.15	4.15	3.98	3.98
NDC-level Drug Fixed Effects	Y	Y	Y	Y
Insurer Observations	32	32	32	32
Insurer-Drug Observations	9520	9,474	9,248	9,211

Table 4 reports the effect insurers' Part D enrollment on their unit profits earned in the commercial non-Part D market. The dependent variable is the change in insurer-NDC drug level average price-per-pill (\$) between the second half of 2005 and the first half of 2006. The key regressor is the change in insurers' Part D enrollment between 2005 and 2006. As discussed in the text, including NDC drug fixed effects allows for the coefficient on enrollment to be interpreted as the enrollment elasticity of unit profits. Covariates include changes in the average number of pills per prescription and changes in each insurer's exposure to the pharmacy (the state-level market share of the pharmacy, weighted by the fraction of the insurer's total retail expenditures in each state). The sample of drugs comprises the top 1000 drugs ranked by expenditures observed in the claims. The instrument for Part D enrollment is potential Part D enrollment. In all specifications, insurer-NDC-level observations are weighted by the number of claims for the NDC observed for that insurer. Parentheses report standard errors clustered at the insurer level. \* significant at 10%; \*\* significant at 5%; \*\*\* significant at 1%

Table 5. Enrollment Effect on Retail Drug Profits-per-Pill, by Branded/Generic Status

		Dependent Variable: $\Delta$ (Drug Price-per-Pill)					
Comparison Period		First Half 2006			Second Half 2004		
Drug Sample		All	Branded	Generics	All	Branded	Generics
		(1)	(2)	(3)	(4)	(5)	(6)
$\Delta$ Firm's PartD Enrollment (1M)		-0.191** (0.072)	-0.081 (0.075)	-0.297*** (0.090)	0.024 (0.027)	0.049 (0.045)	-0.003 (0.039)
NDC-level Drug Fixed Effects		Y	Y	Y	Y	Y	Y
Average Baseline Unit Price		4.15	7.32	0.71	3.98	6.76	0.60
Insurer Observations		32	32	32	32	32	32
Insurer-Drug Observations		9,474	5,135	4,339	9,211	5,009	4,202

Table 5 reports instrumental variables estimates of the effect of insurers' Part D enrollment on their unit profits earned in the commercial non-Part D market, by branded and generic drug status. The dependent variable is the change in the insurer-NDC drug level average price-per-pill (\$) between the second half of 2005 and the first half of 2006. The key regressor is the change in the insurer's Part D enrollment between 2005 and 2006. As derived in the text, including NDC-level drug fixed effects allows for the coefficients on enrollment to be interpreted as an enrollment elasticity of unit profits. Covariates include changes in the average number of pills per prescription and changes in each insurer's exposure to the pharmacy (the state-level market share of the pharmacy, weighted by the fraction of the insurer's total retail expenditures in each state). The sample of drugs comprises the top 1000 drugs ranked by expenditures observed in the claims. The instrument for Part D enrollment is potential Part D enrollment. In all specifications, insurer-NDC-level observations are weighted by the number of claims for the NDC observed for that insurer. Parentheses report standard errors clustered at the insurer level. \* significant at 10%; \*\* significant at 5%; \*\*\* significant at 1%

Table 6. Enrollment Effect on Pharmacy Prices-per-Pill, by Age Group

Population Model Drug Sample	Dependent Variable: $\Delta \ln$ (Drug Price-per-Pill)					
	Commercially Insured Ages 60 and Over			Commercially Insured Ages 60-64		
	IV	IV	IV	IV	IV	IV
	All	Branded	Generics	All	Branded	Generics
	(1)	(2)	(3)	(4)	(5)	(6)
$\Delta$ Firm's PartD Enrollment (1M)	-0.203*** (0.056)	-0.019 (0.014)	-0.455*** (0.115)	-0.133** (0.064)	-0.023* (0.012)	-0.293** (0.129)
Predicted $\Delta \ln$ (Drug Profit/Pill): Median	-0.085	-0.008	-0.190	-0.055	-0.010	-0.122
Overall Predicted % $\Delta$ in Expenditures	-0.070	-0.007	-0.150	-0.045	-0.008	-0.100
Insurer Observations	32	32	32	32	32	32
Insurer-Drug Observations	9,520	5,172	4,348	6,259	3,430	2,829

Table 6 reports instrumental variables estimates of the effect of insurers' Part D enrollment on their unit profits earned in the commercial non-Part D market, by age group. Price changes are measured as the difference in the log retail price-per-pill, averaged at the insurer-NDC drug level, between the second half of 2005 and the comparison period noted in column headings. The key regressor is the change in the insurer's Part D enrollment between 2005 and 2006. Covariates include changes in the average number of pills per prescription, changes in the average per-pill wholesale drug price, and changes in each insurer's exposure to the pharmacy (the state-level market share of the pharmacy, weighted by the fraction of the insurer's total retail expenditures in each state). The sample of drugs comprises the top 1000 drugs ranked by expenditures observed in the claims. The instrument for Part D enrollment is potential Part D enrollment. Changes in log prices predicted by the model are reported for a) the insurer at the 50th percentile of enrollment increases; and 2) for the market average, given the distribution of observed enrollment increases. In all specifications, insurer-NDC-level observations are weighted by the number of claims for the NDC observed for that insurer. Parentheses report standard errors clustered at the insurer level. \* significant at 10%; \*\* significant at 5%; \*\*\* significant at 1%

Appendix Table 1. Pre-Part D Trends in Drug Prices by Insurer Size

	Dependent Variable: $\Delta$ (Drug Price-per-Pill)				
	(1)	(2)	(3)	(4)	(5)
Insurer's Total Number of Claims	0.0051 (0.0082)				
Insurer's Log Total Number of Claims		0.0002 (0.0007)			
Insurer's Total Rx Expenditures (\$M)			0.0001 (0.0001)		
Insurer's Log Total Rx Expenditures				0.0004 (0.0007)	
Medium Insurer					-0.0049 (0.0038)
Large Insurer					-0.0035 (0.0029)
$\Delta$ Log Avg Quantity per Rx	-0.0590*** (0.0139)	-0.0592*** (0.0139)	-0.0590*** (0.0139)	-0.0592*** (0.0139)	-0.0593*** (0.0138)
$\Delta$ Log Exposure to Pharmacy	-0.0074 (0.0174)	-0.0055 (0.0166)	-0.0078 (0.0174)	-0.0067 (0.0170)	-0.0045 (0.0136)
$\Delta$ Log AWP of Drug	0.1476*** (0.0079)	0.1476*** (0.0079)	0.1476*** (0.0079)	0.1476*** (0.0079)	0.1476*** (0.0079)
Constant	0.0323*** (0.0030)	0.0310*** (0.0090)	0.0322*** (0.0029)	0.0277** (0.0116)	0.0370*** (0.0025)
R-squared	0.031	0.031	0.031	0.031	0.031
Number of Insurers	33	33	33	33	33
Insurer-Drug Observations	9975	9975	9975	9975	9975

Appendix Table 1 tests whether trends in negotiated retail drug prices prior to the implementation of Part D differed according to the size of insurers. In columns (1) and (2), insurer size is measured as the total number of prescriptions observed in the pharmacy claims in 2005 for that insurer. In columns (3) and (4), insurer size is measured as the total expenditures for all prescriptions observed in the claims in 2005 for that insurer. In column (5), the insurer sample is partitioned into terciles based on the expenditure measure. The indicator for the smallest insurer size is the omitted insurer category. The dependent variable is the change in the log average negotiated price per pill paid to the pharmacy on a given drug by a given insurer between the second half of 2004 and the second half of 2005. Controls include the change in the average per-pill wholesale price of the drug, the change in each insurer's exposure to the pharmacy (the state-level market share of the pharmacy, weighted by the fraction of the insurer's total retail expenditures in each state), and the change in the log average wholesale price of a drug. The sample of drugs comprises the top 1000 drugs, ranked by expenditures, observed in the claims. Parentheses report standard errors clustered at the insurer level. \* significant at 10%; \*\* significant at 5%; \*\*\* significant at 1%

Appendix Table 2. Relationship between Pharmacy's Market Power and Insurer Enrollment Size

Dependent Variable	$\Delta \ln$ (Insurer's Exposure to Pharmacy's Market Share)						
	2004		2006		2007		
	Comparison Period	Model	OLS	IV	OLS	IV	
		(1)	(2)	(5)	(6)	(7)	(8)
$\Delta$ Firm's PartD Enrollment (1M)		0.002 (0.002)	0.001 (0.003)	0.002 (0.001)	0.002 (0.002)	0.004 (0.002)	0.003 (0.003)
Constant		0.063*** (0.010)	0.063*** (0.009)	-0.061*** (0.004)	-0.061*** (0.004)	0.002 (0.011)	0.003 (0.011)
F-Stat			66.03		66.03		66.03
Observations		33	33	33	33	33	33
R-squared		0.002	0.002	0.014	0.014	0.007	0.007

Appendix Table 2 reports estimates of the effect of a change in insurers' Part D enrollment on the change in their exposure to the national pharmacy. Exposure is an insurer-level measure of the state-level market share of the pharmacy, weighted by the fraction of the insurer's total retail expenditures in each state. Changes in the exposure measure are defined between 2005 and the comparison year noted in the column heading. The instrument for Part D enrollment in columns (2), (4) and (6) is potential Part D enrollment, described in detail in the text. Parentheses report standard errors clustered at the insurer level. \* significant at 10%; \*\* significant at 5%; \*\*\* significant at 1%

Appendix Table 3. Enrollment Effect on Retail Drug Price-per-Pill, by Branded and Generic Status.

Dependent Variable	$\Delta \ln$ (Drug Price-per-Pill)							
	Panel A: Linear Enrollment Effects							
	Comparison Period	Second Half 2004		First Half 2006		Second Half 2006		First Third 2007
Drug Sample	Branded	Generics	Branded	Generics	Branded	Generics	Branded	Generics
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$\Delta$ Firm's PartD Enrollment (1M)	0.011 (0.007)	0.008 (0.060)	-0.019 (0.014)	-0.455*** (0.115)	-0.001 (0.018)	-0.548*** (0.150)	0.007 (0.018)	-0.585*** (0.145)
Predicted Median $\Delta \ln$ (Drug Price/Pill)	0.005	0.003	-0.008	-0.190	0.000	-0.236	0.003	-0.252
Predicted % $\Delta$ in Total Expenditures	0.007	0.003	-0.007	-0.150	0.000	-0.178	-0.003	-0.190
Insurer Observations	32	32	32	32	32	32	32	32
Insurer-Drug Observations	5,043	4,205	5,172	4,348	4,631	3,789	4,299	3,134

Appendix Table 3 reports instrumental variables estimates of the effect of insurers' Part D enrollment on log unit drug prices in the commercial non-Part D market, by branded and generic drugs. Price changes are measured as the difference in the log retail price-per-pill, averaged at the insurer-NDC level, between the second half of 2005 and the comparison period. The key regressor is the change in the insurer's Part D enrollment between 2005 and 2006. Covariates include changes in the average number of pills per prescription, changes in the average per-pill wholesale drug price, and changes in each insurer's exposure to the pharmacy. The sample of drugs comprises the top 1000 drugs ranked by expenditures observed in the claims. The instrument for Part D enrollment is potential Part D enrollment. Changes in log prices predicted by the model are reported for a) the insurer at the 50th percentile of enrollment increases; and 2) for the market average, given the distribution of observed enrollment increases. Parentheses report standard errors clustered at the insurer level. \* significant at 10%; \*\* significant at 5%; \*\*\* significant at 1%

Appendix Table 4. Drug Prices and Plan Premiums

	Dependent Variable: ln (2006 Part D Plan Premium)			
	(1)	(2)	(3)	(4)
Plan Deductible	0.292 (0.296)	0.154 (0.326)	0.170 (0.329)	0.121 (0.456)
Low Income Subsidy Plan (Y/N)	-0.429** (0.174)	-0.407*** (0.132)	-0.411*** (0.132)	-0.426** (0.171)
Low Drug Coverage Level	-0.331 (0.236)	-0.257 (0.183)	-0.267 (0.191)	-0.267 (0.255)
Relative Drug Price Index (2005)		4.240 (3.028)	4.722 (3.587)	
Relative Drug Prices Index (2006)			-0.442 (1.436)	
Relative Change in Price Index (05-06)				1.667 (2.265)
Constant	6.239*** (0.232)	6.049*** (0.249)	6.052*** (0.251)	6.222*** (0.230)
Market Fixed Effects	Y	Y	Y	Y
Observations	332	332	332	332
R-squared	0.284	0.379	0.380	0.292

Appendix Table 4 reports validity tests for the quality-adjusted premium instrument. The dependent variable in each specification is the log of premiums for basic and actuarially equivalent Part D plans offered by insurers in the study sample. In each specification, log premiums are regressed on plan characteristics, controlling for market fixed-effects. In column (2), we control for 2005 insurer-level retail drug prices. These prices are contemporaneous to premiums given that 2006 Part D plan premiums were priced in mid year 2005. In column (3), we additionally control for 2006 retail drug prices. In column (4), we remove controls for drug price levels, and instead control for insurer-level index of price changes between 2005 and 2006. Parentheses report standard errors clustered at the insurer level. \* significant at 10%; \*\* significant at 5%; \*\*\* significant at 1%

Appendix Table 5. Non-Linear Enrollment Effect on Retail Drug Profits-per-Pill, by Branded/Generic Status

		Dependent Variable: $\Delta(\text{Drug Price-per-Pill})$					
Comparison Period		Second Half 2004			First Half 2006		
Drug Sample		All	Branded	Generics	All	Branded	Generics
		(1)	(2)	(3)	(4)	(5)	(6)
$\Delta$ Firm's PartD Enrollment (1M)		0.009 (0.036)	0.009 (0.049)	0.015 (0.052)	-0.228** (0.083)	-0.105 (0.084)	-0.354*** (0.099)
$\Delta$ Firm's PartD Enrollment <sup>2</sup> (1M)		-0.002 (0.007)	-0.001 (0.010)	-0.004 (0.011)	0.046** (0.017)	0.020 (0.017)	0.072*** (0.021)
NDC-level Drug Fixed Effects		Y	Y	Y	Y	Y	Y
Average Baseline Unit Price		3.85	6.51	0.61	4.07	7.15	0.72
Insurer Observations		33	33	33	33	33	33
Insurer-Drug Observations		9,939	5,438	4,501	10,265	5,611	4,654

Appendix Table 5 reports instrumental variables estimates of the effect of insurers' Part D enrollment on their unit profits earned in the commercial non-Part D market, by branded and generic drug status. The dependent variable is the change in the insurer-NDC drug level average price-per-pill (\$) between the second half of 2005 and the first half of 2006. The key regressor is the change in the insurer's Part D enrollment between 2005 and 2006. As derived in the text, including NDC-level drug fixed effects allows for the coefficients on enrollment to be interpreted as an enrollment elasticity of unit profits. Covariates include changes in the average number of pills per prescription and changes in each insurer's exposure to the pharmacy (the state-level market share of the pharmacy, weighted by the fraction of the insurer's total retail expenditures in each state). The sample of drugs comprises the top 1000 drugs ranked by expenditures observed in the claims. Instruments for Part D enrollment include potential Part D enrollment, and the quality-adjusted premium measure, which are described in the text. In all specifications, insurer-NDC-level observations are weighted by the number of claims for the NDC observed for that insurer. Parentheses report standard errors clustered at the insurer level. \* significant at 10%; \*\* significant at 5%; \*\*\* significant at 1%

Appendix Table 6. Non-Linear Enrollment Effect on Pharmacy Prices-per-Pill, by Age Group

Population Model Drug Sample	Dependent Variable: $\Delta \ln$ (Drug Price-per-Pill)					
	Commercially Insured Ages 60 and Over			Commercially Insured Ages 60-64		
	IV	IV	IV	IV	IV	IV
	All	Branded	Generics	All	Branded	Generics
	(1)	(2)	(3)	(4)	(5)	(6)
$\Delta$ Firm's PartD Enrollment (1M)	-0.262*** (0.053)	-0.027** (0.013)	-0.564*** (0.115)	-0.183*** (0.070)	-0.035*** (0.011)	-0.382*** (0.142)
$\Delta$ Firm's PartD Enrollment <sup>2</sup> (1M)	0.053*** (0.011)	0.005** (0.003)	0.115*** (0.024)	0.037** (0.015)	0.007*** (0.002)	0.077*** (0.029)
Predicted $\Delta \ln$ (Drug Profit/Pill): Median	-0.100	-0.010	-0.215	-0.070	-0.014	-0.146
Overall Predicted $\% \Delta$ in Expenditures	-0.083	-0.010	-0.172	-0.058	-0.011	-0.116
Insurer Observations	33	33	33	33	33	33
Insurer-Drug Observations	10301	5639	4662	6834	3756	3079

Appendix Table 6 reports instrumental variables estimates of the effect of insurers' Part D enrollment on their unit profits earned in the commercial non-Part D market, by age group. Price changes are measured as the difference in the log retail price-per-pill, averaged at the insurer-NDC drug level, between the second half of 2005 and the comparison period noted in column headings. The key regressor is the change in the insurer's Part D enrollment between 2005 and 2006. Covariates include changes in the average number of pills per prescription, changes in the average per-pill wholesale drug price, and changes in each insurer's exposure to the pharmacy (the state-level market share of the pharmacy, weighted by the fraction of the insurer's total retail expenditures in each state). The sample of drugs comprises the top 1000 drugs ranked by expenditures observed in the claims. Instruments for Part D enrollment include potential Part D enrollment, and the quality-adjusted premium measure, which are described in detail in the text. Changes in log prices predicted by the model are reported for a) the insurer at the 50th percentile of enrollment increases; and 2) for the market average, given the distribution of observed enrollment increases. In all specifications, insurer-NDC-level observations are weighted by the number of claims for the NDC observed for that insurer. Parentheses report standard errors clustered at the insurer level. \* significant at 10%; \*\* significant at 5%; \*\*\* significant at 1%