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Is Drug Coverage a Free Lunch? Cross-Price Elasticities and the Design of Prescription Drug Benefits

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Abstract

Recently, many US employers have adopted less generous prescription drug benefits. In addition, the U.S. began to offer prescription drug insurance to approximately 42 million Medicare beneficiaries in 2006. We use data on individual health insurance claims and benefit data from 1997-2003 to study the effects of changing consumers' co-payments for prescription drugs on the quantity demanded and expenditure on prescription drugs, inpatient care and outpatient care. We allow for effects both in the year of the co-payment change and in the year following the change. Our results show that increases in prescription drug prices reduce both the use of and spending on prescription drugs. However, consumers substitute the use of outpatient care and inpatient care for prescription drug use, and the expenditure reductions on prescription drugs are largely offset by the increases in outpatient spending.

Keywords: drugs, elasticity, substitution, cost-sharing, insurance

JEL Classification: D12, I10, M52

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1. Introduction

In the past 15 years, national spending on prescription drugs has grown dramatically, far outpacing the growth rate of hospital spending and physician spending during the same period.¹ In response to these rapid increases in drug spending, many health insurance plans have reduced the generosity of their prescription drug benefits. Consequently, patients have been paying substantially more out of their own pockets for prescription drugs in recent years. Ostensibly, benefit designers are seeking to reduce drug spending by increasing the price faced by consumers, the copayment.

A number of studies have investigated the relationship between cost-sharing and spending on drugs (Joyce et al., 2001; Goodman et al., 2004; Huskamp et al., 2003; Soumerai et al., 1987, 1991; Harris et al., 1990; Johnson et al., 1997; Tamblyn et al., 2001; Motheral and Fairman, 2001). Almost all of these studies suggest that higher cost-sharing reduces pharmaceutical use. In addition, some studies have found that higher drug cost-sharing results in worsened health status (Johnson et al., 1997), and more adverse health events such as emergency room visits, nursing home admissions, or hospital admissions (Soumerai et al., 1991; Tamblyn et al., 2001; Balkrishnan et al., 2001). These findings suggest that the reduction in drug spending may come with unintended consequences. Achieving reduced drug use may come at the cost of worse health consequently leading to use of expensive care such as the emergency room or inpatient hospital care.

Two economic theories provide potential explanations of why and how changes in the price of drugs affect people's health and the demand for other types of health care. First, the household

¹ Spending on prescription drugs in 2004 was nearly five times the level in 1990. Spending on physician and hospital services were a little over twice as high in 2004 as in 1990. Source: Center for Medicare and Medicaid Services, "National Health Expenditure Web Tables," <http://www.cms.hhs.gov/NationalHealthExpendData/downloads/tables.pdf>.

production model introduced by Becker (1965) provides a framework for analyzing the determinants of many behaviors including health-related behaviors. In this approach, prescription drugs are one of many inputs to a health production function. Drugs and other types of medical goods such as physician visits or hospital care are substitute (or complementary) treatments for some diseases. For example, clinical evidence suggests that either antidepressant medication or psychotherapy can be used as first line treatment for mild to moderate Major Depressive Disorder (MDD). Theory suggests that, to the extent that these other inputs are substitutes (complements) to drugs in producing health, increases in drug prices will result in increases (decreases) in the consumption of other medical goods.

The basic household production theory is static. Grossman's health capital theory (1972) is derived from Becker's basic model, and provides a dynamic framework for analyzing medical care demand. In this theory, individuals inherit an initial stock of health capital that depreciates over time and can be increased by investment. Gross investments in health capital are produced by inputs such as medical care, diet, exercise, etc. Prices of medical goods affect people's demands for medical care in each period. This model suggests that the effects of changes in input prices on consumers' demands for medical care have a dynamic component through their effects on health capital.

Taken together, these theories suggest that changes in drug prices will have effects not only on the demand for drugs, but also on the demands for substitute and complementary services. It also suggests that there will be a dynamic aspect to the effects of changes in drug prices. Adjustment will not occur instantaneously, but over time.

These theories are consistent with the observed outcomes of worsening health status and increased adverse health events following increases in consumer cost-sharing for prescription

drugs. A substantial fraction of drug consumption is directed at managing chronic conditions (hypertension, hyperlipidemia, diabetes, etc). In responding to changes in prescription drug prices, people with chronic illness reduce compliance with drug therapies. In the short-term there may not be serious adverse health events as a result. However, in the longer term poor compliance could lead to health stock transitions that manifest as poor health outcomes, which lead to additional medical use and expenses. For example, if an increase in drug co-payments causes people to fall out of compliance with their drug therapy for hypertension, they may be more likely to suffer heart attacks, strokes, and other complications, leading to hospitalization, physician care, additional medication, and higher health care spending.

The discussion above suggests that a fuller understanding of the effects of out-of-pocket drug price changes on health spending and health outcomes requires an examination of the dynamic structure of demand for health care services. This is our focus. We use a large panel dataset of health insurance claims and benefit design information to identify the effects on drug spending, outpatient spending, and inpatient spending of changes in workers' employer-provided prescription drug benefits.

There are two central findings. First, there is substantial substitution between prescription drugs use and the use of outpatient care. Increases in out-of-pocket drug prices lead to decreases in the demand for drugs, but lead to increases in demand and spending on outpatient care. We do not find detectable changes in inpatient spending as a result of increases in drug co-payments overall. However, for the small group of people who use inpatient services, the increase in drug co-payment leads to a substantial increase in inpatient spending.

Second, we find strong dynamic own-price effects for drugs and dynamic substitution effects for outpatient care. The dynamic price effects are substantially larger than the

contemporaneous effects. The effect of increasing drug co-payments on total health care spending is significantly smaller than is the direct effect on drug spending. This suggests a substantial offset effect: increased co-payments for pharmaceuticals result in savings on drugs, but that effect is substantially offset by increased spending incurred as patients substitute outpatient care for pharmaceuticals.

The paper is organized as follows. Section 2 provides relevant institutional facts and findings from prior literature. The data used in the study are described in Section 3. Section 4 describes the empirical strategy, including estimation methods. Results are discussed in Section 5. Finally, Section 6 contains a summary and conclusions.

2. Background

2.1 Background on Prescription Drug Insurance Benefit Design

Aggregate spending in the U.S. on outpatient prescription drugs has increased rapidly, both in absolute terms, and compared to the trends for spending on hospital services and physician services. Private drug spending has increased by 15-20 percent per year, starting in the 1990s.² Prescription drug spending is the third largest component (after hospital and physician services) of national health care expenses at \$162.4 billion. This amount is almost 5 times larger than the amount spent in 1990, and drugs now account for more than 11% of total health care spending.³

In response to these large and rapid increases in drug expenditures, many employers and insurance plans adopted more stringent prescription drug benefit designs, imposing greater cost-sharing on patients for the use of prescription drugs. By far the most common form of cost-

² Kaiser Family Foundation and Health Research and Educational Trust, Employer Health Benefits: 2004 Annual Survey, Sept 2004.

³ Source: Center for Medicare and Medicaid Services, "National Health Expenditure Web Tables," <http://www.cms.hhs.gov/NationalHealthExpendData/downloads/tables.pdf>.

sharing for prescription drugs is fixed co-payments. Drugs are typically divided into groups called “tiers,” and each drug in a given tier has the same co-payment. In a two-tier plan, consumers pay a lower co-payment for generic drugs and a higher co-payment for branded drugs. In a three tier plan, a further distinction is made between “preferred” and “non-preferred” branded drugs. There is a higher co-payment for non-preferred drugs. A typical co-payment schedule for a 3-tier plan is a \$5 co-payment for each generic drug prescription, \$10 for preferred branded drugs, and \$25 for non-preferred branded drugs. Moreover, starting in the late 1990s insurance plans further differentiated the co-payments for drugs purchased at walk-in pharmacies (“card plan” purchases) and at mail order pharmacies. The mail order part of the plan typically requires that ninety days supply of the drug be purchased at once. Co-payments are set so that mail-order prescription purchases cost less per day than do card plan prescription purchases.

These changes have led to an increase in out-of-pocket payments by consumers over time. According to statistics from the Kaiser Family Foundation⁴, the average co-payment for generic drugs increased from \$7 per prescription in 2000 to \$9 in 2003 (a 28.6% increase) for workers with employer-sponsored health plans. Co-payments for preferred branded drugs increased from \$13 to \$21 per prescription from 2000 to 2004 (a 61.5% increase), and for non-preferred drugs increased from \$17 to \$33 per prescription (a 94.1% increase) over the same period.

2.2 Cost-sharing and the Demand for Pharmaceuticals

A number of studies have evaluated the effects of increased prescription drug co-payments on the use of and spending on prescription drugs and on health status. The results from most of these studies are show that increased cost-sharing resulted in lower drug use and spending.

A prominent paper in this field of research is Joyce et al. (2002). Using individual claims data from 25 large employers with 75 distinct insurance plan-years from 1997-1999, the authors

⁴ Kaiser Family Foundation: Prescription Drug Trend, Oct 2004.

studied the effects of patient cost-sharing and formulary restrictions on spending for generic drugs, branded drugs and on patients' out-of-pocket spending. The study exploits cross-sectional variation in plan-level drug co-payments to identify the effects of cost-sharing and formulary restrictions. The results show that higher co-payments for prescription drugs are associated with significantly lower drug spending: the estimates imply that a doubling of the co-payment is associated with reductions in drug spending of 19%-33%.

Using a similar study framework and dataset, Goodman et al. (2004) further explore the effects of prescription drug benefits on the use of the eight most commonly used therapeutic classes of drugs. The key independent variable in this study is the generosity of prescription drug benefits, which is calculated as the insurance plan-level price index for a standardized "market basket" of drugs. The estimation results in this study imply that doubling the co-payment for each therapeutic class is associated with lower use of almost all therapeutic classes of prescription drugs, although patients' demands might be more responsive for "nonessential" drugs such as NSAIDs and antihistamines than to "essential" drugs such as antihypertensive drugs and antidepressants.

Other studies on this topic include Huskamp et al. (2003), Soumerai et al. (1987, 1991), Harris et al. (1990), Johnson et al. (1997), Tamblyn et al. (2001), and Motheral and Fairman (2001). Johnson et al. (1997) estimates the effects of increased cost-sharing by exploiting the difference in the drug co-payment changes in two large Medicare HMO plans. A difference-in-difference estimation strategy is used. Soumerai et al. (1987) uses time-series analysis to study the effect of imposing caps to the number of drug prescriptions in one state's Medicaid program. They find that among the 10,734 continuously enrolled patients, the cap caused a sudden and sustained drop of 30% in the number of prescription filled. Tamblyn et al. (2001) investigates the

effects of the introduction of prescription drug cost-sharing policy in Quebec Province, Canada in 1996. Using an interrupted time-series study design, they study the effects of instituting a cost-sharing schedule of a deductible plus 25% coinsurance on the use of essential and nonessential drugs for poor and elderly population. They find significant reductions in the use of essential drugs (9%-14%) and in less essential drugs (15%-22%) for this population after the introduction of this drug cost-sharing schedule.

In sum, previous studies use various designs, measurements, statistical models and data samples and come to a consensus that increased cost-sharing is associated with lower drug use and spending.

2.3 Substitution and Dynamic Price Effects

Our study focuses on two important economic questions: what are the substitution and dynamic price effects of pharmaceutical co-payments on the demand for medical care?

The notion of substitution among different types of medical care has long been recognized in the health economics literature. However, most of the existing studies on this topic focus on the substitution between outpatient and inpatient care. For example, Davis and Russell (1972) use data for 48 states on outpatient visits and inpatient admissions in nongovernmental and nonprofit hospitals to study the substitution between all outpatient and inpatient services and find evidence of substitution between these two types of care. Helms et al. (1978) study the effects of imposing small co-payments for out-of-hospital services on some Medicaid beneficiaries in California, and find the co-payment requirement decreased physician visit demand by 8 percent, but increased hospital service demand by 17 percent, suggesting the substitution of inpatient care for outpatient services when the price of the latter changed.

Evidence from studies on drug insurance policy changes suggests the existence of substitution effects. For example, Balkrishnan et al. (2001) study the effects of increased cost-sharing on prescription drugs for 2,411 Medicare HMO enrollees in 1998. They find this resulted in a 25.2% increase in annual inpatient admissions in the first year. Soumerai et al. (1991) study the effects of imposing caps on the number of prescriptions for a small sample of elderly people with chronic illness in New Hampshire and find a significant increase in nursing home admissions after the drug use limit was introduced. Tamblyn et al. (2001) find that a reduction in essential drug use (induced by higher co-payments) was associated with significantly higher adverse events such as hospitalization and long-term care admission.

Lichtenberg (1996, 2001) looks at the effect of changes in the quantity and type of drugs prescribed by physicians on the changes in the use of other medical inputs such as hospitalization and surgical procedures. He finds that the number of hospital stays, bed-days and surgical procedures declined most rapidly for those diagnoses with the greatest increase in the total number of drugs prescribed. In Lichtenberg's work the changes in drug use were driven mostly by physicians' prescribing behavior and the diffusion of new prescription drugs; therefore, his empirical results can be best interpreted as the "biological" or "technological" substitution between drugs and other medical inputs in health production. Duggan (2004) investigates whether newer and more expensive antipsychotic drugs offset, and therefore substitute for, the use of other health care services such as inpatient care. He finds no evidence of an offsetting effect for anti-psychotic drugs.

We study consumers' demand for medical goods and the substitutability of drugs for other medical inputs. Using consumer demand theory as our conceptual framework, we use variation in out-of-pocket prices facing patients for the use of prescription drugs and other medical goods

to identify substitutability between drugs and other medical services. Since Grossman's health capital theory suggests the existence of dynamic price effects, we estimate a demand system with a lagged price structure. This demand system represents a reduced form of a full dynamic model of medical care demand and health production. The contemporaneous and lagged prices on the demand equations allow us to estimate short and long term effects of insurance price changes. To our knowledge, this is the first study on this topic to allow for dynamic price effects.

3. Data

We use data from the Medstat MarketScan database. MarketScan is the largest private sector health care database in the U.S, containing paid claims of more than 7 million privately insured individuals, and over \$13 billion in annual healthcare expenditures. Medstat had contracts with over 40 large employers for the submission of the health insurance data for their employees over the period 1990-2003. Neither employers nor health plans are identified by name in the database. The database contains longitudinal data for each person, including person and family identifiers, enrollment history, uses of inpatient care, outpatient care and prescription drugs, health expenditure, and detailed health insurance coverage information from 1990 to 2003.

We link information from five different files in the Medstat database from 1997-2003: 1) the enrollment file, containing patients' demographics and detailed information on their health plan enrollment history, 2) the employer benefit plan design file, containing summary benefit descriptions for major medical and prescription drugs benefits for many health plans, 3) the hospital inpatient claims file, containing individual hospital claims aggregated to the level of the hospital stay and providing information on diagnosis, treatment, length of stay, and basic payment information, 4) the outpatient service claims file, containing individual outpatient claims aggregated to the level of each outpatient visit with information on diagnosis, treatment

procedures and payment information, and 5) the outpatient pharmaceutical claims file, containing a claim for each prescription filled by each person with information on days of prescription drug supplied, national drug codes, therapeutic classes and payment information.

3.1 Sample Selection

More than 40 individual employers contributed data to the MarketScan databases over time. However, not every employer submitted all five files to Medstat in a given year. We use only firms with complete information in all of the above five files. Moreover, in our empirical work, we estimate models with dynamic price effects (1 lag), and person-specific fixed effects. This requires at least three consecutive years of full information. Therefore we selected from the overall database firms which had complete information from all of the five files mentioned above for at least three consecutive years. This removed a large number of firms from our analysis, principally because a large number of firms did not submit prescription drug data. For example, in 1997 only 19 out of 53 employers (an employer may have one or multiple health plans in a specific year) submitted prescription drug claims data; in 2000, only 24 out of 45 employers submitted drug data; and in 2003, 38 out of 45 employers submitted drug data. After applying the 3-consecutive-year requirement, 16 employers remain.

Of these 16 employers, two used coinsurance in their prescription drug benefits design and the other 14 used co-payments as their cost-sharing mechanism. Since these two cost-sharing mechanisms represent different incentive strategies in insurance benefit designs, and have different effects on the demand for medical care, we limited our analysis to the fixed co-payment insurance plans, which represent the most common benefit design for pharmaceuticals (Kaiser 2006). Last, the employer benefit plan design file contains some missing or inaccurate information for prescription drug benefits or medical benefits for some employers. We delete the

firms with unclear or missing insurance benefits information from our analysis sample, an additional three firms.

There are eleven employers which have three consecutive years of full information, clear insurance benefit information, and use co-payments as their cost-sharing mechanism for prescription drugs. These employers offer multiple insurance plans of varying generosity. Of these eleven firms, nine had a single, uniform prescription drug benefit plan: that is, all employees faced identical prescription drug plans at any given time. We focus only on the plans at these nine employers. For consumers covered by these firms, any change in consumer out-of-pocket price for prescription drugs comes about not from employees switching drug plans to change their own out of pocket benefits, but from employers uniformly changing the benefits of all of their employees. This selection strategy left us with 97 insurance plans from the nine large employers.⁵

On average there are 4.3 years of data from each firm. There are in total 1,304,687 individuals who have ever enrolled at least one full year in these nine firms. We further restrict our selection of individuals by examining only those individuals who have been continuously enrolled for at least three years during the 1997-2003 period.⁶ This selection criterion rules out about 56% of individuals. Last, people older than 65 are excluded because of the complexities introduced by Medicare coverage, potential outside Medigap coverage, and coordination of benefits issues. Finally, we have a panel data set of 1,713,879 person- years for 526,086 people in 97 different insurance plans at nine different employers, spanning a seven-year period from 1997 to 2003.

⁵ Including these two firms which have non-identical prescription drug benefits in a given year and are deleted from our study sample doesn't change the main results of this paper.

⁶ Continuously enrolled here means continuously enrolled in any of the firm's health plans. We are not dropping people who switch among health plans.

Given all of the observations eliminated by the inclusion criteria, there is a natural concern about the representativeness of the data in the analysis sample. We therefore make a number of comparisons of the analysis sample with the full sample from the Medstat database and the population in the U.S. with employer sponsored health insurance.

In *Figures 1 through 4*, we compare demographic and spending variables for people in our study sample with the full sample in the Medstat database and with the employer-insured U.S. population. We get the information on age, regional location of residency, annual total medical spending and annual pharmaceutical spending for the employer-insured US population from the Medical Expenditure Panel Survey from 1997 - 2003. *Figure 1* suggests that the age distribution in the Medstat sample (excluding those age greater than 65) represents the US population quite well, except that it underrepresents the age group 25-44 by 6.8%, and overrepresents the age group 55-64 by 4%. The same conclusion holds for the comparison of our study sample with the US population – our sample is slightly older. The Medstat sample and our study sample are less representative in terms of regional distribution compared with the US population. The South region is over-represented in these two samples, and Northeast and West are under-represented. This is described in *Figure 2*.

Figures 3 and 4 indicate that the Medstat sample and the MEPS sample for the US population follow similar trends in per-capita total spending and pharmaceutical spending, but enrollees in the Medstat sample on average spend \$146 more for prescription drugs, and \$300 more for all types of medical care. For our study sample, enrollees on average spend \$380 more for prescription drugs, and \$945 more for all types of medical care, compared to an average employer-insured person.

Overall, while there are some differences, it does not appear that the analysis sample we use differs critically from either the Medstat or the U.S. populations.

3.2 Measures

Our dependent variables are quantity and total spending on prescription drugs, outpatient care and inpatient care. For prescription drugs, quantity is the sum of days supplied from all prescriptions filled from a particular year for a patient.⁷ Similarly, quantity of outpatient services is the total number of outpatient visits, and the quantity of inpatient services is the total number of inpatient admissions.

Total spending on prescription drugs, outpatient services and inpatient services are calculated as the yearly spending per enrollee. This measure of spending is the sum of spending by the insurer in the database and the required out-of-pocket spending by the insured person. We are not able to observe whether or not the consumer actually made their out-of-pocket payment, neither are we able to observe the operations of coordination of benefits.

Since we use an individual fixed effects model in our estimation, only time-varying socio-demographic variables are used. The effects of time-constant variables such as race, sex, education, etc. are absorbed in the individual fixed effect. In our estimation, we use indicators of urban residence, retirement status, a set of individual fixed effects, and a set of year fixed effects. Because we enter both individual and year fixed effects, we cannot also enter age into our estimating equations. We do wish, however, to allow spending to grow at different rates for people in different age groups; therefore, we construct a set of interactions between dummies for age category and a linear time trend. We separate people into seven age categories: 0-10 years, 11-18, 19-29, 30-39, 40-49, 50-59, and 60-64.

⁷ Our results are similar if we define quantity as the number of prescription fills.

Our primary independent variable of interest is the out-of-pocket price faced by consumers for prescription drugs. In the presence of health insurance, the prices faced by a consumer for health services are determined by the consumer's health plan benefit design. Ideally, one would like to include all the relevant aspects of the prescription drug benefit design in the analysis. These are six variables describing the benefit design in the Medstat database (generic co-payment, preferred brand co-payment, and non-preferred brand co-payment, separately for card and mail order). While this is a rich source of descriptive information, these measures are highly collinear, therefore it isn't possible to separately identify their effects in a regression. Further, there are 18 changes in drug benefit design that occur in the data, making it highly unlikely we could identify parameters for six variables describing the drug benefit (see *Table 2*).

Thus, we construct an out-of-pocket price index for prescription drugs for each health plan in each year. For each plan-year, the price index is a weighted average of the out-of-pocket co-payments for that plan's tiers: the generic co-payment, the preferred brand co-payment, and the non-preferred brand co-payment. For a plan with only two tiers, we use that plan's brand co-payment as both the preferred and non-preferred brand co-payment. In addition, since plans often specify different co-payments for the card plan and the mail order plan, we differentiate between those two modes of delivery in the price index. The formula for the price index for plan j in time t is:

$$\begin{aligned}
P_j^t = & Copay_{j,G}^{t,Card} * W_{Card,G} + Copay_{j,PB}^{t,Card} * W_{Card,PB} + Copay_{j,NPB}^{t,Card} * W_{Card,NPB} \\
& + Copay_{j,G}^{t,Mail} * W_{Mail,G} + Copay_{j,PB}^{t,Mail} * W_{Mail,PB} + Copay_{j,NPB}^{t,Mail} * W_{Mail,NPB}
\end{aligned} \tag{1}$$

$Copay_{j,G}^{t,Card}$, $Copay_{j,PB}^{t,Card}$ and $Copay_{j,NPB}^{t,Card}$ are plan j 's co-payment for one prescription of generic drugs, preferred brand drugs and non-preferred brand drugs, respectively, purchased

using insurance card at walk-up pharmacies. $Copay_{j,G}^{t,Mail}$, $Copay_{j,PB}^{t,Mail}$ and $Copay_{j,NPB}^{t,Mail}$ are plan j 's co-payment for one prescription of generic drugs, preferred brand drugs and non-preferred brand drugs respectively purchased by mail order. The $W_{m,n}$ s are quantity-based weights for generic, preferred brand and non-preferred brand drugs for card and mail-order purchases, calculated using prescription drug claims data for all the enrollees in all the years in the nine study firms. For example, the weight on generic, mail-order, $W_{Mail,G}$, is the proportion of prescriptions in our whole sample which are generic drugs ordered by mail-order.

Table 1 contains the weights for the six categories. For example, the table shows that $W_{Mail,G}$, the weight on generic mail-order co-payment, is 0.0603. Since each employer in our data offers a uniform prescription drug benefit at any given time, this price index changes for a consumer only when his employer changes its prescription drug benefit design.

Because of the potential for input substitution in the production of health, the demand equations for each type of medical care are functions of drug prices and the prices of other medical services. Insurance benefits for outpatient and inpatient services are more complicated than those for prescription drugs. Common cost-sharing devices for medical services take the form of a combination of deductibles, coinsurance rates for spending above the deductible, a co-payment for one physician office visit, and a stop-loss limit beyond which consumers don't pay any more. The budget sets for these medical services are therefore complicated and nonlinear. It is difficult to construct a single price measure for these services that would correctly reflect the true out-of-pocket prices consumers pay. Instead we include the deductible for medical services and the co-payment for outpatient visits as our out-of-pocket price measures for medical services.

Table 2 provides a description of the history of prescription drug benefits for the nine firms, and the drug price index for the health plans within each employer. Each firm changed its

prescription drug benefits at least once, with a total of 18 price changes occurring during the 1997-2003 period. As is clear in *Table 2*, each benefit change is associated with increases in co-payments and therefore made the prescription drug insurance less generous. This is consistent with the national trend of increasing drug co-payments over the past decade. In *Figure 5* we compare the average co-payments for generic, preferred branded and non-preferred branded drugs of the nine firms in our study with those of the employer-sponsored prescription drug insurance plans for the U.S. population. The estimates of drug co-payments for U.S. population are from the Employer Health Benefits Annual Survey by Health Research & Educational Trust (HRET) in 2000-2003⁸.

Table 3 gives definitions of all the variables used in the demand estimation, and *Table 4* provides summary statistics of these variables. Our dependent variables are the spending and use variables for prescription drugs, outpatient care and inpatient care. We construct the total spending variable as the sum of spending on these three types of medical care. In *Table 4* we also report the number of observed zeroes in the corresponding quantity and spending variables. In our sample, 96% of the inpatient observations have zero values, compared to 30% for prescription drug use, and 17% for outpatient care. In this sample 87% of individuals never used any inpatient care during the time span when data are available for them, compared with 13% for prescription drugs and 6% for outpatient care. We describe our empirical strategy for coping with the large number of zero values in the next section.

By construction, there is no switching within employer among drug plans, since all of our employers have uniform drug plans over time. But switching among medical plans is still of some concern. Employers have, on average in our data, 2.76 medical plans available to their

⁸ The Employer Health Benefit Annual Survey is funded by Kaiser Family Foundation and studied jointly by HERT and Kaiser Family Foundation since 1987. It collects health insurance benefits information of approximately 2,000 randomly selected employers in all major industries.

employees. **Table 5** gives summaries insurance plan switching rates for each firm. This table shows that the switching rates are fairly small. One might interpret the low switching rates in **Table 5** as indicating that, although people may self-select into insurance plans, it seems that these selections are determined mostly by stable health or preference factors at the baseline year. If this is true the individual fixed effects will largely account for unobservable factors which drive selection.

4. Empirical Strategy

4.1 Model Specification

We estimate equations relating the demand for and the total spending on prescription drugs, outpatient care, and inpatient care to the out-of-pocket prices paid by consumers for drugs and to the medical plan design characteristics. The basic estimation model is

$$Q_{it}^j = \mathbf{b}_1 + \mathbf{b}_2 P_{it}^d + \mathbf{b}_3 P_{it-1}^d + \mathbf{b}_4 P_{it}^m + \mathbf{b}_5 P_{it-1}^m + X_{it} \mathbf{d} + \mathbf{a}_i + \mathbf{g} + \mathbf{e}_{it} \quad (2)$$

In this equation Q_{it}^j denotes the demand for health input j by person i in period t , where j indexes prescription drugs, outpatient care, and inpatient care. $P_{it}^d, P_{i,t-1}^d$ are indexes of the contemporaneous and lagged patient out-of-pocket prices for prescription drugs $P_{it}^m, P_{i,t-1}^m$ stand for prices for other types of medical care, such as outpatient and inpatient care. The X_{it} capture all non-price time-varying variables that also affect peoples' demand for medical care, such as retirement status, urban-rural location, and age group-year interactions. The error term can be decomposed into three separate elements. \mathbf{a}_i captures unobservable and unchanging individual heterogeneities in medical demand, such as individuals' preferences for using medical goods,

inherited generic traits, etc. g_t captures the general trend effects in demand over time, and e_{it} stands for all other random factors which might affect demand, such as random health shocks.

Since we are concerned with dynamic effects of out-of-pocket drug prices on both drug demand and demand for other medical services, throughout our analysis we include both the current drug price index and one lag of the drug price index. We include a lag of prices in our demand model for two reasons. First, there is substantial evidence that people treat health as an investment, as suggested by Grossman's health capital theory (1972). Thus, a one-time change in drug prices will not only alter consumption in the current period, but also in future periods. Including a lag in drug prices is a simple way to allow for this effect. Second, drug consumption may be sticky. Previous literature suggests that there might exist substantial persistence in the habits of patients' use and doctors' prescribing of prescription drugs (Coulson and Stuart, 1992; Hellerstein, 1998; Coscelli, 2000). As the result patients and their doctors may not alter prescribing and filling behavior quickly in response to a change in drug prices.

In this model, the "long-run" effect of a change in out-of-pocket prescription drug prices is the sum of the effects of both the contemporaneous and lagged price variables. We are unable to include a lag structure longer than 1 year, since this would require including only firms with four or more years of usable data, and would thus reduce our sample size greatly⁹.

The spending measure we use throughout is the total spending on the relevant service. For example, total spending on drugs is the consumer's out-of-pocket spending plus the insurance plan's spending, subject to the caveats above. It is important to distinguish our regression of total spending on price with the more familiar expenditure function approach to estimating demand. Our total spending regressions are not expenditure function regressions, because the left-hand-

⁹ Adding two lags in the regression will cause the losing of three firms which have only three years of data from our estimation sample. Total number of individuals would be 355920, almost a third reduction from 526086 in our sample with only one lags.

side spending variable is not the *consumer's* total spending but the consumer's total spending plus the plan's total spending. Therefore, our price elasticity of spending is *not* the consumer's price elasticity of demand plus one, as it is in consumer theory. The spending variable here is more like a quantity index: it weights each prescription by the *total* (as opposed to out-of-pocket) price paid for the prescription.

4.2 Econometric Issues

Throughout the analysis, we use fixed effects to control for unobserved individual characteristics. This choice is motivated by several considerations. The consumer-specific information we have access to is quite sparse. For example, we have no income information and the only health status information we have is that which we can infer from the claims data.¹⁰ Second, though there is little switching among medical plans over time, there is still the prospect of adverse selection: consumers may already have selected into their health plans at baseline. Thus, we hope to control for unobserved consumer characteristics and to mitigate adverse selection by including individual fixed effects. Since firm fixed effects are contained in the span of the individual fixed effects, this strategy also means that we are identifying the effects of the out-of-pocket drug prices solely from the variation induced by changes in the drug plans within a firm over time.

We use linear fixed effect models for both the quantity demand and expenditure equations. This approach is straightforward; however, there are a number of econometric issues associated with the use of linear fixed effect estimators, given the nature of some of our data. The measures of quantity are counts. Some of the expenditure variables have significant probability mass at

¹⁰ We could attempt to estimate health status from the claims data by using a “before” period. However, that would require shortening our panel, which we are loathe to do.

zero. We therefore also employ fixed effect count data model and fixed effect Tobit model to estimate the quantity and spending equations, respectively.

Another issue is serial correlation. The errors in the demand equations are likely to be serially correlated when the individual fixed effects and time trends do not completely control for the persistence in consumer's demand. We know of no straightforward parametric method for correcting for serial correlation in the non-linear settings discussed above. We therefore use the block bootstrap method to adjust standard errors for serial correlation.

Last, a large number of zeroes are observed in both quantity and spending for inpatient care (most people are not hospitalized in a given year). As will be explained in more detail in the following section, in estimating the individual fixed effect Poisson model, individuals for whom the dependent variable is always zero make no contribution to the likelihood function. Similarly, in the fixed effect Tobit estimation, the symmetric trimming procedure employed by this estimator causes substantial loss of observations for inpatient demand equation. The parameter estimates from fixed effect Poisson and fixed effect Tobit estimators are not biased because of the data loss. However, it is likely that these estimates for the inpatient demand equation reflect demand responses for a smaller group of individuals who are inpatient care users, and therefore may not be representative of the entire population. We therefore aggregate the data into broader groups and re-estimate the inpatient demand using fixed effects for these groups. The results from these regressions provide an alternative look at inpatient demand. Indeed, we do find significant differences between the parameter estimates from individual and group fixed effect models.

4.2.1 Fixed Effect Count Data Model

As *Table 3* documents, our quantity variables are counts and some of them are frequently zero.

We therefore use fixed effects count data methods to model the m.

Specifically, we use Hausman, Hall and Griliches' (1984) conditional maximum likelihood estimator (HHG). The HHG estimator is based on a conditional mean assumption,

$I_{it} = \exp(\mathbf{a}_i + x_{it}\mathbf{b}) = \exp(\mathbf{a}_i)\exp(x_{it}\mathbf{b})$, where the fixed effect takes the multiplicative form $\exp(c_i)$. Estimation is fairly straightforward, since the individual fixed effects parameters c_i are conditioned out in the log-likelihood function. This estimator also has strong robustness properties. Consistency requires only that the conditional mean assumption is correct. Except for the conditional mean, the distribution of y_{it} given (\mathbf{a}_i, x_{it}) is completely unrestricted. That means the estimates are consistent, even in the presence of overdispersion, underdispersion, or serial correlation (Wooldridge, 1999).

4.2.2 Fixed Effect Tobit Model

It would be inappropriate to use linear regression models for the spending variables, given the large number of zeros observed for each. Obviously, these zeroes arise when consumers do not use any of the services in the relevant category in a year. We therefore use fixed effect Tobit models to estimate the parameters of the spending models.

Although there has been a large literature on identification and estimation of linear panel data models with fixed effects, fixed effect limited dependent models have not been studied as much. Honoré (1992) proposed a semi-parametric estimator for estimating the fixed effect Tobit model. The idea for this estimator is to restore the symmetry of the distribution of the dependent variable which was destroyed by censoring. In a panel data context, the censored regression model can be described by

$$\begin{aligned}
y_{it}^* &= x_{it} \mathbf{b} + \mathbf{a}_i + \mathbf{e}_{it} \\
y_{it} &= \max\{0, y_{it}^*\}
\end{aligned}
\tag{3}$$

Honoré (1992) proposed that under the assumption of independent and identically distributed errors $(\mathbf{e}_{it}, \mathbf{e}_{is})$, where t, s denotes errors in different periods, the conditional distribution of (y_{it}^*, y_{it}^*) given (x_i, \mathbf{a}_i) is distributed symmetrically around the 45° -degree line through $(x_{is} \mathbf{b}, x_{it} \mathbf{b})$ (or equivalently through the 45° -degree line through $((x_{is} - x_{it}) \mathbf{b}, 0)$). As discussed in Honoré's (2000) paper, this i.i.d. error assumption can be relaxed, and the estimator is still applicable under only conditional exchangeability and conditional stationarity assumptions. As this is true for any value of \mathbf{a}_i , it is also true for the conditional distribution of (y_{it}^*, y_{it}^*) given x_i .

Honoré proposed two estimators - the trimmed least absolute deviation estimator (LAD) and trimmed least square estimator (LS) - for fixed effect Tobit models. These two estimators are essentially the generalization of Powell's (1984) least absolute deviation estimator and Powell's (1986) trimmed Least Square estimators for censored regressions, in the setting of panel data with fixed effects. Each estimator is consistent and asymptotically normal under fairly weak assumptions. Moreover, these estimators do not impose any parametric structure on the distribution of error terms.

As later pointed out by Deaton (1997), Powell's (1984, 1986) censored LAD estimator and censored least square estimator can be calculated easily by repeated application of linear least squares or least absolute deviations regression algorithms. This point was further reinforced by Chay and Powell (2001). Simply put, these estimators can be achieved by iterating between the "symmetric trimming" of the dependent variables using estimates from the previous iteration and least squares or median regression using the "trimmed" data.

Honoré's least square and LAD estimators are replications of Powell's methods in a panel data framework. The slight difference between them resides in how the data are properly trimmed. To derive our fixed effect Tobit estimates, we use the identically censored least squares (ICLS) estimator of Honoré as described by Clay and Powell (2001).

4.2.3 Serial Correlation in Panel Errors

Estimating health care demand using longitudinal data is subject to a potential serial correlation problem. Several studies show that consumers' health spending is persistent from one year to the next (Eichner, McClellan and Wise, 1996; Vliet, 1992; Pauly and Zeng, 2003). Bertrand, Duflo and Mullainathan (2002) evaluate several methods of estimating consistent standard errors in the presence of serial correlation in difference-in-difference models. They find that the block bootstrap method works well in producing precise standard errors in the case when T is small and N is large.

We use the block bootstrap method to correct our standard errors for serial correlation. The block bootstrapping is done at the individual consumer level, so that we are correcting for serial correlation and heteroskedasticity at this individual level. Since we have to drop one year of data because of the lagged prices, our panel data has small T (3.3 on average per person) and large N (526,086), in which setting Bertrand, Duflo and Mullainathan (2002) found the bootstrap to be a good choice. The bootstrap has the merit of avoiding strong parametric assumptions about the structure of the error variance matrix. The serial correlation problem and possible solutions to it for nonlinear panel models, such as fixed effect Poisson models and fixed effect Tobit models which we will estimate in this study, are still not fully explored in theory.

4.2.4 Grouped Fixed Effects

Ideally, we would like to use individual level fixed effect models to estimate all the demand and spending equations. However, for inpatient admissions, 87.35% people in our sample have zero admissions in each of the years we observe them. Mechanically, these observations would be dropped in the HHG estimator for inpatient demand. Similarly, the symmetric trimming requirement implied by the ICLS estimators for fixed effect Tobit model also suggests a substantial loss of individual data for inpatient spending equation. Although there is no bias associated with this if the assumptions of the underlying models are true, the sample thus used for estimation is highly un-representative, as it is older and sicker.

One approach is to use fixed effects for groups of individuals, where there are some non-zero values in the dependent variables in each group. However, this remedy also comes with a price- the broader fixed effect parameters can't completely absorb all individual heterogeneity. In addition to the individual fixed effects specifications, for inpatient care, we also estimate grouped fixed effects.

The groups are created by dividing individuals into age group-sex-firm categories. There are 126 ($=7$ age groups \times 2 sexes \times 9 firms) groups. For the inpatient equation, our estimates for inpatient use involve a regression of admissions on these group dummies, time dummies, time-age group interactions, pharmaceutical price index, medical benefit design variables, and the dummy variables for each of these groups into a more conventional Poisson regression. For the inpatient spending equation, we run a conventional Tobit regression including the same independent variables.

5. Results

Figure 6 shows the time series movements of per-capita prescription drug spending and average out-of-pocket prices for prescription drugs over the study period, 1997-2003. These drug prices are the weighted average of yearly pharmaceutical price indexes at the firm level. Over this period, both the average co-payment and the average spending per enrollee have increased. Furthermore, the correlation over time between detrended average co-payment and detrended average spending is 0.7435 (N=7) and the OLS regression coefficient from a regression of spending on co-payment and a time trend is 14.16 (also N=7).

Table 6 summarizes the fixed effect estimation results for the spending equations, and **Table 7** summarizes the fixed effect results for the quantity demand equations. In each table we present results from individual linear FE models, individual FE Tobit and FE Poisson models, and the grouped fixed effect Tobit and Poisson models. In each table, the point estimates of marginal effects are reported. The parameter significance indicators are based on block bootstrapped standard errors which cluster residual errors at either individual or group dummy level. For each demand variable only the coefficients of contemporaneous and lagged drug prices are reported. Coefficients for other control variables are available from the authors upon request.

The coefficients for prescription drug demand through all specifications yield consistent results. Increases in out-of-pocket drug prices cause fewer days of drug use and reduce spending on drugs. For example, the individual FE Tobit estimates suggest a \$1 increase in drug price reduces total drug spending by \$33.50 in the first year after the price change, and a further reduction of \$46 in the second year after the price change. This corresponds to a “short-run” elasticity of -0.5 and “long-run” elasticity of -0.8 for drug spending. The regression results for days of drug supply follow a similar pattern. These results concur with previous findings that

more stringent drug cost-sharing benefits are associated with reductions in drug use and drug spending. One distinguishing contribution of our analysis is its examination of the dynamic effects on consumers' demand. The results from our study show that there are significant and strong lagged price effects on the demand for drugs. Moreover, the lagged price effects exceed the instantaneous effects. These results suggest that there are substantial adjustments to drug consumption in the long term, along with considerable shorter term stickiness.

Through all specifications, results from both the OLS and Tobit estimations suggest that consumers facing higher drug co-payments substitute to outpatient services in both the short and long run. Specifically, from the individual fixed effect Tobit results in *Table 6*, a one dollar increase in the out-of-pocket drug price index increases per-capita outpatient spending by \$7.70 and in the first year and \$19 in the second year after price changes. These estimates correspond to "short-run" and "long-run" cross-price elasticities of 0.07 and 0.18, respectively. Estimates from the outpatient quantity demand equation are less consistent. The fixed effect Poisson results in *Table 7* indicate that a one dollar increase in the out-of-pocket drug price index leads to 0.026 fewer outpatient visits in the first year, but 0.075 more outpatient visits in the second year after the price change. This corresponds to cross-price elasticities of -0.02 and 0.07 with drugs, respectively. One limitation associated with the measure of the number of outpatient visits variable is that outpatient visits are not homogeneous. An outpatient surgery counts the same as a physician office visit. It is possible that in the short-run consumers reduce the use of less expensive outpatient visits, such as a trip to the doctor's office to get a prescription written, but seek more expensive alternative treatments in the outpatient setting. In the long-run, both the quantity and spending on outpatient care go up as the result of increases in drug prices. This sustained substitution into the use outpatient care in a longer period might signal the worsening

of consumers' underlying health status. That is, the reduction in prescription drug use in both periods affects patients' health, and in turn generates more demand for outpatient care in the long-run.

The individual linear FE results for inpatient demand indicate that there is no significant relationship between the use of inpatient admissions and prescription drug prices. However, the individual FE Tobit results suggest a different story. There are large positive price effects on inpatient spending in both the first year and second year after the price change. A \$1 increase in drug price causes a \$183 increase in inpatient spending in the first year, and another \$800 increase in the second year. Note that in the FE Tobit estimation for inpatient spending, nearly 90% of observations are dropped because of the symmetric censoring. Therefore, the FE Tobit results are essentially the regression coefficients for the 10 percent of individuals who had positive spending on inpatient care. Taking into account that fact that this group of people spend on average \$15,000 annually on inpatient care, the resulting elasticities for the Tobit estimates are 0.19 and 0.82 for the short and long run, respectively. These estimates suggest that sick people are sensitive to changes in drug prices. When drug price increases, these people end up spending more on hospital care. Presumably, the substitution into hospital care is driven mostly by a deterioration in health status.

Note that this strong inpatient substitution effect goes away when we look at the grouped FE Tobit results, where every observation is included in the estimation. The coefficients for both short-run and long-run drug prices change signs in this group FE Tobit model. As discussed previously, these results might be of some concern because the individual heterogeneity might not be completely controlled for by the group dummy variables. Nevertheless, the large discrepancies in parameter estimates from the individual Tobit model and group Tobit model

suggest that the demand responses for sick people differ substantially from those of an average person.

A comparison of the results from linear fixed effect models with those from the FE Poisson and FE Tobit models indicates that the linear FE models produce estimates which may be systematically biased toward zero. These results suggest that using simple linear regression models for prescription drug use could result in understating consumers' demand responses.

6. Summary and Conclusions

We have estimated consumers' responses to increased cost sharing for pharmaceuticals accounting for the contemporaneous and lagged responses of drug, outpatient, and inpatient quantity and spending to increases in drug co-payments. Our results show that increased consumer cost sharing for prescription drugs reduces both use and spending on prescription drugs. We also find dynamic adjustment by consumers: the effects one year after a co-payment increase are substantially larger than the contemporaneous effects. We also find that consumers substitute to outpatient care in response to rising drug prices. These effects also have a significant dynamic component: there is substantially more substitution to outpatient care one year after an increase in pharmaceutical cost sharing. There is no significant substitution between drugs and inpatient care for an average person. However, we find large substitution effects into inpatient care for the small group of people who are users of inpatient care.

These results are intuitively plausible. We expect there to be own price effects. Theory and intuition also tell us that consumers will not likely adjust instantaneously to changes in their out of pocket costs. It is also plausible that consumers substitute outpatient care for medications. This may happen directly, as consumers pursue other treatments for their conditions. It may also

occur because consumers experience more adverse health events as a result of decreased drug consumption, thereby leading to more use of outpatient care.

In total, we find that the expenditure savings on prescription drugs are largely offset by increases in outpatient spending. A \$1 increase in drug price reduces drug spending by \$33.50 in the first year, and \$46 in the second year. This amounts to a total reduction of spending by almost \$80 in a two year period. However, total medical spending decreases by about \$48 (\$24+\$24) in this two year period. Thus, higher drug co-payments save money on drug spending, but cost money on outpatient and possibly inpatient spending and have much smaller effects on overall spending.

These findings shed light on the efficiency of current insurance benefit designs in both the private and public sectors. The trend toward increased consumer cost sharing for prescription drugs should be carefully examined in light of these findings. Our results may also have relevance for the new Medicare prescription drug benefit (Part D). The findings from our study suggest that high consumer cost-sharing might not be as effective a mechanism for controlling spending as previously thought.

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Figure 1: Age Group Distribution for Employer-Insured Population, Medstat Sample and Our Study

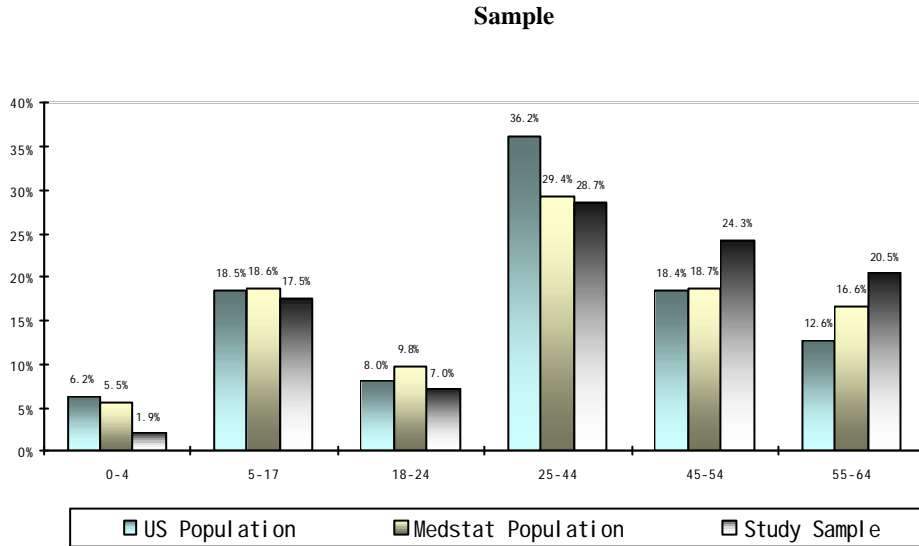
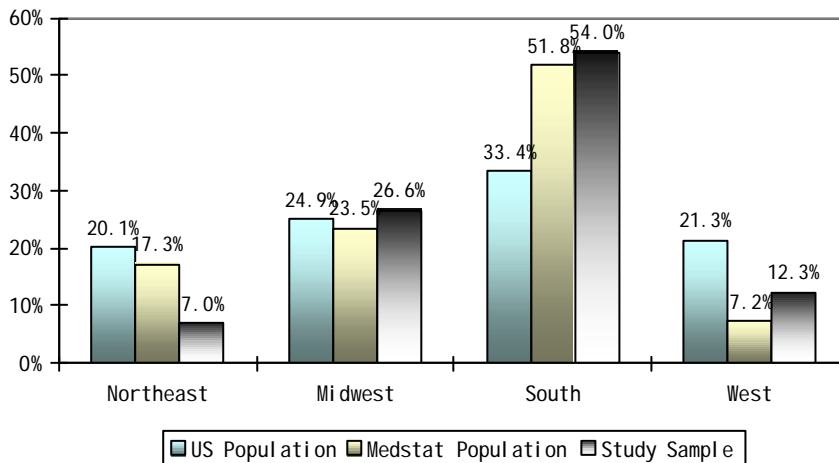
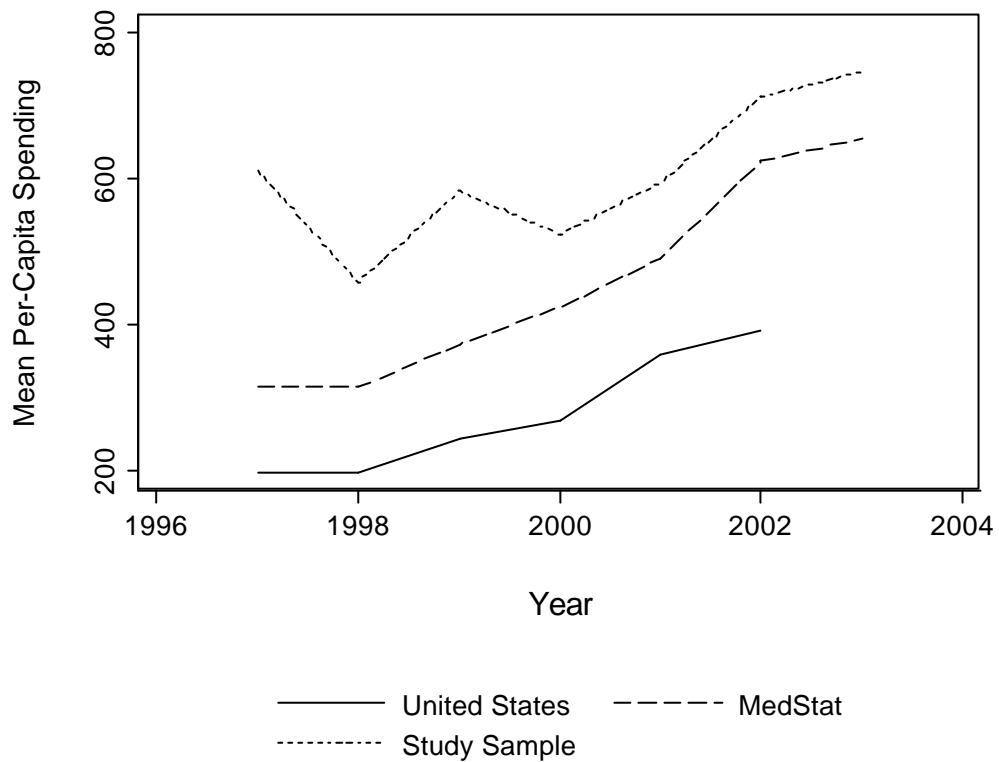


Figure 2: Regional Distribution for Employer-Insured Population, Medstat Sample and Our Study Sample

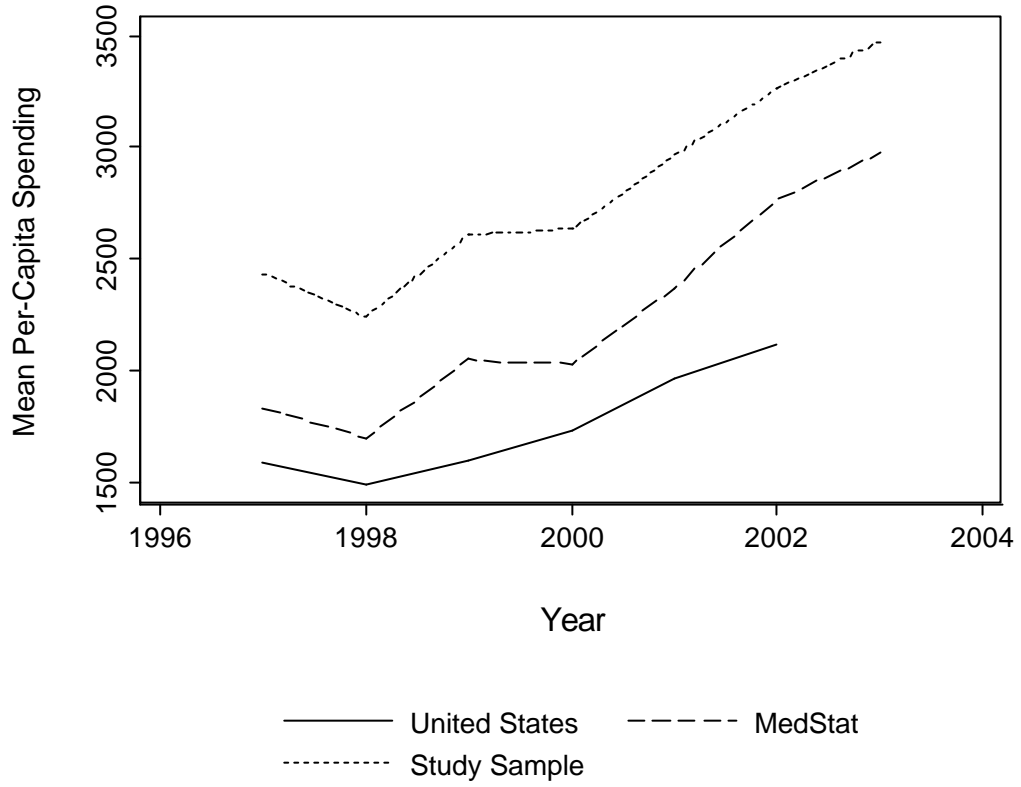


**Figure 3: Per Capita Pharmaceutical Spending for Employer-Insured Population,
Medstat Sample and Our Study Sample**



Note: The prescription drug spending data for year 2003 isn't available in the MEPS data at the time of this research.

**Figure 4: Per Capita Total Medical Spending for Employer -Insured Population,
Medstat Sample and Our Study Sample**



Note: The prescription drug spending data for year 2003 isn't available in the MEPS data at the time of this research.

Figure 5: Average Co-payments for Generic, Preferred Branded and Non-Preferred Branded Drugs

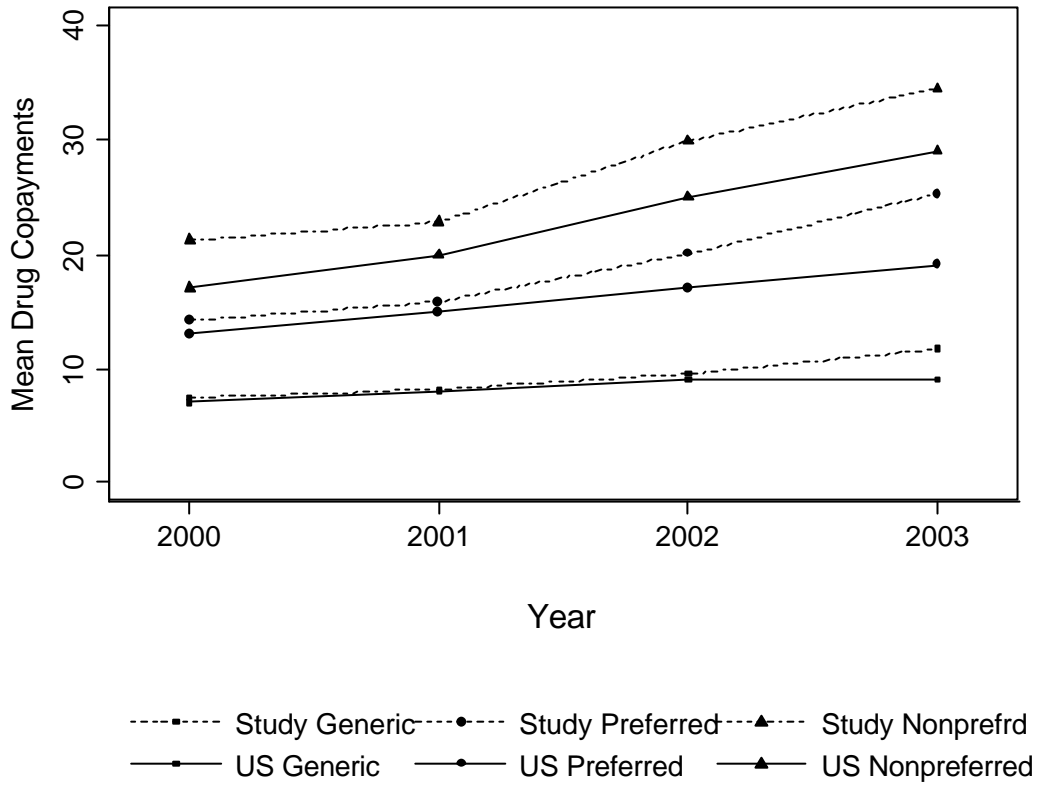


Figure 6: Trend in Prescription Drug Spending and Prices

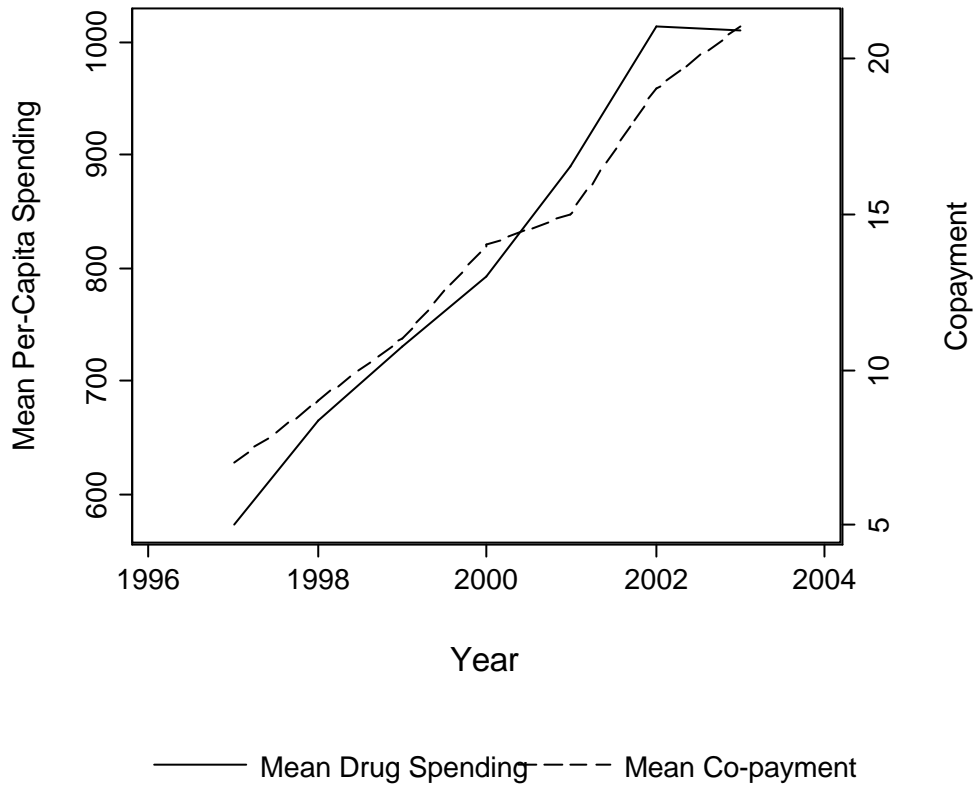


Table 1: Weights Used to Calculated Price Index for Prescription Drug Purchase

	Generic	Preferred Brand	Non-Preferred Brand
CARD PLAN	0.3580	0.3681	0.0908
MAIL-ORDER	0.0603	0.1025	0.0204

Table2: Prescription Drug Benefits for the Firms in Study Sample

		1997	1998	1999	2000	2001	2002	2003
FIRM1	Card Plan	\$4-\$4	\$4-\$8	\$4-\$8				
	Mail Plan	\$4-\$4	\$4-\$8	\$4-\$8				
	Drug Price	4.00	6.33	6.33				
FIRM2	Card Plan			\$12-\$12	\$12-\$12	\$12-\$12	\$12-\$12	\$12-\$16
	Mail Plan			\$12-\$12	\$12-\$12	\$20-\$20	\$20-\$20	\$20-\$36
	Drug Price			12.00	12.00	13.47	13.47	17.27
FIRM3	Card Plan		\$5-\$10	\$5-\$15	\$5-\$15-\$25	\$5-\$15-\$25	\$7-\$20-\$40	
	Mail Plan		\$6-\$12	\$6-\$18	\$10-\$30-\$40	\$10-\$30-\$40	\$14-\$40-\$70	
	Drug Price		8.22	11.25	14.08	14.08	19.87	
FIRM4	Card Plan			\$5-\$10-\$25	\$5-\$10-\$25	\$5-\$15-\$30	\$5-\$15-\$30	\$5-\$15-\$30
	Mail Plan			\$10-\$20-\$45	\$10-\$20-\$45	\$10-\$30-\$70	\$10-\$30-\$70	\$10-\$30-\$85
	Drug Price			11.31	11.31	15.60	15.60	15.90
FIRM5	Card Plan			\$4-\$12	\$8-\$18	\$10-\$20-\$30	\$10-\$25-\$35	\$11-\$27-\$42
	Mail Plan			\$12-\$36	\$24-\$54	\$20-\$40-\$60	\$20-\$50-\$70	\$22-\$54-\$84
	Drug Price			12.09	19.21	20.20	23.72	26.27
FIRM6	Card Plan	\$7-\$7	\$9-\$9-\$15	\$9-\$9-\$15				
	Mail Plan	\$7-\$7	\$9-\$9-\$15	\$9-\$9-\$15				
	Drug Price	7.00	9.67	9.67				
FIRM7	Card Plan	\$5-\$10	\$8-\$16	\$8-\$16	\$8-\$16-\$25	\$8-\$16-\$25	\$10-\$20-\$40	\$10-\$20-\$40
	Mail Plan	\$5-\$10	\$8-\$16	\$8-\$16	\$8-\$16-\$25	\$8-\$16-\$25	\$15-\$30-\$60	\$15-\$30-\$60
	Drug Price	7.91	12.66	12.66	13.66	13.66	19.78	19.78
FIRM8	Card Plan	\$5-\$10	\$5-\$10	\$5-\$10	\$5-\$10	\$6-\$15	\$6-\$15	\$6-\$15
	Mail Plan	\$8-\$15	\$8-\$15	\$8-\$15	\$8-\$15	\$9-\$25	\$9-\$25	\$9-\$25
	Drug Price	8.71	8.71	8.71	8.71	12.65	12.65	12.65

	Card Plan	\$10-\$10	\$10-\$10	\$10-\$20
FIRM9	Mail Plan	\$15-\$20-\$30	\$15-\$20-\$30	\$20-\$45-\$60
	Drug Price	11.74	11.74	19.80

Note: The \$X-\$Y or \$X-\$Y-\$Z structures represent the 2-tier and 3-tier co-payment schedules, respectively, where \$X denotes co-payment for generic drugs; \$Y denotes co-payment for brand drugs in a 2-tier schedule, and the co-payment for preferred brand drugs in a 3-tier schedule; \$Z denotes co-payment for non-preferred brand drugs in a 3-tier schedule. Drug prices are the calculated price index for each insurance plan based on the co-payments and population weights of each types of prescription.

Table3: Descriptions of variables

Variable	Description
RXS	Annual spending on prescription drugs
OUTS	Annual spending on outpatient services
INS	Annual spending on inpatient services
DAYSUPP	Prescription drugs demanded (in days)
NUMV	Number of outpatient visits
NUMADM	Number of inpatient admissions
RXP	Out of pocket prices for prescription drugs
DEDUCT	Deductible for medical services (for outpatient and inpatient)
COPAY	Co-payment for one physician office visit
AGEGROUP	Age group, 1=0-10, 2=11-18, 3=19-29, 4=30-39, 5=40-49, 6=50-59, 7=60-64
AGE1YEAR- AGE7YEAR	The interaction of age group with Year
RETIRE	1 = yes, 0 = no
URBAN	Residence place of employees, 1 = Urban, 0 = Rural
YEAR97- YEAR03	Dummy variables for year of data. Year97=1 if year=1997, etc.

Table4: Summary Statistics of Variables

Variable Name	Mean	Std	% of Zeroes	% of People Always Users	% of People Always Non-User
RXS	624	1756	29.52%	52.24%	12.94%
OUTS	1614	5156	16.72%	68.18%	6.05%
INS	714	6918	95.78%	0.42%	87.35%
TOTS	2951	10242	13.00%	74.83%	4.45%
DAYSUPP	331	584	29.52%	52.24%	12.94%

NUMV	16.61	27.83	16.72%	68.18%	6.05%
NUMADM	0.06	0.33	95.78%	0.42%	87.35%
RXP	15.30	5.32	-	-	-
DEDUCT	134	168	-	-	-
COPAY	6.50	7.06	-	-	-
URBAN	0.78	0.41	-	-	-
RETIRE	0.18	0.39	-	-	-
AGE1TIME	0.71	2.13	-	-	-
AGE2TIME	0.83	2.31	-	-	-
AGE3TIME	0.57	1.96	-	-	-
AGE4TIME	0.98	2.48	-	-	-
AGE5TIME	1.47	2.93	-	-	-
AGE6TIME	1.58	2.98	-	-	-
YEAR98	0.05	0.22	-	-	-
YEAR99	0.11	0.32	-	-	-
YEAR00	0.19	0.39	-	-	-
YEAR01	0.22	0.41	-	-	-
YEAR02	0.22	0.41	-	-	-
YEAR03	0.17	0.38	-	-	-
N			1713879		
# of Individual			526086		

Table 5: Percentage of People Switched Between Insurance Plans

		<i>97</i>	<i>98</i>	<i>99</i>	<i>00</i>	<i>01</i>	<i>02</i>	<i>03</i>
Firm 1	Medical	-	0	0				
	Drug	-	0	0				
Firm 2	Medical			-	0	0.11	1.8	2.18
	Drug			-	0	0	0	0
Firm 3	Medical		-	0	2.88	1.11	1.14	
	Drug		-	0	0	0	0	
Firm 4	Medical			-	0.13	0.19	0.29	0.97
	Drug			-	0	0	0	0
Firm 5	Medical			-	0	7.13	0	0
	Drug			-	0	0	0	0
Firm 6	Medical	-	0	0				
	Drug	-	0	0				

Firm 7	Medical	-	1.13	0.99	0.92	1.17	0.39	0
	Drug	-	0	0	0	0	0	0
Firm 8	Medical	-	0.45	0.34	0.14	0.58	0	0
	Drug	-	0	0	0	0	0	0
Firm 9	Medical					-	0	0
	Drug					-	0	0

Note: The number in each cell represents the percentage of people within a firm who switch their medical insurance benefits or prescription drug insurance benefits from previous year to the indicated year.

Table 6: Estimation Results for Spending Equations

	FE OLS		FE Tobit		FE Tobit	
	Point Estimate	Elasticity	Point Estimate	Elasticity	Point Estimate	Elasticity
<u>Rx Spending</u>						
RxP	-20.34*** (0.52)	-0.50	-33.5*** (1.03)	-0.82		--
RxP_1	-13.49*** (0.75)	-0.33	-12.7*** (1.27)	-0.31		
<u>Outpatient Spending</u>						
RxP	10.28*** (2.34)	0.10	7.7*** (2.96)	0.07		--
RxP_1	12.70*** (4.05)	0.12	11.3*** (3.46)	0.11		
<u>Inpatient Spending</u>						
RxP	-3.45 (4.32)	-0.07	183.2** (96.68)	0.19 †	-5.47** (2.35)	-0.12
RxP_1	-4.97 (5.56)	-0.11	616.6*** (99.16)	0.63 †	-6.43*** (2.54)	-0.14
<u>Total Spending</u>						
RxP	-13.51*** (5.01)	-0.07	-24.0*** (6.60)	-0.12		
RxP_1	-5.76 (6.01)	-0.03	-0.4 (7.56)	0.00		--
Obs	1713879		1713879		1713879	
# Group	526086		526086		126	
Individual FE	Yes		Yes		No	

Note: † These elasticities are calculated using the mean dollar of inpatient spending for inpatient care users. The average inpatient spending for users is \$15083 per year, and for the whole population is \$714 per year.

Significance is based on bootstrapped standard errors which cluster errors at either individual or group level.

- *** Statistically significant at 0.001 level.
- ** Statistically significant at 0.05 level.
- * Statistically significant at 0.1 level.

Table 7: Estimation Results for Quantity Demand Equations

	FE OLS		FE Poisson		<i>FE Poisson</i>	
	Point Estimate	Elasticity	Point Estimate	Elasticity	Point Estimate	Elasticity
<u>Days Rx Supply</u>						
RxP	-11.47*** (0.18)	-0.53	-15.7*** (0.3)	-0.73		--
RxP_1	-8.60*** (0.21)	-0.40	-10.6*** (0.2)	-0.49		
<u>Outp Visits</u>						
RxP	0.05*** (0.01)	0.05	-0.026** (0.013)	-0.02		--
RxP_1	0.18***	0.17	0.101***	0.09		

	(0.02)		(0.01)			
<u>Hospital Admission</u>						
RxP	-0.0003 (0.0002)	-0.08	-0.0026 (0.0016)	-0.03 †	-0.0006*** (0.0002)	-0.15
RxP_1	-0.0005** (0.0002)	-0.13	-0.0043** (0.0018)	-0.05 †	-0.0009*** (0.0002)	-0.23
Obs	1713879		1713879		1713879	
# Group	526086		526086		126	
Individual FE	Yes		Yes		No	

Note: † These elasticities are calculated using the number of inpatient admissions for inpatient care users. The average inpatient admission for users is 1.29 admissions per year, and for the whole population is 0.06 per year.

Significance is based on bootstrapped standard errors which cluster errors at either individual or group level.

- *** Statistically significant at 0.001 level.
- ** Statistically significant at 0.05 level.
- * Statistically significant at 0.1 level.