CHARGING LOW-INCOME FAMILIES FOR HEALTH INSURANCE: HOW DOES IT IMPACT CHOICES?

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As health care reform moves forward in the United States, one common feature of virtually all proposals is to expand coverage for low income populations not through a traditional public insurance model, but rather through an "defined contribution exchange" mechanism. Under this approach, low income individuals have a choice of a number of options for their insurance coverage. Individuals receive a subsidy to purchase insurance that was tied to the lowest-cost plan (or some index of low-cost plans) and pay some part of the difference if they chose a more expensive plan.

This major departure from the traditional free/single-choice public payer model raises a number of important questions related to firm behavior and competition, adverse selection, and welfare, but the key initial question is: How do low-income consumers respond to differences prices across plans? If price sensitivity is low, for example, then choice is less likely to lead to cost-reducing competition among plans. If price sensitivity is very much related to enrollee health, then it suggests that plans that have different prices may also enroll patients of different levels of health, leading to adverse selection in which only sicker patients remain in more expensive plans (Cutler and Reber, 1998).

In this paper, we study the plan choice of low-income enrollees in Massachusetts' Commonwealth Care program that was established as part of the state's health reform in April, 2006. Enrollees in Commonwealth Care were given a choice of up to four Medicaid Managed Care Organizations (MMCOs) from which they could receive their coverage. For about half of enrollees (those below the poverty line), this decision had no financial implications. But for the remainder, enrollees were charged not only a base contribution rate, but the differential cost of their plan choice over the lowest-cost plan in their area. The financial implications of this decision were non-trivial; the average differential in 2007 between the non-lowest-cost plan and the lowest-cost plan across areas and income groups was $18.83 per month, and the maximum was $116 per month. For the relevant population, those with incomes from 100% to 300% of poverty line, these are meaningful amounts.

Most relevantly for our study, there was a major shift in the pricing of plans for open enrollment in June 2008. Before that time, all enrollees below 150% of the federal poverty line received health insurance for free regardless of cost. However, after June 2008, enrollees between 100% and 150% of poverty were responsible for the full differential in cost between the lowest-cost plan and the plan that they choose. In addition to this major change, most other plan types for individuals above 150% of poverty also experienced an increase in the base contribution rate charged enrollees, due to new bids by the MMCOs. At that open enrollment, the cost of the average plan increased by about $18.83, with a standard deviation of $30.22; in addition, the range in cost between the highest- and lowest-cost plan increased from 2007 to 2008. However, the ranking of many plan prices were reversed from the previous year, such the correlation with previous-year prices was negative for many plan types (Table 1).

We have created a unique data set using information from the state of Massachusetts on the enrollment decision of each Commonwealth Care recipient over 2007 and 2008. We have data both on those enrollees who were already in a plan as of June 2007 and faced the decision over whether to switch plans ("prior enrollees"), and those who were newly choosing across plans through 2008 ("new enrollees"). For each enrollee we have information on their income and an index of their underlying medical-spending risk. We then estimate conditional logit models on these data, using the changes in pricing described above, to understand how price differentials impact plan choice.

We find that these low-income populations are highly sensitive to plan price differentials. Our central estimate, accounting for endogeneity and heterogeneity, suggests that an increase in monthly out-of-pocket costs of $10 for a given plan in 2008 decreases enrollment by 11.4% for prior enrollees who are deciding whether to switch plans and by 19.4% for new enrollees who are deciding on a plan after entering Commonwealth Care. These are indeed substantial numbers of enrollees for a relatively small difference in price. Because out-of-pocket costs are low in general and even free for many consumers, the implied price elasticities of demand are -0.71 for prior enrollees and -0.75 for new enrollees.

Our paper proceeds as follows. Section I describes the institutional background on Commonwealth Care and the price changes we study. Section II discusses our data and baseline results. Section III presents discusses evidence related to endogeneity. Section IV extends the results to account for patient health and heterogeneity, and Section V concludes.

I: INSTITUTIONAL BACKGROUND

The groundbreaking health care reform passed in Massachusetts in 2006 had a number of important features, including a mandate on individuals to purchase insurance and a reform of non-group and small group insurance markets. Most important for our purposes, the law established the Commonwealth Care program for those in families with incomes below three times the poverty line (roughly $30,000 for singles and $60,000 for a family of four at the time of the law's passage). Only individuals who were not eligible for other coverage (employer-sponsored insurance or Medicaid) could enroll.

Starting in mid-2007, the first full year of the program, individuals were placed in one of six "plan types" depending on their income. Plan types were differentiated by the patient cost-sharing imposed in the plan and by enrollee contribution rates.[1] Those below poverty were in plan type I; those who were 100-150% of poverty were in plan type IIA; and those who were 150-200% of poverty were in plan type IIB. For those 200-300% of poverty, there was a choice in 2007 between two different benefits structures: plan type III with higher copayments and a lower premium cost to enrollees, and plan type IV with lower copayments and a higher premium cost to enrollees; within plan types III and IV there was a division into IIIA/B and IVA/B at 250% of poverty.

Enrolling individuals had a choice of up to 4 Medicaid Managed Care Organizations (MMCOs); in some areas of the state the choice set was smaller due to limited regional coverage of some MMCOs. For example, by 2008, enrollees in Western Massachusetts generally only had 2 choices, while enrollees in Northern Massachusetts predominantly had 4 choices and enrollees elsewhere, including Boston, had on average 3 choices. In 2007, individuals below 150% of the poverty line were enrolled in plan type I or plan type IIA for free, and could enroll in any of the available MMCO at no personal cost. Individuals in the remaining plan types had to pay a base contribution for the lowest-cost plan available in their area, as well as the full differential in the cost of choosing any plan that was above the lowest-bidding plan. The base contribution was $35 for plan type IIB, $70 for PT IIIA, and $105 for PT IIIB. Enrollees choosing plan types IVA and IVB had the same base contributions as those in plan types IIIA and IIIB, respectively, since they were of the same income groups, but they because they chose the lower copayment plans, premium contributions for all plans in IVA or IVB were above the base contribution.

In addition to enrollees who explicitly chose a plan, individuals below 150% of the poverty line who were deemed eligible but did not choose a plan were auto-enrolled. The auto-

[1] Cost sharing arrangements and coverage of services were specified and mandated by the Commonwealth Connector Authority.

enrollment algorithm could randomize auto-enrollees among several possible plans but was weighted towards low-cost plans. The MMCOs themselves made bids in early 2007 for the prices they would charge the state for each demographic group in plan types IIA/B, IIIA/B, and IVA/B.[2] Costs were calculated for each region based on the demographic composition of each region. Incentives to bid low therefore came from both the assignment of auto-enrollees and the financial incentives for enrollees to choose low-cost plans.

The system then changed in several important ways for open-enrollment in June, 2008. First, those in plan type IIA still could sign up for the lowest-cost option for free, but now had to pay the full differential for choosing a more expensive plan in their area. Second, plan types IVA and IVB, the more expensive plans with lower copayments for individuals at 200-300% of FPL, were discontinued. Once again, prices changed as MMCOs made bids for each demographic group in plan types IIA/B and IIIA/B.

The resulting pricing changes are illustrated in Table 1. This Table shows, for each plan type, the mean and standard deviation of the change in contribution for the typical enrollee to stay in the same plan in 2008 as in 2007, as well as the mean and standard deviation of the change in contribution for the typical enrollee to move from the lowest- to highest-priced plan in the area. Enrollees in plan types IIA and IIB experienced an average increase in contributions of $9.14 and $14.04, respectively, while enrollees in plan types IIIA and IIIB experienced an increase in contributions of $48.66 and $50.85, respectively. While plan types IVA and IVB were discontinued in 2008, enrollees in these plan types still would have experienced average price increases of $16.08 and $21.47, respectively, assuming that they continued in plan types

[2] Because plans could only jointly bid for plan types IIA and IIB, for example, the only difference in effective enrollee contributions between plan types IIA and IIB is due to the difference in base contribution rates, so that prices in plan types IIA and IIB are collinear. The same is true of plan types IIIA and IIIB and of plan types IVA and IVB.

IIIA and IIIB, respectively. The range between the lowest- and highest-cost plans also increased from 2007 to 2008. Most notable are enrollees in plan type IIA, who faced no price differential in 2007 and then a $24.15 differential in 2008. Finally, for plan types other than IIA/B, there was a negative correlation between 2007 and 2008 prices. Although it is beyond the scope of this paper to fully characterize firm pricing behavior, we will argue later that this is consistent with firms responding to auto-enrollment in plan type IIA (which has price a collinear with IIB) and general consumer inertia after initial enrollment in the higher plan types.

II: DATA AND BASELINE RESULTS

Data

In order to assess the impact of this change in relative plan prices on plan choice, we have collected three sets of data with the helpful assistance of the staff of the Massachusetts Health Connector. The first set of data is information on all those who were enrolled in Commonwealth Care continuously from June 2007 to September 2008, a total of 75,184 “prior enrollees.” We exclude from our analysis 4,971 prior enrollees who had only one plan in their choice set, as well as an additional 3,256 who were auto-enrolled and therefore did not choose their plan. Table 2 shows the number of remaining 66,957 enrollees in each plan type and MMCO for both 2007 (as of June 2008 prior to open enrollment) and in September 2008 (after open enrollment), by plan type and MMCO. Among the MMCOs, we see that Boston Medical Center HealthNet Plan (BMC) had the largest group of enrollees, while Fallon Community Health Plan (Fallon) was just entering the market at this time. Also, the greatest proportion of enrollees above poverty was just above poverty, in plan type IIA or at 100-150% of poverty.

For each prior enrollee, we have data on demographic characteristics (age and sex), health care utilization, area of residence, original plan choice in 2007, and new plan choice in 2008. Area of residence was categorized by the Commonwealth Connector Authority into 5 regions and 38 areas. Of note, we have three actuarial measures of health risk for prior enrollees, one based on enrollee demographics, another based on health care utilization, and a third based on a combination of enrollee demographics and health care utilization.

The second dataset contains information on all first-time enrollees during 2008, which includes a total of 52,305 “new enrollees” above poverty. We exclude an additional 3,161 with only one plan in their choice set and another 10 auto-enrollees (virtually all auto-enrollees were below poverty). Table 3 presents similar information on the remaining new 49,134 new enrollees in 2008 as does Table 2, breaking down numbers of new enrollees by plan type and MMCO. For each new enrollee, we have similar data as for prior enrollee, with the exception of health utilization data and previous plan choices, since they are enrolling in Commonwealth Care for the first time. Finally, our third set of data is at the plan level categorized by area, plan type, and MMCO. These data contain enrollee contributions for each plan choice available to them based on their income and area of residence.

Using this information, we construct conditional logit models to describe the discrete choice that old and new enrollees faced. In our base specification, we modeled the utility of prior enrollees as a function of whether the enrollee had chosen that plan previously in 2007, the contribution price of the choice in 2008 during open enrollment, and plan dummies:

$$u_{ij}^{old} = \alpha_1 Same_{ij} + \alpha_2 P_{08,j} + \beta Plan_j + \varepsilon_{ij},$$

where i indicates the enrollee and j indicates the plan. $Same_{ij}$ is a dummy for whether prior enrollee i was enrolled in plan j in 2007; $P_{08,j}$ is the contribution price for plan j in 2008; $Plan_j$ is a vector of plan dummies (with BMC normalized to 0); and ε_{ij} is an error term distributed independently and identically as extreme value.

For new enrollees, we use a similar specification, although one that is simpler than for prior enrollees because they do not have a prior plan:

$$u_{ij}^{new} = \gamma P_{08,j} + \delta Plan_j + \varepsilon_{ij},$$

In our baseline results we assume that prices are set exogenously and also do not consider heterogeneity in preferences based on income group or health status, issues we will address below. For models here and later, we use robust standard errors accounting for clustering at the area level.

Baseline Results

The first column of Table 4 presents results for the baseline regression of prior enrollee choices. We find highly significant coefficients both on whether the enrollee has been previously enrolled in the choice (the same-plan dummy) and on price, where the same-plan dummy is by far the most important determinant of choice. The relative magnitudes of the same-plan dummy coefficient and the price coefficient imply that a price difference of about $84 would be required for a prior enrollee to be equally likely to switch out of his prior plan, all else equal.

The implication of the price coefficient is presented in the bottom rows of Table 4. Weighted across the population and over choices by the likelihood of an enrollee to pick a choice, we estimate that on average a $10 increase in the price of a plan lowers the probability that an individual chooses that plan by 8.7% in relative terms. This is a very large effect, but the $10 increase is also a relatively large change compared to baseline, especially considering that some enrollees pay nothing. As a result of many of these low prices, the implied average elasticity, calculated for each individual and choice by standard calculations (Train 2003) and similarly weighted, is somewhat lower at -0.67.

The second column in Table 4 present regression results for new enrollees. The coefficient for this population is smaller, although we cannot directly compare effects to the prior enrollees since we do not have a control for prior plan choice. But the implied effect of a change in plan prices is larger: we find that a $10 increase in the price of a plan lowers the odds that the plan is chosen by a new enrollee by 15.4%, for an elasticity of -0.72.

III: ENDOGENEITY

In our baseline regressions we have take prices as exogenous, but it is certainly plausible that MMCOs could be making bids endogenously. Although plans in each plan type are required to provide the same level of coverage for the same types of services at the same copayments, there still may be omitted variables such as reputation, advertising, or the number of physicians or hospitals accepting a given plan. For endogeneity to influence our baseline results, we must have endogeneity at the plan- and area-specific level, because we have included plan dummies that will account for any statewide differences in plan preferences. Furthermore, we note that some aspects of plan price are exogenously set by the state, such as the baseline contribution, and that plans may only technically bid by demographic group in each region, which was converted to an “effective bid” for each region by the Commonwealth Connector Authority. Endogeneity in price across areas that arises from MMCO bidding is therefore restricted to the space that such a bidding mechanism can cover.

Nonetheless, it appears that firms are able to set prices in a manner that is consistent with (potentially constrained) profit maximization. We have shown above that once enrollees are in a given plan, they are very likely to stay in that plan in the next year, despite potentially large differences in price.[3] If plans do not systematically differ by much in quality, it can be shown in a model of horizontal differentiation that initially cheaper plans will subsequently become relatively more expensive than other plans because of the larger mass of consumers it attracted in the first period.[4] This is consistent with what we find in Table 1, where there is in fact a negative correlation between prices in the initial and subsequent years for higher plan types. At the same, we note that the Commonwealth Connector Authority auto-enrolls members in plan type IIA mechanistically based on prices and not on which plan that member was auto-enrolled in previously. Thus, we see that for plan type IIB, which has a collinear price with plan type IIA, there is a positive correlation in prices between the two years.

We will evaluate the possible impact of such endogeneity in prices in two ways. First, we will attempt to add control variables that capture such endogeneity and evaluate the effect on the price coefficient of interest. Second, we will adopt a control function approach by using plausible instrumental variables, namely the price of the plan in other regions, first used by Hausman (1997).

Control Variables

It is difficult to think of control variables that would capture all endogeneity over area, plan type, and time, but we may evaluate the effect of including variables that potentially capture

[3] Although results are not shown here, the propensity to stay in a previous plan remains even when considering enrollees who have previously been auto-enrolled but are now choosing plans.
[4] In order to have prices differ in the first period of such a model, we would rely on some degree of imperfect pricing ability, such that firms cannot price exactly at equilibrium. Alternatively, some parameters yield mixed strategy equilibria.

some of the effect of omitted variables and see if the coefficient on price changes as a result. First, we consider the price of a plan in the previous year for prior enrollees. If omitted variables do not change over time for a given plan and area, or if they change in a manner independent of pricing, then previous-year price will capture all endogeneity. This does not seem plausible, given the predicted and observed firm pricing behavior that we have noted above, but perhaps previous-year price reflects some of endogeneity that affects prices and choices. Second, we consider the price of another (non-collinear) plan type by the same MMCO in the same area. If unobserved variables are the same across different plan types for a given MMCO and area (i.e., higher-income enrollees enjoy the same unobserved benefits as lower-income enrollees), then this measure would be plausible control variable.

Results are presented in Table 5. In the third column, model 4 considers our first control variable of previous-year price. We find a positive and significant coefficient, which is nearly 60% of the magnitude of current price. This may reflect unobserved differences across plans and areas, or alternatively, it may reflect decisions in which enrollees incorrectly *infer* quality from previous-year prices as a "rule of thumb," either by irrationality or imperfect information. When comparing the coefficient on current-year price with the baseline regression (model 1), there is no evidence of a statistically significant difference either by the Hausman test or by bootstrapping[5]. The fourth column (model 5), presents results using both previous-year price and the price of plan types IIIA/B for enrollees in plan type IIA.[6] Again, we find no evidence of a difference in coefficients either by the Hausman test or by bootstrapping. Finally, we evaluate the effect of omitting statewide plan dummies in the second column (model 3) and, as expected, find a significant difference in price coefficients with the baseline model.

[5] We perform bootstrapping in cases where the Hausman test statistic is negative in this finite sample. We perform 250 bootstrap repetitions for each model (using the same random numbers for set of 250 repetitions).
[6] We perform similar regressions for each plan type, all with similar results.

Next, we consider another set of tests using control variables on the choices of new enrollees who enrolled before or after June 2008. Significantly, during open enrollment, prices for the new fiscal year were announced and implemented only during June 2008, so that new enrollees who enrolled before June 2008 would not know of later prices but would obviously be affected by them, while those who enrolled after June would know both current and previous prices but only be financially affected by current prices. In the first two columns of Table 6, we evaluate whether previous-year price affects those who enrolled after June 2008. Although as before we find a large positive and significant coefficient on previous-year price, we cannot show that this changes our estimate of current-year price. In the next two columns, we evaluate whether future price affects the decisions of those who enrolled before June. We find a positive but much smaller coefficient that interestingly also causes a change in the current-year price coefficient.[7] Taken together, this evidence is consistent with some endogeneity but also a likely larger behavioral "rule of thumb" referencing of previous prices that consumers use to make decisions.

Instrumental Variables and Control Function Approach

We next use a control function approach to account for possible endogeneity. This approach is appealing because of its simplicity and easy applicability to individual-level data, especially given the importance of the effect of previous-plan enrollment in choices (Petrin and Train 2006). In order to form the control function, we perform a first-stage regression using instruments of prices in other regions under the same MMCO, which was first introduced by Hausman (Hausman 1997) and subsequently used elsewhere (Goolsbee and Petrin 2004). Such

[7] As a robustness check, we perform another regression on the choices that prior enrollees made in 2007 and including 2008 prices as future prices, and we find similar results.

instruments will be valid if we allow firms to be able to price discriminate across regions and if we assume that any statewide endogeneity in plans is absorbed by plan fixed effects. We then enter the first-stage residuals as a "control function," which we assume to enter linearly into utility.[8] For example, for prior enrollees, we formulate the model as

$$u_{ij}^{old} = \alpha_1 Same_{ij} + \alpha_2 P_{08,j} + \beta Plan_j + \lambda \mu_j + \varepsilon_{ij}.,$$

where μ_j is the control function formed from the first-stage residuals.

Table 7 presents results using this control function approach. The first column (model 9) shows a significantly positive coefficient on the control function for prior enrollees. Following Rivers and Vuong (Rivers and Vuong 1988), we take the significant coefficient on the control function to indicate evidence of endogeneity. We also interpret the positive sign to mean that plan-areas with higher quality are priced more expensively and note that "quality" may also indicate the network externalities of having enrolled many other people in the region, independent of the effect of the same-plan dummy for the individual. Finally, we note that the coefficient on price is slightly more negative but not by an impressively amount compared to the baseline model. The second column in Table 7 shows that previous-year price remains significant even when controlling for endogeneity, again suggesting that previous price also enters independently into decision-making. The third column shows similar results for new enrollees. Finally, the bottom two rows show the implied average own-price elasticities and the percentage responses to a $10 increase in own-price. After accounting for endogeneity, responses are generally slightly greater than the baseline figures in Table 4.

[8] Thus far, we have only considered a homogeneous utility function with fixed coefficients, but this is also a valid model under heterogeneous utility if we assume that the error term ε_{ij} in our model consists of a normal component that is correlated with price with mean $\lambda\mu_j$ and an i.i.d. extreme value component. We will address this below under heterogeneity.

IV: HETEROGENEITY

We have thus far described enrollee decisions in aggregate, but in order to evaluate the existence and importance of potential adverse selection, we must also characterize how behavior might differ among individuals, including individuals distinguished by health risk. We will explore heterogeneity in preferences, specifically with regards to health risk, first by extending the baseline conditional logit model to account for risk and income, and then by quantifying unconditional and conditional variance of the distribution of random price coefficients in mixed logit models.

Extension of Baseline Models

To account for health and income groups, we extend our baseline models for prior enrollees to

$$u_{ij}^{old} = \alpha_1 Same_{ij} + \alpha_2 P_{08,j} + \alpha_3 P_{07,j} + \alpha_4 P_{08,j} Q_i + \alpha_5 P_{08,j} PT_i + \beta Plan_j + \varepsilon_{ij},$$

where $P_{07,j}$ is the contribution price for plan j in 2007; Q_i and PT_i are vectors of dummies corresponding to quartiles of health (with lower quartiles signifying healthier patients) and plan type for enrollee i, respectively; and ε_{ij} is identically and independently distributed as extreme value.[9] Likewise, we specify a more general model for new enrollees:

$$u_{ij}^{new} = \gamma_1 P_{08,j} + \gamma_2 P_{08,j} Q_i + \gamma_3 P_{08,j} PT_i + \delta Plan_j + \varepsilon_{ij},$$

where the variables are similarly defined.

Note that for prior enrollees, we use original plan type, which can include the discontinued plan types IVA and IVB. Because enrollees of the same income could have

[9] We also took two alternative approaches, both with similar results. First, we evaluated a more general model with interactions between income and health. Second, we implemented a "local maximum likelihood" approach by fitting the baseline model at each quantile of health risk, smoothed by neighboring quantiles with an Epanechnikov kernel similar to that of Fan et al (1998).

originally chosen either IIIA or IVA, or either IIIB or IVB, this also serves as an indicator of preferences in addition to income. For prior enrollees, we also consider two measures as proxies for health, one defined by health care utilization and another by demographics; for new enrollees, we only have access to the demographic-based measure, since we have not yet observed their health care utilization.

Table 8 presents results for these regressions. In general, we see that enrollees who are poorer and healthier are more price sensitive, suggesting that they assign less value to health insurance relative to its price. For prior enrollees, we generally do not see much change in the price coefficient with health as we do with income group.[10] In contrast, we see a much greater relationship between health risk and price sensitivity among new enrollees, who also do not show much difference in price sensitivity by income group among plan types above IIA. In addition, we note that health based on utilization data (models 13 and 15) is more accurate in characterizing price preferences for insurance than the demographic-based measure (models 12 and 14) for prior enrollees. Although this is not surprising, it makes the relationship between health and price sensitivity all the more striking for new enrollees, since health for new enrollees is only estimated by demographic measures.[11]

In addition to conditioning income and health risk, we are also able to observe another aspect of preference heterogeneity by comparing prior enrollees who originally chose plan types IVA or IVB versus those who chose plan types IIIA or IIIB for income groups at 200-300% of poverty. Plan-type choice in 2007 between low-premium (plan types IIIA and IIIB) and high-

[10] Of course, the support of prices is different for different income groups, such that it is not possible for an enrollee in plan type IIIA to purchase a plan for free, for example. Thus, coefficients may take on different meanings for different plan types, since enrollees are choosing within pre-specified ranges in proportion to their incomes, as defined by the federal poverty level. In this sense, heterogeneity of preferences according to income may be overstated.

[11] As a robustness check, we verify that the difference in price responsiveness according to health is not due to different enrollees entering over time by running the same general “new enrollee” regressions (models 16 and 17) on prior enrollees who were making decisions in 2007.

premium (plan types IVA and IVB) predicts price sensitivity in 2008 even when the choice is no longer available: enrollees who chose the high-premium, low-copayment plans are much less price-sensitive than enrollees with the same incomes but with the opposite choices in the previous year.

Finally, we also adjust for endogeneity by regressing on the control function described above. When comparing models that adjust for endogeneity (models 14, 15, and 17) with corresponding ones that do not (models 12, 13, and 17, respectively), we again see a significantly positive coefficient on the control function. As before, the price coefficient is more negative when adjusting for endogeneity, but notably, we see that the magnitude of this change is less than those observed when comparing heterogeneous subgroups.

Predicted Responses to Price Changes

As we did before, we use our results in Table 8 to calculate implied price responses among enrollees, both in terms of own-price elasticities and in terms of a response to a $10 increase in own-price. In addition to aggregate responses, we also specifically evaluate responses within subgroups of plan type and risk. Table 9 presents elasticities, and Table 10 presents percentage responses to a $10 increase in own-price. In both tables, we present results for the same models as before in columns, while we subgroup each population into rows according to health and plan type.

As expected, total aggregate responses are similar to those estimated before, although slightly different due to the nonlinearity of changes when accounting for heterogeneity. Also, we see that price responses, both in terms of elasticities and percentage responses to a $10 price change, uniformly increase when accounting for endogeneity with our control function. More interestingly, we see the translation of the price-coefficient effect of heterogeneity onto price responses. For example, we see that in model 17, new enrollees who are in the healthiest quartile would decrease their probability of choosing a plan by 24.1% in response to a $10 increase in the price of that plan, while the sickest quartile would only decrease this probability by 16.8%. On the other hand, as we saw with price coefficients, there is not much difference in responses among the prior enrollees according to health. Within income groups, elasticities and percentage responses to a $10 price increase largely follow the same trend, but we of course note that trends can be reversed when comparing different income groups, since prices are much higher for higher-income groups, and therefore we would be comparing higher absolute increases for a given percentage increase.

Most strikingly, we see the large differences in implied responses when comparing prior enrollees of the same income who originally chose plan types IVA or IVB over plan types IIIA or IIIB, respectively. Despite the discontinuation of plan types IVA/B, so that all enrollees with incomes 200-300% of poverty had to choose from the same options in plan type IIA or IIIB in 2008, we see roughly twice the price responsiveness among enrollees who originally chose IIIA or IIIB, the low-premium, high-copayment option, than among those who chose IVA or IVB, respectively. For example, in model 15, prior enrollees originally in plan type IVA would respond by -7.6% to a $10 increase, while those originally in plan type IIIA would respond by -17.3%, corresponding to elasticities of -0.87 and -1.94, respectively.

Unconditional and Conditional Heterogeneity

In our above analysis we see evidence of substantial heterogeneity in price responsiveness, some of which is conditional on health and income, but some of which appears to reflect preferences independent of observable characteristics. In an attempt to quantify the degree of heterogeneity in the price responsiveness, we fit mixed logit models to explicitly quantify the variance of a distribution of random price coefficients. We consider variants of the following the mixed logit model for prior enrollees:

$$u_{ij}^{old} = \tilde{\alpha}_{1,i} Same_{ij} - \tilde{\alpha}_{2,i} P_{08,j} + \alpha_4 P_{08,j} Q_i + \alpha_5 P_{08,j} PT_i + \tilde{\beta}_i Plan_j + \lambda \mu_j + \varepsilon_{ij}.$$

In this model, coefficients on the same-plan dummy, price, and statewide plan dummies are allowed to be random (and potentially correlated) among enrollees. We will assume that the coefficients on the same-plan dummy and negative price are log-normally distributed, so that they are guaranteed to be positive, while we assume that the coefficients on the statewide plan dummies are normally distributed.

We first estimate an unconditional model without any price interactions with health or income to quantify the unconditional variance of the coefficient on price, which reflects the unconditional heterogeneity of price responsiveness. Next, we fit other models conditioning the price coefficient by interactions with health and/or income, and we interpret their corresponding estimated variances as measures of conditional heterogeneity.[12] Estimation was done by simulated maximum likelihood using Halton sequences.[13]

Results of the mixed logit models are shown in Table 11, with means and standard deviations of the estimated distributions of random parameters in the first and second panels, respectively. We also report the fixed coefficient on the control function, which now can be interpreted as the conditional mean of the potentially random effect of endogeneity, given the

[12] Since we assume that the coefficients on the price interactions are fixed, the conditional variance of each subgroup is by construction the same in each model. Alternatively, we tried to fit an "unconditional" model for each subgroup, but we unfortunately could not get such a model to converge, possibly reflecting the increased data requirements of mixed logit. These barriers may potentially be surpassed with panel data.

[13] Because of the computational intensity of the mixed logit model, we randomly reduced the estimation sample to one-eighths of its original size. Our findings are robust to separate random reductions.

random coefficients on the plan dummies (see footnote 8).[14] Our results reject the null hypothesis of fixed coefficients for price in all models, including when we condition on interactions both with health and income.

Importantly for our purposes, we calculate in each model the variance of the (log-normally distributed) coefficient around negative price, and find that it indeed decreases as we condition on observable correlates of heterogeneity. Conditioning on health alone leaves us with a variance that is 65.1% of the unconditional variance, while conditioning on income results in a 38.0% variance.[15] Conditioning on both health and income leaves us with 28.2% of the unconditional variance, an arguably significant number.

IV: CONCLUSION

Using a unique dataset that includes all enrollees in the Massachusetts Commonwealth Care program during 2008, we examine how the availability of health-plan choice and financial responsibility affect the decisions of low-income families seeking subsidized health insurance. Specifically, we evaluate the response of low-income consumers to differences in the price of insurance, and whether price sensitivity differs among them. We find that low-income consumers are indeed price-sensitive. Own-price elasticities of -0.71 and -0.75 for prior enrollees and new enrollees, respectively, belie the fact that many of the plans that we study are quite cheap and some are even free, so that a percentage increase in price for these plans may represent small absolute increases. Indeed, a $10 increase in the monthly premium of a plan will on average decrease enrollment by 11.4% for prior enrollees potentially switching plans and by 19.4% for

[14] We omit reporting estimates for fixed coefficients on the interaction terms due to space; they are similar to estimates in Table 8.
[15] Here, we take income literally as percentage brackets of FPL (i.e., we do not consider plan types IIIA and IVA, or IIIB and IVB, separately), because we want to condition variance only on observable characteristics.

new enrollees who have decided to enroll for the first time. At the same time, we find that once enrolled, there is a substantial propensity for enrollees to continue in the same plan, equivalent to a price difference of about $84.

We explore the possibility of endogenous prices, given some evidence that firms do indeed take advantage of the reluctance for prior enrollees to switch, with prices being negatively correlated from the initial year to the next. Although there is evidence of endogeneity, its impact on price sensitivity is likely small. More importantly in terms of magnitude, we observe substantial heterogeneity in price sensitivity behind average estimates. Specifically with regards to health-associated heterogeneity, new enrollees are much more likely to differ in price responsiveness, with sicker new enrollees being less sensitive, and therefore are more prone to adverse selection. However, there is also likely a significant amount of heterogeneity that we cannot condition on observable characteristics, although it is sometimes revealed through prior choices (such as the choice to enroll in a high-premium, low-copayment plan).

Taken together, these results suggest that defined contribution plans may provide a strong incentive for firms to lower costs. In addition, if significant heterogeneity reveals true differences in preferences, then providing the option to choose among plans represents an additional welfare gain. However, because healthy new enrollees are substantially more price-sensitive than sicker ones, we may see adverse selection among new enrollees that may become entrenched once these consumers enroll and are likely to stay enrolled in the same plan over time. Price competition following these dynamics may be limited, and insurance providers who gain an early share of the healthy population may be able to extract rents from the reluctance of enrollees to switch plans by raising prices.

This analysis may be extended in several ways. First, we might explicitly model firm behavior and structurally test whether firms are following pricing behavior that exploits consumer inertia and considers the possibility of adverse selection among new enrollees. Second, as more data become available from the Commonwealth Connector Authority, we may evaluate the dynamics of insurance choices in the resulting panel data. In addition, such a panel data structure would provide a richer set of observations upon which to characterize heterogeneity in preferences by random coefficients, in a mixed logit model or perhaps even more non-parametrically (Berry and Haile 2009; Chernozhukov, Fernández-Val et al. 2009; Hoderlein and White 2009). Finally, we might attempt to evaluate the welfare effects of introducing a defined contribution exchange, as opposed to a single-choice arrangement, for low-income enrollees. Such an approach would consider the benefits of cost reduction and providing choice to heterogeneous consumers, weighed against the risk of adverse selection (Einav, Finkelstein et al. 2007; Chetty 2008; Einav, Finkelstein et al. 2008).

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APPENDIX: TABLES

Table 1: Changes in Enrollee Contribution by Plan Type

Plan Type	Change	2007 Range	2008 Range	Correlation
IIA	$9.14 ($13.78)	$0 ($0)	$24.15 ($10.04)	N/A
IIB	$14.04 ($25.63)	$40.19 ($14.49)	$47.62 ($20.48)	0.624
IIIA	$48.66 ($39.95)	$47.72 ($24.11)	$56.99 ($29.89)	-0.554
IIIB	$50.85 ($40.46)	$48.96 ($33.09)	$57.1 ($29.79)	-0.554
IVA	$16.08 ($42.15)	$38.74 ($20.8)	$57.92 ($30.48)	-0.530
IVB	$21.47 ($39.27)	$33.65 ($16.56)	$55.49 ($30.15)	-0.530
Total	$18.83 ($30.22)	$24.52 ($26.07)	$39.93 ($24.36)	

Enrollee contributions changed from “2007” to “2008” in July 2008. Numbers in the first column represent the means and standard deviations of the change in enrollee contribution for each plan averaged across areas, plan type, and insurer. The next two columns represent the difference in cost between the most expensive and cheapest plans for each area and plan type in 2007 and 2008. The last column presents the correlation between 2007 and 2008 cost. Plan types are as follows: IIA for those with incomes 100-150% of poverty, IIB for those 150-200% of poverty, IIIA or IVA for those 200-250% of poverty, and IIIB or IVB those 250-300% of poverty in plan type IIIB. Plan types IVA and IVB corresponded to lower-copayment, higher-premium options that were discontinued in 2008. To calculate the change in cost and 2008 cost range for IVA and IVB, we assume that enrollees continued in the corresponding plan by the same insurer and in the same area in IIIA and IIIB, respectively. Plan type IIA was free in 2007.

Table 2: Switches in Existing Enrollment from June 2008 to September 2008
By Plan Type and Medicaid Managed Care Organization (MMCO)

	June 2008 MMCO					September 2008 MMCO				
Plan Type	BMC	Fallon	NHP	Network	Total	BMC	Fallon	NHP	Network	Total
IIA	10,087	1,277	5,710	9,597	26,671	11,225	1,181	4,853	9,412	26,671
IIB	10,283	486	2,127	10,854	23,750	11,831	486	2,249	9,184	23,750
IIIA	3,956	420	787	3,749	8,912	3,354	1879	2966	2,813	11,012
IIIB	1,811	236	407	1,722	4,176	1,853	892	1312	1,467	5,524
IVA	986	130	368	616	2100					
IVB	657	89	179	423	1348					
Total	27,780	2,638	9,578	26,961	66,957	28,263	4,438	11,380	22,876	66,957

Numbers represent enrollees by June 2008 and the same enrollees who may have switched plans by September 2008 ("prior enrollees"), excluding those with one plan in their choice set and auto-enrollees. Enrollees with incomes 100-150% of poverty were in plan type IIA, those 150-200% of poverty in plan type IIB, those 200-250% of poverty in plan type IIIA (lower premium but higher copayment) or plan type IVA (higher premium but lower copayment), those 250-300% of poverty in plan type IIIB (lower premium but higher copayment) or plan type IVB (higher premium but lower copayment). Enrollee contributions changed from "2007" to "2008" in July 2008. Plan types IVA and IVB no longer were available in 2008. MMCOs include Boston Medical Center HealthNet Plan (BMC), Fallon Community Health Plan (Fallon), Neighborhood Health Plan (NHP), and Network Health.

Table 3: New Enrollment in 2008
By Plan Type and Medicaid Managed Care Organization (MMCO)

	MMCO				
Plan Type	BMC	Fallon	NHP	Network	Total
IIA	7,630	892	3,066	5,201	16,789
IIB	8,454	433	1,518	7,873	18,278
IIIA	3,422	983	1,738	3,161	9,304
IIIB	1,812	503	926	1,522	4,763
Total	21,318	2,811	7,248	17,757	49,134

Numbers represent numbers of new enrollees in each plan type for each MMCO in 2008, excluding those with one plan in their choice set and autoenrollees. MMCOs include Boston Medical Center HealthNet Plan (BMC), Fallon Community Health Plan (Fallon), Neighborhood Health Plan (NHP), and Network Health. Plan types are as defined above: IIA for those with incomes 100-150% of poverty, IIB for 150-200% of poverty, IIIA (low premium, high copayment) and IVA (high premium, low copayment) for 200-250% of poverty, and IIIB (low premium, high copayment) and IVB (high premium, low copayment) for 250-300% of poverty.

Table 4: Baseline Regressions

Model	Prior Enrollees (1)	New Enrollees (2)
Coefficients		
Same plan as in previous year	3.635*	
	(0.073)	
Price	-0.044*	-0.027*
	(0.002)	(0.002)
Fallon	-0.310*	-0.793*
	(0.129)	(0.223)
NHP	0.154	-0.565*
	(0.125)	(0.178)
Network Health	-0.084	-0.012
	(0.098)	(0.199)
Implied price responses		
Price elasticity	-0.667	-0.718
	(1.032)	(0.656)
Percent response to $10 price increase	-8.7%	-15.4%
	(7.7%)	(2.5%)

*Significantly different from 0 at the 5% level

This table reports results from baseline regressions of plan choice on price, plan dummies (Fallon, Neighborhood Health Plan, and Network Health, with BMC as the referent), and for prior enrollees, a dummy for whether the enrollee was already previously enrolled in the choice. Results for prior enrollees (model 1) are shown in the first column, while those for new enrollees (model 2) are shown in the second column. The top panel reports coefficients (standard errors in parentheses) of these regressions, while the bottom panel reports implied average price responses from these regressions weighted over the population and across choices by the likelihood of each enrollee to pick a choice; numbers in parentheses are standard deviations of the entire distribution of responses.

Table 5: Prior Enrollee Regressions with Control Variables for Endogeneity

Model	(1)	(3)	(4)	(5)
Coefficients				
Same plan as in previous year	3.635*	3.574*	3.801*	4.369*
	(0.073)	(0.085)	(0.076)	(0.127)
Price (same year)	-0.044*	-0.042*	-0.044*	-0.056*
	(0.002)	(0.002)	(0.002)	(0.005)
Previous year price			0.026*	-0.021
			(0.006)	(0.051)
Other plan type price**				0.018*
				(0.005)
Fallon	-0.310*		-0.671*	3.007
	(0.129)		(0.152)	(3.167)
NHP	0.154		-0.615*	2.563
	(0.125)		(0.167)	(2.144)
Network Health	-0.084		-0.114	-0.133
	(0.098)		(0.094)	(0.206)
Difference in price coefficient				
Hausman test statistic	Base	78.12	-0.03	-2.84
Hausman test p-value		<0.001	N/A	N/A
Bootstrapped p-value			0.38	0.44

*Significantly different from 0 at the 5% level

This table reports results from regressions on prior enrollee choices, using control variables for endogeneity. The first column (model 1) restates baseline results from Table 4. The second column (model 3) reports results of a regression without statewide plan dummies and reveals that these plan dummies capture statewide omitted variables that affect the price coefficient. The third column (model 4) adds previous-year price to the baseline regression; although it has a positive and significant coefficient, it does not affect the main price coefficient. The last column (model 5) reports results for the subgroup of enrollees in plan type IIA (those 100-150% of poverty) and includes both previous-year price and the current price of plans IIIA/B in the same area. The top panel reports coefficients (standard errors in parentheses) of these regressions, while the bottom panel reports tests on the difference between the coefficients on price in the baseline model (model 1) and the alternative model with control variables for endogeneity. Hausman test statistics and p-values are first reported; if the Hausman test statistics are negative in this finite sample, then bootstrapped p-values are also reported. Taken together, these results imply that other than statewide plan dummies, previous-year price and the price of other plan types do not capture much endogeneity but may reflect variables that enter into enrollee decision making as "rules of thumb."

Table 6: New Enrollee Regressions with Control Variables for Endogeneity

	Enrollees after June 2008		**Enrollees before June 2008**	
Model	(2)	(6)	(7)	(8)
Coefficients				
Price (same year)	-0.024*	-0.024*	-0.032*	-0.031*
	(0.002)	(0.002)	(0.002)	(0.002)
Previous year price		0.011*		
		(0.003)		
Future year price				0.005*
				(0.002)
Other plan type price				
Fallon	-0.821*	-0.982*	-0.679*	-0.666*
	(0.166)	(0.191)	(0.287)	(0.298)
NHP	-0.655*	-0.966*	-0.425*	-0.522*
	(0.176)	(0.213)	(0.199)	(0.206)
Network Health	-0.246	-0.123	0.086	0.057
	(0.200)	(0.201)	(0.208)	(0.199)
Difference in price coefficient				
Hausman test statistic	Base	0.21	Base	5.99
Hausman test p-value		0.64		0.01

*Significantly different from 0 at the 5% level

This table reports results of regressions of new enrollee choices that include control variables for endogeneity. Current prices were updated for the fiscal year in June 2008, so that new enrollees enrolling after June 2008 faced "2008" prices, while those enrolling before June 2008 faced "2007" prices. We perform regressions on the subgroup of those enrolling after June 2008 (models 2 and 6) to see if including "2007" prices affects decisions; and we perform similar regressions on those enrolling before June 2008 to see if "2008" prices affect decisions. Coefficients (standard errors in parentheses) are presented in the top panel; Hausman test results are presented in the bottom two rows. Although "2007" prices affect the decisions of those enrolling after they are relevant, they do not change the coefficient on current ("2008") prices, suggesting at least some use of behavioral "rules of thumb" in decision-making. For those enrolling before June 2008, "future" ("2008") prices are obviously unknown but still are significantly positive and change the coefficient on current prices, suggesting endogeneity.

Table 7: Control Function Approach to Endogeneity

Model	Prior Enrollees (9)	Prior Enrollees (10)	New Enrollees (11)
Coefficients			
Same plan as in previous year	3.628* (0.079)	3.793* (0.083)	
Price (same year)	-0.046* (0.002)	-0.045* (0.002)	-0.029* (0.002)
Previous year price		0.025* (0.006)	
Fallon	0.150 (0.234)	-0.318 (0.223)	0.071 (0.344)
NHP	0.368* (0.150)	-0.446* (0.162)	-0.233 (0.242)
Network Health	-0.252* (0.101)	-0.244* (0.100)	-0.168 (0.192)
Residual from regressing price on same-plan prices in other regions (control function)	0.020* (0.008)	0.015* (0.007)	0.029* (0.010)
Implied price responses			
Price elasticity	-0.711 (1.094)	-0.706 (1.102)	-0.753 (0.679)
Percent response to $10 price increase	-9.1% (8.3%)	-8.9% (8.3%)	-16.2% (2.5%)

*Significantly different from 0 at the 5% level

This table presents results from regressions that use Hausman-type instruments (prices of the same MMCO from other regions) to form residuals ("control functions") that then enter into the main regression. The first two columns present results for prior enrollees, with the first column (model 9) adding the control function to the baseline model (model 1), and the second column (model 10) also including previous-year price, which is still significantly positive but not affecting the current price coefficient (Hausman test not shown). The last column (model 11) adds the control function to the baseline model for new enrollees (model 2). Coefficients (standard errors in parentheses) are presented in the first panel; the last two rows present implied price responses as elasticities and percentage responses to a $10 price increase, again weighted over the population and across choices by the likelihood of each enrollee to pick a choice; numbers in parentheses are standard deviations of the entire distribution of responses.

Table 8: Regressions Accounting for Heterogeneity

Independent Variable	Prior Enrollees (12)	(13)	(14)	(15)	New Enrollees (16)	(17)
Same plan as in previous year	4.007*	4.006*	3.998*	3.998*		
	(0.078)	(0.078)	(0.086)	(0.086)		
Price (same year)	-0.078*	-0.080*	-0.088*	-0.090*	-0.044*	-0.053*
	(0.006)	(0.006)	(0.009)	(0.009)	(0.005)	(0.004)
Previous year price	0.021*	0.021*	0.019*	0.018*		
	(0.004)	(0.004)	(0.003)	(0.003)		
Price × Q2	0.003*	0.001	0.002*	0.001	0.009*	0.004*
	(0.001)	(0.001)	(0.001)	(0.001)	(0.003)	(0.001)
Price × Q3	0.001	0.005*	0.001	0.005*	0.009	0.012*
	(0.001)	(0.001)	(0.001)	(0.001)	(0.005)	(0.002)
Price × Q4	0.003*	0.009*	0.003*	0.009*	0.009	0.012*
	(0.001)	(0.001)	(0.001)	(0.001)	(0.005)	(0.001)
Price × IIB	0.024*	0.024*	0.029*	0.029*	0.004*	0.016*
	(0.004)	(0.004)	(0.005)	(0.005)	(0.001)	(0.003)
Price × IIIA	0.034*	0.034*	0.042*	0.042*	0.012*	0.018*
	(0.009)	(0.009)	(0.012)	(0.012)	(0.002)	(0.005)
Price × IIIB	0.038*	0.038*	0.046*	0.046*	0.013*	0.019*
	(0.009)	(0.009)	(0.012)	(0.012)	(0.001)	(0.005)
Price × IVA	0.050*	0.048*	0.058*	0.056*		
	(0.008)	(0.008)	(0.011)	(0.011)		
Price × IVB	0.050*	0.049*	0.058*	0.057*		
	(0.009)	(0.009)	(0.012)	(0.012)		
Fallon	0.126	0.110	0.925	0.908	-0.764*	0.212
	(0.292)	(0.291)	(0.549)	(0.549)	(0.242)	(0.365)
NHP	0.094	0.088	0.538	0.531	-0.541*	-0.153
	(0.205)	(0.205)	(0.320)	(0.319)	(0.179)	(0.254)
Network Health	0.184	0.182	-0.005	-0.006	-0.005	-0.164
	(0.143)	(0.141)	(0.129)	(0.128)	(0.202)	(0.194)
Control function			0.027*	0.026*		0.031*
			(0.011)	(0.011)		(0.010)

*Significantly different from 0 at the 5% level

This table presents coefficients (standard errors in parentheses) of model that account for heterogeneity in the price coefficient by adding interactions with health quartiles ("Q1" for the healthiest quartile; "Q4" for the sickest) and income groups (defined in the text and in the captions for Tables 1-3). The first four columns (models 12 to 15) are for prior enrollees, while the last two (models 16 and 17) are for new enrollees. Models 14, 15, and 17 include a control function for endogeneity (described in the text and in the caption for Table 7); models 12, 13, and 16 are similar but do not include the control function. For prior enrollees, we consider two measures health – demographic-based (models 12 and 14) and health-utilization-based (models 13 and 15); for new enrollees we only have demographic-based measures available. See the text (pages 14-16) for interpretation of these results.

Table 9: Average Price Elasticities Accounting for Heterogeneity

	Prior Enrollees				New Enrollees	
Subgroup	(12)	(13)	(14)	(15)	(16)	(17)
Total	-0.671	-0.671	-0.712	-0.711	-0.714	-0.747
	(0.979)	(0.987)	(1.018)	(1.025)	(0.653)	(0.651)
Q1	-0.650	-0.763	-0.687	-0.803	-0.666	-0.685
	(0.998)	(1.112)	(1.036)	(1.151)	(0.723)	(0.713)
Q2	-0.673	-0.752	-0.715	-0.794	-0.804	-0.827
	(0.970)	(1.073)	(1.009)	(1.113)	(0.717)	(0.705)
Q3	-0.721	-0.644	-0.765	-0.686	-0.666	-0.705
	(1.007)	(0.938)	(1.047)	(0.978)	(0.571)	(0.577)
Q4	-0.645	-0.530	-0.686	-0.567	-0.679	-0.724
	(0.924)	(0.779)	(0.963)	(0.814)	(0.584)	(0.596)
IIA	-0.141	-0.141	-0.158	-0.157	-0.055	-0.070
	(0.202)	(0.202)	(0.224)	(0.224)	(0.104)	(0.130)
IIB	-0.546	-0.540	-0.566	-0.560	-0.688	-0.746
	(0.819)	(0.815)	(0.851)	(0.847)	(0.142)	(0.144)
IIIA	-1.833	-1.836	-1.939	-1.940	-1.366	-1.383
	(1.045)	(1.064)	(1.048)	(1.066)	(0.389)	(0.361)
IIIB	-2.112	-2.126	-2.247	-2.260	-1.867	-1.891
	(1.297)	(1.331)	(1.306)	(1.339)	(0.545)	(0.507)
IVA	-0.783	-0.791	-0.860	-0.867		
	(0.614)	(0.644)	(0.630)	(0.661)		
IVB	-1.059	-1.057	-1.152	-1.148		
	(0.763)	(0.819)	(0.777)	(0.834)		

This table presents own-price elasticities averaged over the relevant population and weighted by the likelihood of each enrollee to pick a given plan. Standard deviations of the entire distribution of elasticities are given in parentheses. Columns correspond exactly to those in Table 8; with the first four for prior enrollees and last two for new enrollees (see caption in Table 8 for detailed description). Rows represent the relevant population, with “Total” denoting the overall population, “Q1” to “Q4” denoting subgroups selected by quartiles of health (from healthiest to sickest), and “IIA” to “IVB” denoting plan types according to income (see captions for Tables 1-3 for detailed description of plan types). We find greater changes in price sensitivity according to quartiles of health among new enrollees than among prior enrollees. Although plan types IIIA and IVA, and plan types IIIB and IVB, correspond to identical income brackets, those who picked the high-premium, low-copayment plan types IVA and IVB show much greater price sensitivity. Health based on health-utilization measures (models 13 and 15) are better at predicting changes in price-sensitivity than using health based on demographics (models 12 and 14). Finally, accounting for endogeneity with a control function (models 14, 15, and 17) increases price sensitivity slightly.

Table 10: Percentage Response to $10 Price Increase Accounting for Heterogeneity

	Prior Enrollees				New Enrollees	
Subgroup	(12)	(13)	(14)	(15)	(16)	(17)
Total	-10.5%	-10.5%	-11.4%	-11.4%	-17.2%	-19.4%
	(9.8%)	(9.9%)	(10.6%)	(10.6%)	(5.0%)	(6.5%)
Q1	-11.2%	-11.9%	-12.2%	-12.8%	-21.7%	-24.1%
	(10.5%)	(11.0%)	(11.3%)	(11.8%)	(5.0%)	(6.7%)
Q2	-10.3%	-11.5%	-11.1%	-12.4%	-19.0%	-20.9%
	(9.6%)	(10.7%)	(10.3%)	(11.4%)	(4.5%)	(6.1%)
Q3	-10.7%	-10.1%	-11.5%	-11.0%	-15.0%	-17.1%
	(9.8%)	(9.4%)	(10.5%)	(10.1%)	(3.7%)	(5.4%)
Q4	-9.7%	-8.6%	-10.5%	-9.4%	-14.5%	-16.8%
	(9.1%)	(8.0%)	(9.8%)	(8.7%)	(3.6%)	(5.4%)
IIA	-11.0%	-11.0%	-12.3%	-12.3%	-21.9%	-26.9%
	(10.3%)	(10.3%)	(11.4%)	(11.5%)	(3.9%)	(4.0%)
IIB	-7.8%	-7.8%	-8.2%	-8.1%	-14.7%	-15.9%
	(9.6%)	(9.5%)	(10.0%)	(10.0%)	(3.4%)	(3.2%)
IIIA	-16.4%	-16.4%	-17.3%	-17.3%	-15.1%	-15.2%
	(8.3%)	(8.5%)	(8.3%)	(8.5%)	(3.6%)	(3.3%)
IIIB	-13.8%	-13.9%	-14.7%	-14.7%	-14.5%	-14.6%
	(7.7%)	(7.9%)	(7.7%)	(8.0%)	(3.7%)	(3.4%)
IVA	-6.8%	-6.9%	-7.5%	-7.6%		
	(4.6%)	(5.0%)	(4.8%)	(5.1%)		
IVB	-6.8%	-6.8%	-7.5%	-7.4%		
	(4.4%)	(4.8%)	(4.5%)	(4.9%)		

This table presents percentage responses to a $10 increase in price averaged over the relevant population and weighted by the likelihood of each enrollee to pick a given plan. Standard deviations of the entire distribution of percentage responses are given in parentheses. Columns correspond exactly to those in Table 8; with the first four for prior enrollees and last two for new enrollees (see caption in Table 8 for detailed description). Rows represent the relevant population, with "Total" denoting the overall population, "Q1" to "Q4" denoting subgroups selected by quartiles of health (from healthiest to sickest), and "IIA" to "IVB" denoting plan types according to income (see captions for Tables 1-3 for detailed description of plan types). We find greater changes in price sensitivity according to quartiles of health among new enrollees than among prior enrollees. Although plan types IIIA and IVA, and plan types IIIB and IVB, correspond to identical income brackets, those who picked the high-premium, low-copayment plan types IVA and IVB show much greater price sensitivity. Health based on health-utilization measures (models 13 and 15) are better at predicting changes in price-sensitivity than using health based on demographics (models 12 and 14). Finally, accounting for endogeneity with a control function (models 14, 15, and 17) increases price sensitivity slightly.

Table 11: Mixed Logit Regressions

	Price Coefficient Conditioning			
	None	**Income**	**Health**	**Both**
Model	(18)	(19)	(20)	(21)
Coefficient Mean				
Same plan as in previous year**	1.667*	1.635*	1.672*	1.636*
	(0.059)	(0.055)	(0.059)	(0.055)
-Price (same year)**	-2.983*	-2.658*	-2.822*	-2.499*
	(0.077)	(0.164)	(0.089)	(0.133)
Fallon†	0.141	0.636	0.120	0.732*
	(0.238)	(0.357)	(0.234)	(0.357)
NHP†	0.449*	0.706*	0.407*	0.745*
	(0.151)	(0.202)	(0.144)	(0.202)
Network Health†	-0.189	-0.135	-0.198	-0.133
	(0.112)	(0.113)	(0.112)	(0.113)
Control function	0.009	0.018*	0.010	0.021*
	(0.007)	(0.008)	(0.007)	(0.008)
Coefficient Standard Deviation				
Same plan as in previous year**	0.047	0.036	0.049	0.041
	(0.097)	(0.087)	(0.086)	(0.081)
-Price (same year)**	1.013*	0.691*	0.865*	0.570*
	(0.094)	(0.149)	(0.082)	(0.111)
Fallon†	-0.110	-0.130	-0.119	-0.126
	(0.388)	(0.390)	(0.386)	(0.388)
NHP†	-1.560*	-1.541*	-1.616*	-1.566*
	(0.208)	(0.189)	(0.206)	(0.188)
Network Health†	-1.198*	-1.015*	-1.235*	-1.035*
	(0.258)	(0.276)	(0.254)	(0.272)
Implied Moments of Price Coefficient				
Mean	-0.0846	-0.0890	-0.0864	-0.0966
Variance	0.0127	0.0048	0.0083	0.0036
Percent of unconditional variance	Base	38.0%	65.1%	28.2%

*Significantly different from 0 at the 5% level

**Random coefficient distributed log-normally (parameters presented for log coefficient)

†Random coefficient distributed normally

This table presents mixed logit models that quantify the variation in the coefficients by fitting entire distributions of coefficients for random-coefficient variables: we model coefficients on plan dummies as normally distributed and coefficients on the same-plan dummy and negative price as log-normally distributed. Columns correspond whether the price coefficient is also conditioned on observables such as income or health with price interactions with fixed coefficients; the first column represents unconditional variance, while the next three columns condition on income, health, or both, respectively. Distribution means and standard deviations for random coefficients are in the first and second panels; the implied mean and variance of the log-normally distributed price coefficient are in the last panel.