

**Can Risk Adjustment Reduce Selection
in the Private Health Insurance Market?
New Evidence from the Medicare Advantage Program**

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Abstract

Roughly 25 percent of Medicare beneficiaries are enrolled in a private Medicare Advantage plan, a fraction that has been growing steadily in recent years. Because these plans are “at risk” for the cost of their enrollees’ care, they have an incentive to attract and retain differentially healthy enrollees. To address this problem, starting in 2004, measures of enrollee health status and health care utilization were used to determine payments for MA plans. Using both individual and aggregate county-level data, we find no evidence that this risk adjustment reduced positive selection into MA plans. Our estimates suggest that Medicare recipients who join MA plans would have cost approximately \$1,200 less per year if they had remained in FFS than the average FFS beneficiary. As on average MA plans are paid substantially more per enrollee than per capita FFS spending, we estimate that a Medicare beneficiary choosing an MA plan over traditional Medicare increases total Medicare costs by \$2,500. Finally, we propose and provide evidence for a mechanism for our findings: MA plans offering care and cost sharing arrangements that differentially attract and retain healthy patients.

I. Introduction

Since Medicare's inception in 1966, the vast majority of the program's recipients have been enrolled in traditional, fee-for-service (FFS) Medicare, through which the federal government directly reimburses hospitals, physicians, and other health care providers. While there are some advantages to this payment model, one persistent concern has been its effect on the incentives of health care providers. Because it bases payments on the quantity rather than the quality of services, critics of the FFS payment model argue that it encourages providers to deliver services with little clinical benefit.

Partly because of this concern, since the early 1980s the government has contracted with private health insurers to coordinate care for Medicare recipients. The Medicare program pays these Medicare Advantage¹ (MA) plans a fixed amount per month for each enrollee, which gives them a financial incentive to reduce the utilization of low-value services. Throughout the 1980s and 1990s, this capitation payment was typically set equal to 95 percent of average per capita FFS costs in the recipient's county. The capitation payment was further adjusted to account for the recipient's age, gender, and a few other characteristics, but did not adjust for health status.²

Because private insurers were at risk for the cost of their enrollees' care, they had an incentive to provide high-quality care that would keep their patients healthy and using less medical care. However, insurers also had an incentive to design their package of benefits and market their plans to attract and retain Medicare recipients whose expected costs were lower than their demographic characteristics would suggest. Studies found that, because MA enrollees' actual medical costs were well below 95 percent of the FFS average, capitation payments exceeded actual costs and thus MA enrollment actually increased total Medicare spending.(Langwell and Hadley, 1989; PPRC, 1997; Batata, 2004). Partly because of these findings, a comprehensive risk-adjustment model was introduced in 2004. As risk-adjustment

¹ The term Medicare Advantage (MA) was first used for private plans in 2004. From 1997 to 2003, plans were called Medicare + Choice and prior to that were simply referred to as Medicare HMOs. For simplicity, we use the term Medicare Advantage when describing Medicare private plans throughout this paper.

² Over time, the formula has been altered to achieve certain policy goals, such as increasing rural access to private plans, reducing perceived overpayments to certain areas, or trying to stop plan exit from local markets, but aside from the geographic component, the basic structure of the capitation payments remained essentially unchanged until recently (Berenson and Dowd, 2003).

should better align capitation payments and actual medical costs had the enrollee remained in FFS, it should, all else equal, decrease total Medicare spending. Moreover, one would expect the shift to risk-adjusted payments to affect the pattern of selection into MA plans, which would have a stronger incentive (or a weaker disincentive) to attract and retain individuals in worse health.

In our first set of analyses, we use longitudinal data for a sample of nearly fifty thousand Medicare recipients from the 1999 through 2006 annual Medicare Current Beneficiary Survey (MCBS) to estimate the effect of MA enrollment on Medicare spending. When doing this, we account for the non-random selection of Medicare recipients into MA plans by controlling for a rich set of demographic characteristics as well as flexibly controlling for Medicare expenditures in previous years.

Our findings indicate that on average, when a Medicare recipient switches from FFS to MA, Medicare expenditures increase by approximately \$2,500 (or 30 percent) relative to expenditures had the recipient remained in FFS. Approximately half of this effect is attributable to the fact that the average capitation payment is substantially higher than average FFS spending during the sample period. The other half is due to low-cost Medicare recipients differentially enrolling in MA plans, and we find no evidence that this selection declined following the shift to risk-adjusted payments. While our estimates for those switching from FFS to MA may not generalize to the entire MA population, additional analyses suggest that while these overpayments may decline over time, they remain substantial even upon disenrollment.

One limitation with our results is that it ignores the possibility of spillovers to the Medicare FFS population. To the extent that increases in MA enrollment influence practice patterns and thus expenditures for those remaining in Medicare FFS, our individual-level results will not capture this. To explore this possibility, in our second set of analyses we use aggregate county-level data to explore whether changes in MA penetration are significantly related with changes in the average cost of FFS recipients, as one would expect if favorable selection occurred (Batata, 2004). We use annual data from 2000 to 2008, thus spanning a period of substantial MA decline and subsequent increases, and also explore whether this relationship changed over time.

Our findings in our county-level analyses indicate that the marginal cost of MA enrollees / disenrollees is approximately \$1,200 per year less than the average cost of FFS stayers. This estimate matches almost exactly the corresponding one from the MCBS individual-level analysis, suggesting that contemporaneous spillovers from the MA population to the Medicare FFS population are, if they exist, relatively small, and certainly not sufficient to offset the effect for MA enrollees. Furthermore, we once again find little evidence to suggest that the extent of selection, as measured by Medicare expenditures, changed with the move to greater risk adjustment.

Of course, expenditures represent just one measure of the health of Medicare recipients. To further investigate whether MA plans became more likely to enroll individuals in poor health, we utilize individual-level data from the MCBS on self-reported health status and on the presence of chronic conditions. Consistent with our analyses for expenditures, we find little evidence that those in poor health or with chronic conditions (such as previous heart attacks) became more likely to enroll in MA.

In our fourth and final set of empirical analyses, we investigate why MA plans differentially attract low-cost individuals throughout our study period. Using various measures of satisfaction with care, we find that MA enrollees in poor health are significantly less satisfied with their care and out-of-pocket expenditures than their FFS counterparts in poor health. The difference is much smaller, and in many cases reversed, between MA and FFS recipients in excellent health. It therefore appears that MA plans appeal differentially to individuals in excellent health.

Taken together, our results suggest that overpayments to MA plans are substantially greater than is commonly believed and that the move to risk-adjusted payments has not reduced the amount of favorable selection into MA plans.³ Our results take on additional significance when one considers that, as shown in Figure 1, MA enrollment has more than doubled during the last five years, with almost one-in-four Medicare recipients now enrolled in an MA plan, and Medicare payments to these plans exceeding \$100 billion in 2009.

³ As our focus is on the effect on government expenditures, we do not estimate the effect of MA enrollment on total producer and consumer surplus. See Town and Liu (2003) for an examination of these and related issues.

[Figure 1]

The outline of the paper is as follows. In section two we provide a brief background on Medicare Advantage (and its predecessors) and summarize the shift to risk-adjusted plan payments that occurred during our study period. In section three we summarize our MCBS data and the county-level data used in our subsequent empirical analyses. Section four presents our individual-level analyses for the effect of MA enrollment on Medicare expenditures. In section five we use county level data to extend our study period through 2008 and to explore whether there are spillovers to the FFS population that attenuate or amplify our individual-level estimates. In section six, we use various measures of health status to explore whether individuals in worse health or with certain chronic conditions became more likely to enroll in MA plans over time. Section seven summarizes our analyses of the extent to which MA plans differentially appeal to healthy individuals, and section eight concludes and discusses some of the implications of our results and directions for future research.

II. Background on Medicare Advantage and Risk-Adjustment

Since Medicare's inception in 1966, the vast majority of program recipients have been enrolled in traditional fee-for-service Medicare. Under this model, the federal government directly reimburses physicians, hospitals, and other health care providers for services provided to program recipients.

Beginning in the 1980s, Medicare recipients were given the option of enrolling in private plans, which eventually became known as Medicare Advantage (MA) plans, and today nearly a quarter of Medicare recipients are insured through MA plans (see Figure 1). Unlike in the private non-group market for the non-elderly, private MA plans must offer plans on a guaranteed-issue, community-rated basis to any Medicare recipient in their geographic area of operation.

Instead of reimbursing private plans for the medical costs an individual incurred, Medicare instead issues private MA plans a fixed capitation payment meant to approximately cover an individual's costs. The private plans are the residual claimant on the surplus or loss when an individual's actual medical costs fall below or above, respectively, the capitation payment. Proponents argued that MA plans

would thus avoid needless, costly procedures that they believed the FFS reimbursement model incentivized and would ultimately result in substantial savings to the Medicare program. Of course, if MA capitation payments are set above what an individual would have cost the FFS program, then potential savings from outsourcing to private plans will be diminished.

Risk-adjustment of capitation payments

Throughout the 1980s and 1990s, capitation payments were adjusted via a “demographic model,” so-called because it only adjusted for the demographic factors (age, gender, and whether an individual was on Medicaid, Disability or institutionalized) as opposed to underlying medical conditions. For example, MA plans would receive a smaller capitation payment for enrolling a 65-year-old than an 85-year-old, and for someone living on their own in the community than someone in a nursing home. The model was based on cost data from the FFS population and predicted only one percent of the variation in medical costs among the FFS population (Pope et al., 2004). Given that out-of-sample prediction is usually more difficult, it is unlikely that the model explained any more of the variation among the MA population, the group for which it was actually used to base capitation payments.

As such, there was substantial scope for MA plans to increase profits by attracting enrollees who were healthier—and thus cheaper—than the demographic model would predict. Indeed, a large body of literature has shown that during this period MA plans were able to attract patients who were far less costly than the FFS population in general or than the prediction of the demographic model. Estimates suggest that individuals switching from traditional FFS to MA had medical costs between 20 and 37 percent lower than individuals who remained in FFS.⁴ However, as MA capitation payments were on average 95 percent of average FFS per capita spending during this period, MA plans were essentially being compensated for insuring the average Medicare recipient, when they were typically insuring recipients far healthier—and cheaper—than average.

⁴ See, for example, Langwell and Hadley (1989), PPRC (1997), Mello (2003), and Batata (2004).

Federal policymakers reacted to this evidence by enhancing the risk-adjustment procedure. In 2000, ten percent of the payment was based on a model of inpatient diagnoses that explained 6.2 percent of cost variation, with the remaining ninety percent on the demographic model, thereby increasing the explanatory power of the final risk score from one percent to 1.52 percent ($1 \cdot .90 + 6.2 \cdot .10$). A more comprehensive risk-adjustment regime was introduced in 2004 based on the hierarchical condition categories (HCC) model. The model uses spending and utilization data from the FFS population to predict spending, based on demographic characteristics and disease history for the previous year. Several thousand diagnosis codes are condensed into 70 disease categories, plus some interactions, for which a factor is assigned to increase the baseline capitation rate. For instance, HIV/AIDS for a community dweller is assigned a factor of 0.945. With fully phased-in risk adjustment, a plan enrolling someone with HIV/AIDS would receive an additional 94.5 percent of the benchmark rate compared to an otherwise identical enrollee without the condition. This comprehensive model explains 11.2 percent of cost variation among the FFS population.

As the share of the capitation payment based on the HCC model rose from 30 percent in 2004 percent to 50, 75 and 100 percent in 2005, 2006 and 2007 respectively, the actual explanatory power of the final risk score rose from 1.52 percent in 2003 to 4.06 in 2004, 6.10 in 2005, 8.65 in 2006, and finally 11.2 percent in 2007. Of course, no methodology could perfectly forecast health care expenditures given the inherent unpredictability of changes from one year to the next in any individual's health status and health care costs. See Pope et al. (2004) for a detailed description of the HCC model.

Performance of the new model

While there is a large literature examining both the extent of selection into MA plans and the effect of MA plans during the 1980s and 1990s (see Mello et al., 2003 for a review), there is virtually no work explaining these same issues during the most recent decade. In the pages that follow, we aim to fill this gap, but before doing so, it is worth discussing the likely challenges remaining for this or any similar risk-adjustment model. Given the growing share of Medicare recipients in MA plans and the growing

cost of insuring the Medicare population generally, improvements to the risk-adjustment of capitation payments could lead to substantial savings. Moreover, the new insurance exchanges established by the Patient Protection and Affordable Care Act will rely on risk adjustment to compensate private plans for the expected cost of insuring the non-elderly.

Perhaps the greatest challenge in the Medicare Advantage setting is the need to calibrate the model using the FFS population. This reliance arises because MA plans do not submit cost data: they receive their capitation payment and then are essentially left alone to coordinate patients' care in whatever manner they find most efficient, subject to certain regulations. Indeed, basing payments on plans' cost data would undermine the guiding principle of avoiding marginal-cost compensation so as to disincentive needless or low-value care. As discussed earlier, the relatively low predictive power of the HCC model would likely be even lower when taken out-of-sample to the actual MA population.

Second, favorable selection can obviously still occur even when detailed information on individuals' health status is used in calculating capitation payments. The severity of a disease can vary, and if private plans are able to recruit individuals with mild cases of a given disease or if these individuals simply self-select into private plans, then overpayments will persist.

Third, the manner in which a private plan classifies an individual's disease condition affects his payment, and, holding actual health constant, a plan that more zealously codes disease conditions will have larger capitation payments and greater profits. The Center for Medicare and Medicaid Services (CMS) has argued that individuals in MA plans are more likely to have a higher disease score than an individual in FFS, holding actual health constant. For example, approximately one-in-four U.S. residents with diabetes have not yet been diagnosed with the condition. If average spending for newly diagnosed individuals is lower than for those already diagnosed with a condition, then more "intensive coding" (e.g. through testing all enrollees for diabetes upon enrollment) could increase spending for MA enrollees relative to their otherwise identical counterparts in traditional Medicare.

The remainder of the paper investigates whether the shift to risk-adjusted capitation payments has influenced selection into MA plans or the average effect of the plans on Medicare expenditures.

III. Data and Summary Statistics

A. Individual Level Data from the Medicare Current Beneficiary Survey

Our empirical work relies on individual-level data on Medicare spending, self-reported health, and related outcome variables of interest from the Medicare Current Beneficiary Survey (MCBS), a nationally representative survey of Medicare beneficiaries.⁵ The MCBS is based both on CMS administrative data and on surveys from a sample of roughly 11,000 individuals each year. Important for our analysis, the MCBS follows a subsample of respondents for up to four years. We utilize MCBS data from the 1999 through 2006 period, which allows us to examine whether the impact of MA enrollment changed as CMS shifted to risk-adjusted capitation payments.

The MCBS contains information on each respondent's demographic characteristics, educational attainment, and place of residence along with data on total Medicare spending for the year, utilization of medical care, self-reported health, and satisfaction with care. The MCBS also uses CMS administrative enrollment data to determine each month whether respondents are enrolled in an MA plan. During our study period from 1999 to 2006, there were 48,360 unique Medicare recipients surveyed. Because many MCBS respondents are observed for more than one year, the total number of person-year observations was substantially higher at 90,618.

Table 1 provides summary statistics from the 90,618 person-years in the MCBS data, differentiating by coverage type (MA versus FFS) and time period (1999 to 2003, versus 2004 to 2006). We present summary statistics pooled across all years in our data in Appendix Table 1. An individual is classified as being on MA if she is enrolled in MA for half or more of her Medicare-enrolled months in a given calendar year. As the table shows, there are substantial differences between the FFS and MA populations in both periods. For example, Medicare recipients who are under the age of 65, virtually all of whom qualify for Medicare because they are disabled, are significantly more likely to be in traditional

⁵ To be included in the MCBS in a particular year, a person must be enrolled in Medicare in January of that year. Thus a person who enrolled midway through the year could not appear in that year's MCBS.

Medicare than in MA plans. The same is true for Medicare recipients age 85 and up. Women are similarly likely to be in FFS and MA, with blacks somewhat more likely and Hispanics substantially more likely to be MA-enrolled. Individuals residing in metro areas are significantly more likely to be enrolled in MA plans, while college graduates are significantly less likely.

[Table 1]

Thus in terms of their average observable characteristics, FFS and MA recipients differ in many respects during both the early and later years in our study period. The same is true with respect to self-reported health status and total Medicare spending. In both periods, MA recipients are significantly more likely to report that their health status is very good or excellent and significantly less likely to report that they are in poor or fair health. At the same time, however, average Medicare spending is in both periods higher for MA recipients.

B. Aggregate County Level Data

Our county-level analyses utilize administrative data from CMS on county-level per-capita Medicare FFS expenditures for elderly Medicare FFS recipients and on county-level MA penetration in each year from 2000 through 2008. We adjust all expenditure amounts to 2007 dollars using the CPI-U. Average county-level FFS expenditures for the elderly during our nine-year study period are approximately \$6,680 and the average change from one year to the next from 2000 to 2008 in this measure is \$233. Average MA enrollment during the same period is approximately 15.9 percent, with this declining from 17.4 percent to 12.6 percent from 2000 to 2004, and then increasing to 22.0 percent by 2008. One advantage of this pattern is that it allows us to explore whether our estimated relationships are similar when MA enrollment is declining and when it is increasing. As shown in Appendix Table 2, there is considerable variation across counties in each year with respect to the change in the fraction of Medicare recipients in MA plans.

The sources for this county-level data are listed in the Appendix. When constructing the sample, we excluded all counties in Alaska because of inconsistent county definitions across years. We also

excluded a small number of counties that were missing data in two or more years, leaving us with a final sample of 3,110 counties.

IV. The Impact of MA Enrollment on Medicare Expenditures: MCBS Analysis

Empirical Strategy

In this section, we use the individual-level MCBS data to estimate the impact of MA enrollment on Medicare expenditures. As summarized in the preceding section, MA enrollees differ in several observable ways from their counterparts in traditional Medicare. For example, MA recipients are less likely to be on Social Security Disability Insurance (and thus under age 65) and are more likely to reside in metropolitan areas. These same factors will affect average Medicare spending among MA recipients, and it is therefore important to control for them in specifications of the following type:

$$(1) S_{jt} = \alpha_t + \mu * MA_{jt} + \beta * X_{jt} + \rho * H_{jt} + \gamma * METRO_{jt} + \sum \lambda_k * I(State_{jt} = k) + \varepsilon_{jt}$$

In this equation, S_{jt} represents Medicare spending for individual j in year t and MA_{jt} equals the fraction of months in year t during which person j is enrolled in MA. X_{jt} and H_{jt} represent a set of controls for the person's demographic (e.g. age, gender, race) and health characteristics, which are likely to be strongly related to the individual's Medicare spending. $METRO_{jt}$ is an indicator variable that is equal to 1 if he/she resides in a metropolitan area and otherwise is equal to 0, with this variable included to account for the fact that health care spending tends to be higher in metro areas. The specification also includes 51 indicator variables for the person's state of residence, as much previous work has shown that average Medicare expenditures vary substantially across states.

To the extent that there are no systematic unobserved differences between those enrolled in MA plans and their observably similar counterparts in traditional Medicare, one can interpret the coefficient estimate μ as the average impact of MA enrollment on Medicare spending. However, for the reasons mentioned above, it is plausible that there are such differences.

To address this issue, one can exploit the longitudinal nature of the MCBS. Specifically, we examine how total Medicare expenditures changes for individuals who transition from FFS to MA, compared to individuals who remain in FFS. Consider the following specification, estimated on the population of individuals who were enrolled in FFS in year $t-1$.

$$(2) S_{jt} = \alpha_t + \mu * MA_{jt} + f(S_{j,t-1}) + \beta * X_{jt} + \rho * H_{jt} + \gamma * METRO_{jt} + \sum \lambda_k * I(\text{State}_{jt} = k) + \epsilon_{jt}$$

where $f(\cdot)$ is a flexible function and $S_{j,t-1}$ is baseline spending. The identifying assumption is that, conditional on the control variables present in the equation, an individual's decision to join MA is not systematically related to time-varying changes in her demand for Medicare expenditures (ϵ_{jt}). Put differently, this specification assumes that, conditional on our covariates, individuals who remain in FFS ("FFS stayers") serve as a valid control group for individuals who enroll in MA ("MA joiners"). Of course, it is straightforward to expand the sample to include individuals enrolled in MA in the baseline year, which allows one to compare the pattern of spending with MA "stayers" and MA "leavers" as well.

There are at least three limitations worth noting with this approach, all of which we explore below. First, because an individual's decision to enroll in MA plans is endogenous, we cannot with these specifications rule out the possibility that there are unobserved changes that might bias our estimates. For example, individuals may tend to enroll in MA plans just prior to having surgery to get access to the physicians in an MA plan's network. Second, this methodology estimates the effect of MA enrollment from individuals newly joining MA plans. If, for example, Medicare spending for new MA enrollees grows more slowly in subsequent years than it would if individuals remained in FFS, then our results would not generalize to the entire population. Third, this result does not account for the possibility of spillovers. To the extent that changes in MA enrollment influence Medicare spending for those who remain in traditional Medicare, this specification would not capture it.

Enrollment and Expenditure Patterns in the MCBS

Our interest in exploiting the longitudinal nature of the MCBS is motivated by Figure 2, which displays the average Medicare expenditures for individuals based on their MA status in years $t-1$ and t . Between the baseline year and the subsequent year, individuals who remained in FFS had their Medicare expenditures increase by an average of \$1,232.⁶ Individuals who joined MA, in contrast, experienced average growth in expenditures of \$4,165, or nearly \$3,000 more. At the same time, individuals who remained in MA had average expenditures increase \$529, while individuals who left MA experienced expenditure declines of \$682, a difference of more than \$1,200. If the individuals who did not change MA status serve as an adequate control for individuals who changed their MA status, these results suggest that MA enrollment substantially increases total costs to the Medicare program. Below, we explore the extent to which regression to the mean and other potentially important factors influence these estimates.

[Figure 2]

To further explore the pattern of enrollment and disenrollment from MA, Table 2 divides individuals into four distinct categories depending on whether they are enrolled in traditional Medicare or an MA plan in a base year and in the subsequent year. Figure 1 strongly suggests that exit from Medicare managed care was much more common in the early part of our period and that the reverse was true in the latter part of the period. Data from the MCBS, which allow us each year to follow several thousand survey respondents from one year to the next, are consistent with this pattern. As shown in the table, for every 1 person transitioning from traditional Medicare in 2001 to MA in 2002, there were 5.6 exiting from MA plans to return to traditional Medicare. This pattern is almost exactly reversed four years later, with 6.0 MA entrants for every one person leaving an MA plan to return to traditional Medicare.

[Table 2]

⁶ It is worth noting that this growth rate in average expenditures is substantially greater than the growth rate in Medicare's per-capita FFS costs from one year to the next. There are at least two reasons for this. First, by restricting attention to those observed in both $t-1$ and t , we ignore individuals who died in year $t-1$. As much previous research has shown, almost 25 percent of Medicare spending occurs in the last six months of life, and thus average spending in $t-1$ is lower than it otherwise would be. Second, we by definition exclude individuals who join the Medicare program in year t , most of whom are age 65 and thus would reduce average expenditures in that year.

An inspection of the data listed in Table 2 reveals that the first half of our study period was characterized by declining MA enrollment, while the second saw large increases. Combining these two periods, the total number of transitions from traditional Medicare to MA plans was slightly lower than the number of transitions in the opposite direction (641 versus 654).

Using the MCBS data, one can investigate how spending, self-reported health, and other outcome variables of interest change over time, and how these differences vary across groups. Table 3 examines these changes for individuals who are observed in the MCBS in two consecutive years. In addition to excluding individuals who exit the survey in the baseline year, this selection criterion will by definition exclude MCBS respondents who die in the first year or who enter the Medicare program in the second year. Both of these factors will cause the sample to have higher per-capita growth in Medicare spending and smaller improvements (or bigger declines) in average health status across the two years than would exist in the overall Medicare population, which is “refreshed” each year by a cohort of relatively cheap and healthy newly-eligible 65-year-olds.

[Table 3]

Consistent with this and as shown in the first panel of Table 3, average per-capita spending for those Medicare recipients observed in the MCBS in two consecutive years increases by more than 16 percent (from \$7,108 to \$8,286) across the two years. Similarly, the fraction of individuals reporting very good or excellent health declines by 1.7 percentage points while the fraction in poor or fair health increases by 0.6 percentage points.

The next four panels of Table 3 summarize this same information for four distinct groups: those who are in traditional Medicare in consecutive years, those who switch from traditional Medicare into MA, those who switch out of MA, and those in MA in both years. Average spending for those in FFS in both years increased by 20.0 percent in the early period and 15.4 percent in the late period. The increases for those transitioning into MA plans were much larger at 53.7 percent and 93.2 percent, respectively. This substantially higher growth is primarily driven by the much lower average spending of “switchers” in the base year.

The next two panels provide comparable summary statistics for those who exit MA plans and for those who remain in MA in both years. As Panel C shows, Medicare spending actually *declines* when individuals switch out of MA plans, by 6.2 and 14.6 percent during the early and late periods, respectively. This contrasts with the findings of some previous research (PPRC, 1997), which instead found that Medicare spending tends to increase following plan disenrollment because MA recipients are more likely to switch out when their health declines. The final panel shows that on average, Medicare spending grows by just 4.7 percent during the early period and 8.3 percent during the late period for those who remain in MA.

Taken as a whole, these cross-tabs suggest that the effect of MA enrollment on Medicare expenditures has increased between the early and late periods. The difference in dollar increase between the FFS stayers and MA joiners was \$1,537 during the early period and \$3,478 during the later period. Similarly, the difference between the dollar increase for MA stayers and MA disenrollees was \$848 during the early period and \$2,273 during the late period. In the section that follows, we explore the robustness of these cross tabs to formal regression analysis.

Regression Results

The difference between average expenditures for beneficiaries on MA and FFS in our sample is \$713, or 9.1% higher than the average spending for Medicare FFS beneficiaries in our sample and close to the MedPAC estimate of the overpayment to MA plans, which ranged from 3 percent prior to MMA to 12 percent in 2007. However, there are many important observable differences between the MA and FFS populations. Model (1) in Table 4 shows that controlling for some observable differences between the MA and FFS populations (Metro, Black, Hispanic, Female, education, age, and state of residence) decreases the estimated effect of MA enrollment on average Medicare Expenditures to \$257. In this specification and all subsequent ones, our measure of MA enrollment is the fraction of the recipient's eligible months during which they were enrolled in an MA plan.

Model (2) presents the first piece of evidence that failing to account for an individual's health status causes cross sectional estimates to understate the true impact of MA enrollment on average Medicare expenditures. As shown in Table 1, relative to FFS beneficiaries, MA recipients are much less likely to report being in fair or poor health and are much more likely to report being in very good or excellent health. While self-reported health is an imperfect measure of a beneficiaries' true health status, this variable is a very strong predictor of Medicare expenditures. Among the FFS population, a beneficiary who is in "poor" health had total Medicare expenditures that were, on average, more than 5 times greater than the corresponding average for a beneficiary who is in "excellent" health. After adding self reported health fixed effects, the estimate on the fraction of the year that a beneficiary spends on MA more than triples, increasing from \$257 [Model (1)] to \$842 [Model (2)]. In addition, in a specification not summarized in this table, adding controls for whether a beneficiary reported having cancer, stroke, heart problems, or diabetes increases the estimate still further to \$1,023.

Because of the difficult of adequately controlling for all of the differences between MA and FFS beneficiaries, our primary empirical strategy relies on the longitudinal nature of the MCBS. Models (3) through (5) of Table 4 transition from the cross-sectional analysis to the longitudinal empirical strategy. Regression (3) is identical to regression (2), except that (3) restricts attention only to individuals who were present in the MCBS in the previous year. This restriction reduces the sample size by 45% because it excludes both individuals who died or were not included in the MCBS interviews in the previous year. While this longitudinal sample is somewhat older than the full sample (73.5 v 74.3 years, on average) and has somewhat higher average Medicare expenditures (\$8,096 v \$8,286), the coefficient of interest barely changes, moving from \$842 in (2) to \$778 in (3). Models (4) and (5) further narrow the sample, restricting attention only to individuals who are on FFS in the previous year. The coefficient of interest in (4) is 57 percent larger than the corresponding estimate from (3), suggesting that the cross-sectional gap in average Medicare costs between FFS and MA may be larger for recent MA joiners than for the stock of MA beneficiaries, an issue that we explore further below.

Finally, Model (5) shows that including a linear control for baseline spending, a proxy for a beneficiary's underlying health, increases the estimate by 67 percent to \$2,044. This strongly suggests that MA joiners have lower average costs than individuals with comparable self-reported health in the FFS population. Note, however, that the coefficient of interest in (5) is considerably lower than either the dollar difference between the increase in the total Medicare expenditures of MA joiners and FFS stayers shown in Figure 2 [\$2,933] or the estimate if one re-estimates (5) and restricts the coefficient on lagged spending to be 1 [\$3,458], a regression that mirrors a standard difference-in-difference specification.

The difference between the estimate in (5) and the other two estimates reflects the effect of regression to the mean. The average baseline spending for MA joiners is \$5,059, compared with \$6,923 for FFS stayers. Individuals who have low medical spending in one year tend to have low medical spending in subsequent years, but the effect is attenuated over time.⁷ Failing to account for this regression to the mean can cause longitudinal estimates of MA enrollment on Medicare expenditures to overstate the true effect (Mello et al., 2003), an issue we explore further in Table 5.

[Table 4]

In the following table, we summarize the baseline estimates from the switcher regressions and show that, if anything, the estimate in model (5) of Table 4 understates the causal impact of MA enrollment on Medicare expenditures. As a starting point for considering regression to the mean in Medicare expenditures, Appendix Figure 1 shows that among the FFS population, the relationship between Medicare expenditures in years $t-1$ and year t is non-linear. Therefore, controlling for a linear function of lagged spending [Model (5) of Table 4] may yield biased estimates. To address this concern, Models (1) and (2) of Table 5 control flexibly for baseline spending. Specifically, Model (1) includes places beneficiaries into 21 bins based on their baseline total expenditure (1 bin for individuals with non-positive expenditures, and 20 equal size bins for individuals with strictly positive expenditures). Model (2) parallels (1), but also controls for a linear function of baseline spending to account for the fact that

⁷ Specifically, a simple regression of spending in year t on spending in year $t-1$ among the FFS population shows that the coefficient on spending in year $t-1$ is 0.425, and is significantly different from 1 at the 0.01% level.

spending is non-constant within the bins. In both models (1) and (2), the coefficient of interest grows slightly from the estimate of model (5) in Table 4.

Model (3) allows the effect of baseline spending on spending in the subsequent year to differ across Parts A and B. Including this flexibility is potentially important because Part B spending is substantially more persistent than Part A spending and because the ratio of Part A to Part B spending is lower for MA joiners than for FFS stayers.⁸ In Model (3), individuals are placed into 21 bins based on Part A spending (1 bin for individuals with non-positive Part A expenditures, and 20 equal size bins for individuals with strictly positive Part A expenditures) and into 21 bins based on Part B spending. The model also includes a linear control in Part A and Part B spending. In Model (3), our preferred specification, the coefficient of interest grows considerably, and is \$2,590, or 33% of average FFS spending in our sample. The effect is precisely estimated; we can rule out costs of less than \$1,679 with 95% confidence. In addition, the effect does not appear to be driven by extreme outliers or by the skewness of Medicare expenditures.⁹

Our estimate of \$2,590 reflects both the selection of beneficiaries into MA and the fact that, throughout much of our period, benchmark MA reimbursements were set at above average FFS costs. To investigate what fraction of our estimate is driven by selection alone, we use the same observable characteristics in Model (3) to estimate how much each beneficiary would have cost, had he remained in FFS. Specifically, we re-run a regression with all of the control variables as Model (3), restricting attention to individuals who remain in FFS in year $t-1$ and t . We then use these estimated coefficients to predict the cost for individuals who switched to MA. Because these estimates do not use cost data from the MA population, they do not reflect the effect of benchmark MA payments. These estimates suggest

⁸ Specifically, among the population that is in FFS in both t and $t-1$, a simple regression of Part A spending in year t on part A spending in year $t-1$ yields a coefficient 0.295. A parallel regression of Part B spending yields an estimate of 0.671. Note that in the baseline year, MA joiners have a ratio of Part A to Part B spending of 1.16, while the FFS stayers have a ratio of 1.25.

⁹ Specifically, if one re-runs Model (5) with an outcome variable as log total Medicare expenditures for observations with strictly positive Medicare expenditures, the coefficient on MA is 1.63 and has a t-stat of 34. Re-running the identical specification with a dummy variable for having strictly positive Medicare expenditures yields a coefficient on MA of 0.12 and a t-stat of 14. In addition, winsorizing the top and bottom 5 percent of the outcome variable and re-running the same specification yields an estimate of \$3,303 (t-stat 12) for the coefficient of interest.

that if the MA joiner population had remained in FFS, their average Medicare expenditures would have been \$6,983. Comparing this estimate to the average year t Medicare expenditures for the FFS stayer population (\$8,155) suggests that the selection effect accounts for roughly \$1,173. Repeating this exercise separately for the early (1999 – 2003) and late (2004 – 2006) periods suggests that, if anything, the extent of selection increased over time.

Finally, Models (4) and (5) explore the robustness of our results to controlling for two years of baseline spending. Such controls are potentially important for at least two reasons. First, because medical expenditures are highly variable, a single year of baseline spending may be a poor proxy for an individual's underlying health. Second, one potential concern with the switcher regressions is that individuals may defer care until they join MA, causing baseline spending to be abnormally low the year before a beneficiary joins MA. However, models (4) and (5) show that using two years of pre-enrollment data does not appear to influence the coefficient of interest. While changing the sample to include only people who are recorded in the MCBS as being on FFS for two previous years increases the coefficient of interest [Model (4)], adding a second year of baseline spending as a control barely changes the estimate [Model (5)]. It is worth noting that the coefficient estimate for this additional control variable is highly significant with a t-statistic of almost 10.¹⁰

[Table 5]

We now turn to exploring how the estimate has evolved over time with the shift to greater risk-adjustment. Models (1) – (3) of Table 6 present our estimates of the change in Medicare expenditures for MA switchers, controlling flexibly for baseline Part A and Part B spending as described above. Each regression includes an interaction for joining MA in 2004, 2005, or 2006, the three years for which capitation payments were based on the enhanced risk adjustment model. Model (1) controls for only

¹⁰ An alternative approach to address the concern that lagged spending does not fully capture an individual's baseline health is to instrument lagged spending with its own lagged value. If total Medicare expenditure for a person in a year is the sum of a constant term and a year-specific shock that is white noise and independent across years, then this IV strategy will account for the measurement error. A variant of Model (5) with this IV strategy yields a coefficient on the fraction of the year on MA that is to \$3,836 and is precisely measured. In addition, the coefficient on lagged spending (measured in \$1,000) increases to 777, consistent with the measurement error hypothesis.

baseline spending, year FE, and metro area status, (2) adds demographic controls, and (3) includes self reported health status. Across models (1) – (3), the coefficient on fraction of the year in MA ranges from \$1,037 - \$1,277, suggesting that from 1999 – 2003, individuals who joined MA cost approximately \$1,150 more than they were estimated to have cost, had they remained in FFS. The point estimates on the interaction of fraction of the year in MA and a dummy for being 2004 or later suggest that estimated overpayment to MA plans increased during our sample period. This coefficient ranges from \$1,418 to \$1,753, suggesting that the effect of joining MA in the later period is approximately 110 – 170 percent higher than in the earlier period, depending on the specification.

This increase in the effect of MA enrollment on expenditures could originate from two distinct channels: increased benchmark payments or changes in the payment to MA enrollees due to risk adjustment. The Medicare Modernization Act of 2003 increased benchmark payments to all counties, with some counties receiving increases of nearly 50 percent. At the time, MA payments to plans increased from an estimated 103 percent of FFS spending to 107 percent of FFS spending (MedPAC 2004), based on demographic factors. The overpayments continued to increase, reaching 112 percent in 2007 (MedPAC 2008). If one assumes that the estimated overpayment increased by seven percent of FFS costs between the early and late period, then the effect of MA on expenditures would mechanically increase by around \$600. The fact that our point estimates in Models (1) – (3) are all well above \$600 suggests that we find no evidence that the extent of selection decreased during the later part of our period, though it is also worth noting that \$600 is within the confidence interval for the coefficient on this interaction term.

[Table 6]

Alternative Explanations

As discussed above, there are three primary concerns with the individual level analysis from the MCBS. First, because an individual chooses to enroll in MA, an individual's costs had they remained in FFS may not be well approximated by comparing them to other individuals who are similar along observables. While we cannot rule out such a possibility with individual level data, the available

evidence cuts against the hypothesis that an individual's health deteriorates the year that they join MA. For example, as Table 3 shows, between the baseline year and the year of interest, the fraction of FFS stayers in "very good" or "excellent" health declines 1.5 percentage points, but increases 3.3 percentage points for the MA joiners. Similarly, there is a 1.1 percentage point increase in the fraction of FFS stayers who report being in "poor" or "fair" health between the two years of interest, whereas there is a 3.4 percentage point decrease for MA joiners.¹¹ In addition, comparing models (2) and (3) of Table 6 shows that controlling for self reported health has little effect on the estimate of interest.

The second concern is that the estimated effect of MA enrollment on Medicare expenditures that relies on MA switchers may not generalize to the general MA population. To explore this possibility, Models (4) and (5) of Table 6 shed light on how an individual's cost changes as she disenrolls from MA. Model (4) shows that Medicare recipients who leave MA plans cost on average \$1,252 less than observably similar individuals who remain in their plans. Because many of the individuals who leave MA would likely have been enrolled in the program for more than 1 year, these estimates show that our result is not limited only to new MA enrollees. Model (5) shows that, consistent with our earlier estimates for MA joiners, the effect of MA enrollment in this population on Medicare expenditures appears to have increased later in the period, although the effect is not estimated precisely.

Finally, these individual estimates do not capture spillovers of MA on the FFS population. To address this issue and related concerns, we next turn to analysis that utilizes aggregate county level data.

V. The Impact of MA Enrollment on Medicare Expenditures: County-Level Analysis

Empirical Strategy

To address the limitations of this individual level analysis, in this section we utilize annual county-level data on per-capita Medicare spending among those in fee-for-service Medicare.¹² We focus on elderly Medicare recipients, who account for approximately 85 percent of all program beneficiaries,

¹¹ One additional limitation with our study is that we do not investigate the effect of MA enrollment on health status.

¹² No comparable data is available on average per-capita spending at the county level for MA enrollees.

and consider data for the 2000 through 2008 period. We merge this data to annual county-level data on the fraction of Medicare recipients enrolled in MA plans, and summarize our sources for both sets of data in the Data Appendix.

Using both sets of data, we explore whether the change in the fraction of a county's Medicare recipients enrolled in MA plans is significantly related with the change in average per-capita costs among the county's fee-for-service population. If MA recipients would on average cost less than the typical FFS recipient, one would expect increases in MA enrollment to result in increases in average per-capita FFS costs. If instead MA recipients were not favorably selected, or if spillovers to the FFS population were sufficient to offset this, then one would expect no corresponding relationship.

Batata (2004) used comparable county-level data to investigate this same issue during the early 1990s. In this paper, the author utilized methods developed by Berndt (1991) and Gruber et al (1999) to estimate the difference between marginal cost (MC) of new MA enrollees and disenrollees and the average cost (AC) in the fee-for-service population. She demonstrates that, if one assumes the distribution of spending across counties differs only with respect to the mean, then the estimate for β_1 in the following specification yields the difference between MC and AC:

$$(3) \Delta AC_{FFS_{j,t}} = \beta_0 + \beta_1 * \Delta \ln(\% FFS_{j,t}) + \tau_t + \delta_j + \varepsilon_{j,t}$$

If there is favorable selection into MA plans and this is not more than offset by spillovers to the FFS population, then one would expect a negative estimate for β_1 . In this specification, j and t index counties and years, respectively. County effects are included to control for the possibility that there are unobserved time-invariant factors across counties that could bias the estimates while year effects are included to control for common changes nationally in per-capita Medicare expenditures. The key identifying assumption in this specification is that any unobserved factor that influences the change in a county's average FFS expenditures is not systematically related with the corresponding change in MA enrollment.

There are two advantages to our including data for the 2000 through 2008 period when estimating this specification. First, one potential concern with this aggregate county level analysis is that changes in

MA penetration may be correlated with time-varying characteristics of a county that influence average FFS costs. However, as shown in Figure 1, our sample period includes episodes of both a substantial reduction and substantial increase in MA enrollment. To the extent that there are omitted factors influencing FFS expenditures differentially in places with rapid MA growth, it is unlikely that these omitted factors would reverse direction when MA enrollment is declining. Second, this approach allows us to estimate the difference between marginal costs and average costs from MA joiners (when MA enrollment is growing late in the period) and from MA disenrollees (when it is declining), because as we showed in Section 3 these two different groups were the primary drivers of the enrollment changes during this period. To the extent that one is concerned that the effect of MA on spending in the first year does not translate to later years, our results should capture this with a different estimate between the two periods.

Finally, it is worth emphasizing that estimates from (3) recover a different parameter of interest from the MCBS analysis. The causal effect of MA enrollment on Medicare expenditures from the individual level data reflect both (a) the fact that MA joiners had different baseline costs from FFS stayers and (b) the fact that baseline payments to MA plans were different from average FFS spending. Because these county regressions focus only on FFS spending, they recover only (a), thus allowing us to directly explore whether selection became more or less prevalent in the later years of our sample.

Regression Results

The specifications summarized in Table 7 shed some light in this issue. In each specification, we weight each county-year observation by its share of the U.S. Medicare population in that year, and cluster our standard errors at the county level to account for possible serial correlation in the error term. Expenditure amounts are adjusted to 2007 dollars using the Consumer Price Index for all Urban Consumers (CPI-U), and average per-capita FFS expenditures during our period are approximately \$6,700. There are 3,110 counties included in the sample, with all counties in Alaska and a few other counties with missing data in 2 or more years excluded. With nine years of data and our use of a first-difference model, there are 8 observations for each county.

In the first specification, the dependent variable is the change in a county's per-capita Medicare FFS costs from one year to the next. The estimate of -1,191 for β_1 is highly significant with a t-statistic of 6.8, and suggests that Medicare spending for the marginal MA enrollee/ disenrollee *had they remained in FFS* would on average have been approximately \$1,200 less than for the average FFS enrollee. Put another way, the marginal cost is approximately 18 percent less than the average cost of FFS recipients. This county-level estimate corresponds quite closely to the MCBS estimate in the previous section, which suggested that the selection effect can explain \$1,173 of the observed difference in Medicare expenditures between the FFS stayers and MA joiners.

One limitation with these results is that the changes may largely reflect selection on readily observable factors, such as age. If, for example, MA enrollment growth is driven by differential increases among individuals in their late 60s, then one would expect an increase in average FFS costs even in the absence of the favorable selection described above. The second specification accounts for this by instead using adjusted per-capita costs as the dependent variable. While the estimate of -914 is smaller than the corresponding one in the first specification, it remains highly significant. More importantly, this adjustment reduces the estimate by less than one-fourth.

In the next two specifications, we examine whether the selection appears to be greater from either Part A or Part B expenditures. Results from Batata (2004) indicated that most of the selection occurred on Part A expenses. In contrast to this, our findings suggest an almost equal amount of selection from Part A and Part B expenditures. This could potentially be attributable to the risk adjustment methodology used during our period, as it compensated plans more for MA enrollees with recent hospital admissions.

In the fifth specification, we investigate whether the amount of selection appears to differ in the first few years of the period from the latter several years by interacting our key explanatory variable with a POST2004 indicator. At just 166, the estimate for the coefficient on this interaction is small when compared with the main effect estimate of -1,268, and it is statistically insignificant. Thus there is little evidence to suggest that the move to greater risk adjustment beginning in 2004 substantially affected the amount of selection.

One limitation with our MCBS results is that we are able to consider data through just 2006. Thus in our sixth and final specification, we explore whether the relationship is different in 2007 and 2008 by interacting our key explanatory variable with an indicator for whether the year is 2007 or 2008. The estimate for the coefficient on this interaction is once again small and statistically insignificant, suggesting that the amount of favorable selection is similar in this latter part of the period when the full risk adjustment methodology was being used.

We explored the robustness of our results in a number of ways. For example, we estimated the model using the change in the log of per-capita FFS expenditures as the dependent variable and obtained a qualitatively similar estimate of -0.143 (with a t-statistic of 6.5), suggesting that marginal costs are almost 15 percent lower than average costs. Additionally, when we focused exclusively on the larger counties with more than 10 thousand Medicare recipients in every year, the magnitude of our estimate for β_1 was slightly higher at -1269.

Taken together, the results in this section indicate that there was favorable selection into MA plans during the 2000 to 2008 period, which is consistent with the evidence for the 1980s and 1990s. Perhaps more importantly, this selection has persisted in recent years, despite the move by CMS to a reimbursement methodology that uses risk-adjustment to set capitation payments.

And finally, our results provide little evidence to suggest that positive spillovers to the Medicare FFS population are sufficient to offset these expenditure effects. If increases in MA enrollment caused FFS Medicare expenditures to decrease, we would expect that the MCBS results (which do not account for this possibility), to show a larger selection effect than the county results. Instead, our estimates for the difference between MC and AC from the county regressions are quite similar to what one would expect given our MCBS results if there were no spillovers at all; both approaches suggest that selection can account for roughly \$1,200 of the observed difference in cost between the average cost of the FFS population and the marginal cost of the MA population. In addition, because (3) studies the impact of changes in MA penetration on contemporaneous changes in FFS costs, this regression does not capture spillovers that take more than one year to diffuse to the FFS population. However, re-estimating (3) and

adding a lagged value of the variable of interest shows that, once contemporaneous changes in MA penetration have been taken into account, there is no evidence that lagged changes in MA penetration influence average FFS costs.¹³

VI. Evidence on Selection Patterns by Health Status

Empirical Strategy

Having found strong evidence of favorable selection into MA plans from analysis on spending, we now look at other possible indicators of selection, both before and after the introduction of risk adjustment. In this section, we use the MCBS to determine if (a) individuals who join MA plans are in better health than individuals who remain in FFS plans and (b) if the stock of individuals enrolled in MA plans are healthier than individuals in FFS plans. In both cases, we explore the extent to which the patterns changed over the sample period. For this analysis, we focus on two indicators of health: self-reported history of various chronic and acute illnesses and self-reported subjective measures of individual's health.

By differentially increasing the capitation payment for individuals with certain conditions, risk adjustment should increase plans' incentives (or reduce their disincentives) to attract and retain individuals with a history of specific acute or chronic illness. Although the MCBS does not contain the richness of detail used by CMS for actual risk adjustment, we focus on several acute and chronic illnesses that are included in the MCBS in every survey year that are also included in the MA risk adjustment calculation.¹⁴ The baseline specification is

¹³ Specifically, the coefficient on lagged have change in log percent FFS is -\$221 and has a p-value of 0.389.

¹⁴ The acute illnesses include an episode in the past year of cancer, hip fracture, heart attack, and stroke. The chronic illnesses include a history of rheumatoid arthritis, angina, diabetes, Parkinson's disease, Alzheimer's disease, and emphysema/asthma/COPD. Although not comprehensive relative to the CMS risk adjustment model, these health conditions affect 50 percent of our sample, thus covering a large share of the population with risk scores above the baseline. We restrict our sample to survey respondents residing in the community, because institutional respondents are asked a slightly different set of questions about their health. In addition, MA risk adjustment calculates different scores depending on whether the enrollee is institutionalized.

$$(4) \quad MA_{jt} = \alpha_t + \mu * Condition_{jt} + \beta * Condition_{jt} * Year \geq 2004 + \gamma * METRO_{jt} + \sum \lambda_k * I(State_{jt} = k) + \varepsilon_{jt}$$

where MA_{jt} is a dummy variable for person i 's enrolling in MA in year t , $Condition_{jt}$ is a dummy variable for having a given disease or self-reported health, μ , and $Year \geq 2004$ is a dummy variable for t being 2004 or later. Each regression is estimated as a probit, with marginal coefficients reported, and we estimate this specification using both the full sample and restricting to a subsample of individuals who were on FFS in year $t-1$. The coefficient of interest is β , which captures the extent to which an individual with a given condition is more likely to enroll in MA in the 2004 and later period, relative to individuals without the condition.

Results

Table 8 summarizes the results of this analysis. In Panel 1, we examine acute and chronic conditions. For the “stock” analysis (under heading (A)), in all but one case, the coefficients of interest are not statistically different from 0, indicating that we cannot reject the null of no differential change in the probability of being enrolled in MA. In the switching equation (under heading (B)), we see no relationship between any of the health indicators and MA enrollment, either before or after 2003. To give a sense of the power of the estimates, we would need to see a change in the enrollment of beneficiaries with emphysema of around 1.5 percentage points in the stock equation and 0.4 percentage points in the switcher equation, from a sample mean of around seven percent in both specifications, for an estimate to be significantly different from zero at the 90 percent confidence level.

We then divide the sample according to the self-reported health status of the respondent: excellent, very good, good, fair, and poor. If MA plans do not experience favorable selection or have experienced less favorable selection since the introduction of comprehensive risk adjustment, we might expect to see that reflected in the enrollment rates of individuals with the aforementioned health issues. In Panel 2 on Table 8, we see that people reporting being in fair or poor health are significantly less likely

to be enrolled in an MA plan than people in excellent health (the omitted category) before 2004. In 2004 and later, we do see the pattern reversed for people in fair health, and people in good health are also more likely to report being in an MA plan than people in excellent health. However, we do not see that pattern in the switcher regression.

To summarize, using various measures of health of the Medicare population, we find little evidence to support the hypothesis that MA plans became more likely to enroll individuals with high health risks following the shift to risk adjustment, which is consistent with our expenditure results above.

[Table 8]

VII. How do MA plans attract healthier Medicare beneficiaries?

In the face of strong evidence of favorable selection in MA plans, it is worth exploring how such selection is achieved. In contrast to private plans for the non-elderly in the non-group market, Medicare Advantage plans may not legally deny enrollment to any Medicare beneficiary living in the plan's geographic area of operation. As such, the overt denials and "cream-skimming" commonly associated with private non-group plans are unlikely to explain the large selection effects found above and in much previous work on Medicare managed care.

We show in this section that MA plans appear, with respect to their cost-sharing and provider performance, to appeal to healthier beneficiaries. As such, utility-maximizing individuals might choose to enroll in MA plans when healthy and disenroll if they fall ill.¹⁵

Empirical Strategy

¹⁵ Of course, individuals need not always be the active party in the selection process, and past studies have shown that MA plans appear to advertise in ways that target the healthy (Mehrotra, Grier and Dudley, 2006; Neuman et al., 1998).

In every year of our sample, the MCBS asks a subsample of enrollees in the survey to rate different aspects of their medical care.¹⁶ We investigate how satisfaction varies by MA enrollment and health in the following equation:

$$(5) \text{ Satisfaction}_{it} = \beta \text{MA}_{it} * \text{H}_{it} + \gamma \text{MA}_{it} + \delta \text{H}_{it} + \theta \text{X}_{it} + e_{it}$$

where *Satisfaction* is a categorical variable coded one (“very dissatisfied”) to four (“very satisfied”), *MA* is an indicator variable for MA enrollment (equal to 1 for those with half or more months in MA and 0 otherwise), *Health* is a categorical variable for self-reported health coded one (“poor”) to five (“excellent”), **H** is a vector of the health-category fixed effects, and **X** is a vector of covariates based on the standard demographic risk-adjustment model (gender, age, disability and Medicaid status) as well as state and year fixed effects.

The coefficient on the interaction term (β) sheds light on whether MA plans may be differentially attractive to healthy enrollees. A positive coefficient suggests that at least some of the positive selection documented in the earlier sections could arise from healthy individuals differentially choosing to enroll in MA plans, as it would suggest that healthier individuals are more satisfied with their MA plans than their counterparts in worse health.

For three reasons, this section focuses on estimating cross-sectional regressions using the entire sample of enrollees who answered the satisfaction questions, instead of focusing on those switching between MA and FFS. First, unlike costs, which are based on entirely different systems in FFS and MA, there is no a priori reason individuals would answer satisfaction questions differently in MA or in FFS. Second, it may take some time to fully evaluate one’s health plan and those who have just switched may have inadequate experience to judge. Third, we will generally not focus on average satisfaction differences (γ , in equation 5), as one could imagine that those enrolling in MA instead of the traditional option may generally be, say, merely harder to please, but on the interaction term (β_2). Alternative

¹⁶ Unlike most of the variables used in their previous sections of the paper, these questions are only asked of those enrollees alive and Part-A- and Part-B-eligible all twelve months of the survey year.

explanations are obviously still possible, but they would have to explain why individuals who are especially unhappy when sick would be differentially attracted to MA plans.

Main results

The first nine rows of Table 9 show the estimated coefficients on the MA main effect and the MAxHealth interaction term from equation (1). Note that we have demeaned the Health variable, so that the MA main effect estimate represents the difference in satisfaction for someone with average health (between “good” and “very good,” or three and four on the five-point scale).

The first row shows the results based on satisfaction with your “overall medical care.” The coefficient on the MA main effect is negative and significant, though, as discussed, caution should be exercised before interpreting it causally. Although not shown in the table, the health fixed-effects indicate that satisfaction among those in FFS is a positive function of health—not surprising, as individuals are likely to be in a more positive frame of mind in general when they feel they are in good health.

Our main focus, however, is the positive and significant coefficient on the MAxHealth interaction term. The coefficient indicates the positive relationship between health and satisfaction exhibited among FFS enrollees is even stronger among MA enrollees. Put differently, the difference in satisfaction between a healthy and sick person is significantly greater in MA than in FFS.

Strikingly, this pattern of a negative and significant main effect (MA enrollees report lower satisfaction on average) and a positive and significant interaction effect (differences in satisfaction between the health and sick are greater in MA than in FFS) is repeated throughout the table. The MA coefficient is negative and significant for all nine categories except for the ability to receive one’s medical care in the same location (for which it is positive and significant). The interaction coefficient is positive for eight of the nine categories, positive and significant for six, and never negative and significant.

As discussed earlier, people attracted to MA plans may simply have different satisfaction patterns than people who tend to remain in FFS, independent of the medical care they receive. While we prefer the cross-sectional specification already estimated, the tenth row of Table 9 reports the estimate from a

fixed-effects regression otherwise identical to equation (1), using the “overall satisfaction” as the dependent variable.

The negative main effect essentially disappears. This result arises in large part because the year an enrollee switches coverage types, whether from FFS to MA or MA to FFS, is associated with an increase in satisfaction, perhaps simply reflecting the novelty of a new situation (results available upon request). However, the interaction effect is still positive, though has a p -value of 0.117.

[Table 9]

Additional results

One difficulty in interpreting the results in Table 9 is that there are not meaningful units to the satisfaction variable. In the next set of results, we now use an indicator variable for whether a respondent reports being “dissatisfied” or “very dissatisfied” with his medical care. Furthermore, instead of interacting the MA indicator variable with the linear health variable, we interact it with each health fixed effect:

$$(6) \text{ Prob(Dissatisfied)} = \sum_{m=1}^5 \beta_m MA_{it} \cdot H_{mit} + \gamma MA_{it} + \delta H_{it} + \theta X_{it} + e_{it}$$

This specification allows us not only to test another functional form, but to see where exactly in the health-status distribution the satisfaction differences between FFS and MA are greatest, and to investigate whether the relationship is monotonic.

Figure 3 displays the results when satisfaction with “overall medical care” serves as the underlying outcome variable of interest, and plots health status category on the x-axis and the probit coefficient (reported as marginal changes in probability) on the corresponding $MA \times Health\text{-category}$ interaction term on the y-axis. Whereas only about three percent of all enrollees report being unsatisfied with their care, among enrollees in poor health those in MA are 1.5 percentage points more likely to report dissatisfaction than are those in FFS. While MA enrollees are more dissatisfied than FFS enrollees regardless of health status, the difference decreases monotonically with health status.

[Figure 3]

Figure 4 plots the corresponding figure for dissatisfaction with out-of-pocket costs. Again, dissatisfaction with MA relative to FFS decreases with health status (though is not perfectly monotonic). Here, the difference in dissatisfaction actually flips, with MA outperforming FFS for those in the best two health categories. We omit the remaining figures for the sake of space, but they generally show the same qualitative picture, as one would expect based on the results in Table 9.

[Figure 4]

These two figures as well as the results in Table 9 suggest that MA plans are differentially attractive to those in good health. While we caution taking the level differences between MA and FFS too literally, Figure 4 suggests that, relative to enrollment in FFS, healthy (sick) beneficiaries will save (lose) money in MA, which would provide individuals a strong incentive to self-select into MA plans based on their health status.

Finally, we examine whether enrollees “vote with their feet.” Each year an enrollee would have the option of switching from FFS to MA or MA to FFS, and assuming he makes that decision based in part on his current satisfaction, we would expect, for example, that FFS enrollees are more likely to remain in FFS than are MA enrollees to remain in MA, and that this difference is greatest for those in poor health.

The final row in Table 9 shows the results from a probit estimation of whether an individual remains in his current coverage type the following year on the covariates included in equation (1). Note that the sample shrinks, as this estimation requires that individuals be in the sample for two consecutive years. If enrollees “vote with their feet,” then the coefficients should have the same pattern as the other satisfaction regressions in the table, and indeed they do.

As in the other sections, we also looked to see if the results differed across time periods. We do not report the results here, but we found nothing to suggest satisfaction levels across plan type and health status changed in the period after the introduction of risk adjustment.

Taken together, the results in this section provide evidence that the package of benefits, cost-sharing, and service MA plans provide differentially appeal to individuals in good health. Thus it is perhaps not surprising that the individuals who choose to enroll in MA have substantially lower Medicare expenditures than their counterparts who remain in FFS.

VIII. Discussion and Conclusion

Previous literature has demonstrated that enrollees into Medicare Advantage plans cost Medicare more than if they had remained in FFS, particularly since the Medicare Modernization Act, which set the capitation rate at an amount at least equal to the average local FFS costs. MedPAC's estimates of overpayments have grown steadily since MMA, from 103 percent of FFS spending, to 107 percent just after passage, to 114 percent in 2009. The Affordable Care Act seeks to reduce the overpayment by setting the benchmark reimbursement rate at 95 percent of local FFS costs in the highest spending counties, rising up to 115 percent of local FFS costs in the lowest spending counties. The CBO estimates that this change will reduce the value of extra benefits not covered by FFS to MA enrollees by half in 2019.

Yet generous benchmarks are only one mechanism for overpayment. The other avenue is through favorable selection into plans, in which plans enroll beneficiaries with expected costs below the reimbursement rate. Such favorable selection has been well documented in the literature, and as a consequence, Medicare began risk adjusting payment rates to private plans in 2000, gradually increasing the complexity of the risk adjustment model as well as the share of plan reimbursement adjusted for risk. Effective risk adjustment would leave high benchmarks as the only source of overpayment. But our research finds that favorable selection, and consequently overpayments, have not fallen since the introduction of the most comprehensive model of risk adjustment in 2004.

We observe virtually no increase in enrollment in MA plans of people with one of a variety chronic or acute illnesses, nor for people in poor health. Using county-level data, we find similar increases in average per-capita FFS costs for given increases in MA enrollment that are almost identical

to those implied by our MCBS analyses. Most strikingly, we find that overpayments for new MA enrollees have doubled since 2004, exceeding the estimated effects of higher benchmarks. Such selection may seem hard to explain, given the limits on benefit design and marketing that MA plans face. Yet we find on a number of satisfaction measures with health care quality and cost, the difference in satisfaction levels between healthy and sick enrollees is higher for MA enrollees than FFS.

Effectively combating favorable selection is not a policy matter restricted to Medicare. Similar questions confront the new health insurance exchanges, to be established in 2014 by the Affordable Care Act. The exchanges, estimated to cover 24 million people in the individual and small group commercial market by 2018, will exhibit many of the same features as the MA market: mandated benefits packages, guaranteed issue, restrictions on premium variations, and considerable direct purchase by individuals. Also, to protect against adverse selection, the new law mandates cross subsidization between insurers through risk adjustment. In light of the results presented here, one question is how well a risk adjustment mechanism will reduce adverse selection in the exchanges.

There are at least three key differences between the MA market and the exchanges, with respect to risk adjustment, that are important to note. First, FFS Medicare claims provide an extraordinarily rich data source on health and spending on its beneficiary population for its risk adjustment model. A comparable data source does not exist for the commercial market, which leads to the question of how risk adjustment will be implemented, and whether one can reasonably expect it to perform near the level of Medicare's model. Second, there will be no comparable plan on the exchange to FFS Medicare, which bears the brunt of adverse selection in Medicare. If risk adjustment on the exchange does not eliminate adverse selection, the question of how that will manifest itself is not obvious. Possible alternatives are lower quality care across the exchange for high risk individuals as plans seek to avoid them, or perhaps the emergence of plans that do provide quality care to high risk individuals but at higher premiums that incorporate the increased risk that risk adjustment misses. Third, the health profile of the non-Medicare population is very different from that of the Medicare population. Rates of chronic illness, and thus, predictable risk, are considerably lower among the non-elderly. The 2008 National Health Interview

Survey finds that 25 percent of the elderly population considers themselves in fair or poor health, compared to eight percent of the nonelderly population (Adams 2009). Such differences may reduce the magnitude and the incentives for adverse selection on the exchanges relative to MA.

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Table 1: Cross Sectional Summary Statistics

	Early Years		Late Years	
	MA	FFS	MA	FFS
64 and under years old	0.068 (0.251)	0.127 (0.332)	0.077 (0.267)	0.144 (0.351)
65 - 69 years old	0.202 (0.401)	0.152 (0.359)	0.148 (0.355)	0.159 (0.366)
70 - 74 years old	0.256 (0.437)	0.230 (0.421)	0.260 (0.439)	0.209 (0.407)
75 - 79 years old	0.227 (0.419)	0.206 (0.404)	0.222 (0.415)	0.192 (0.394)
80 - 84 years old	0.137 (0.344)	0.149 (0.356)	0.165 (0.372)	0.152 (0.359)
85 and over years old	0.111 (0.314)	0.138 (0.344)	0.128 (0.334)	0.144 (0.351)
College Graduate	0.130 (0.336)	0.144 (0.351)	0.137 (0.344)	0.156 (0.363)
Not a HS Graduate	0.306 (0.461)	0.336 (0.472)	0.298 (0.457)	0.291 (0.454)
Black	0.100 (0.300)	0.093 (0.290)	0.108 (0.310)	0.096 (0.295)
Hispanic	0.036 (0.187)	0.021 (0.142)	0.039 (0.194)	0.017 (0.129)
Female	0.569 (0.495)	0.573 (0.495)	0.590 (0.492)	0.563 (0.496)
Metro Area	0.966 (0.181)	0.717 (0.450)	0.962 (0.190)	0.719 (0.449)
Health Poor or Fair	0.218 (0.413)	0.288 (0.453)	0.217 (0.413)	0.273 (0.446)
Health Very Good or Excellent	0.458 (0.498)	0.391 (0.488)	0.445 (0.497)	0.408 (0.492)
Total Expenditure	\$7,920 (4897.373)	\$7,445 (16797.480)	\$9,706 (8442.128)	\$8,566 (18102.570)
Sample Size	8,582	49,455	4,620	27,961

Note: early years are 1999 - 2003. Late years are 2004 - 2006.

The unit of observation is a person in a year. Drops all person-year observations in which a person is eligible for only Part A or Part B for any part of the year or if the person is eligible for Medicare only because of having ESRD. An individual in a given year is classified as being on MA if she is on MA for more than half of the months for which she is Medicare eligible in that given year.

The means (standard deviations) are reported for each variable.

Table 2: Frequency Distribution of Transitions

Baseline Year	Next Year	FFS	FFS	MA	MA	Total Remain in Sample	Leave Sample	Total in Sample
		→ FFS	→ MA	→ FFS	→ MA			
1999	2000	6,096	73	72	1,150	7,391	4,549	11,940
2000	2001	6,026	32	177	1,056	7,291	4,475	11,766
2001	2002	6,165	33	189	929	7,316	4,277	11,593
2002	2003	6,142	37	91	822	7,092	4,375	11,467
2003	2004	5,967	62	33	842	6,904	4,367	11,271
2004	2005	5,845	83	39	849	6,816	4,096	10,912
2005	2006	5,547	321	53	805	6,726	4,150	10,876
1999 - 2002		24,429	175	529	3,957	29,090	17,676	46,766
2003 - 2005		17,359	466	125	2,496	20,446	12,613	33,059
Total		41,788	641	654	6,453	49,536	30,289	79,825

Note: Unit of observation is a person in a year. Drops all person-year observations in which a person is eligible for only Part A or Part B for any part of the year or if the person is eligible for Medicare only because of having ESRD. An individual in a given year is classified as being on MA if she is on MA for more than half of the months for which she is Medicare eligible in that given year.

Table 3: Summary Statistics by Transition Type

Panel A: FFS → FFS						
	Early Years			Late Years		
	<u>t</u>	<u>t+1</u>	<u>Change</u>	<u>t</u>	<u>t+1</u>	<u>Change</u>
Total Expenditure	\$6,383	\$7,659	\$1,276	\$7,616	\$8,792	\$1,176
Health Very Good or Excellent	0.399	0.384	-0.015	0.416	0.405	-0.011
Healthy Poor or Fair	0.280	0.291	0.011	0.266	0.273	0.007
N		24,429			17,359	

Panel B: FFS → MA						
	Early Years			Late Years		
	<u>t</u>	<u>t+1</u>	<u>Change</u>	<u>t</u>	<u>t+1</u>	<u>Change</u>
Total Expenditure	\$5,237	\$8,050	\$2,813	\$4,994	\$9,648	\$4,653
Health Very Good or Excellent	0.403	0.436	0.033	0.388	0.398	0.010
Healthy Poor or Fair	0.308	0.274	-0.034	0.285	0.306	0.021
N		175			466	

Panel C: MA → FFS						
	Early Years			Late Years		
	<u>t</u>	<u>t+1</u>	<u>Change</u>	<u>t</u>	<u>t+1</u>	<u>Change</u>
Total Expenditure	\$7,679	\$7,204	-\$475	\$10,393	\$8,873	-\$1,520
Health Very Good or Excellent	0.456	0.437	-0.020	0.349	0.340	-0.008
Healthy Poor or Fair	0.218	0.218	0.001	0.256	0.295	0.039
N		529			125	

Panel D: MA → MA						
	Early Years			Late Years		
	<u>t</u>	<u>t+1</u>	<u>Change</u>	<u>t</u>	<u>t+1</u>	<u>Change</u>
Total Expenditure	\$7,871	\$8,245	\$374	\$9,078	\$9,831	\$753
Health Very Good or Excellent	0.472	0.453	-0.019	0.483	0.451	-0.031
Healthy Poor or Fair	0.207	0.216	0.008	0.190	0.202	0.012
N		3,957			2,496	

Note: Early years are 1999 - 2003. Late years are 2004 - 2006.

Table 4: Cross Sectional Spending Results

	(1)	(2)	(3)	(4)	(5)
Fraction of Year on MA	257*	842***	778***	1225**	2044***
	(147)	(138)	(178)	(491)	(459)
Lagged total expenses, \$1,000					367***
					(19)
Live in Metro Area?	1133***	1208***	1276***	1332***	885***
	(184)	(176)	(215)	(227)	(190)
Black	2571***	2406***	2691***	2866***	2219***
	(349)	(335)	(427)	(507)	(425)
Hispanic	635	741*	499	-210	-291
	(443)	(430)	(500)	(630)	(495)
Female	-708***	-889***	-933***	-956***	-731***
	(145)	(1380)	(172)	(199)	(166)
Mean of DepVar	8,096	8,096	8,286	8,178	8,178
Health FE?		X	X	X	X
Sample	Full	Full	In Sample	FFS Lst	FFS Lst
			Lst Yr	Yr	Yr
Observations	90,617	90,617	49,535	42,428	42,428

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Data used for this table is from the Medicare Current Beneficiary Survey (MCBS). The outcome variable is total spending, adjusted using the CPI to be 2007 dollars. Standard errors clustered at the individual level. Year FE are included in all specifications. Age FE include a dummy variable for each age, measured in years and are included in all regressions. Education FE include dummy variables for each educational attainment group (no schooling, nursery school to 8th grade, 9th to 12th grade, but no diploma, high school graduate, vocational, technical, business, etc., some college, but no degree, associate's degree, bachelor's degree, post graduate degree, as well as dummy variables for the question not being answered properly) and are also included in each regression. Finally, each specification includes state of residence fixed effects. Health fixed effects include dummy variables for each self reported health status, as well as separate dummy variables for a missing or unknown variable. Models (1) and (2) include all relevant MCBS observations. (3) restricts attention only to individuals who were in the MCBS in the previous year. (4) and (5) consider only individuals in FFS in the previous year.

Table 5: Switcher Results: Full Period

	(1)	(2)	(3)	(4)	(5)
Fraction of Year on MA	2082*** (469)	2186*** (463)	2590*** (464)	2763*** (953)	2740*** (940)
Lagged total expenses, \$1,000		276*** (56)		377*** (25)	324*** (25)
Twice lagged total expenses, \$1,000					172*** (18)
Lagged Part A Spending, \$1,000			76 (76)		
Lagged Part B Spending, \$1,000			862*** (70)		
Live in Metro Area?	868*** (193)	817*** (189)	849*** (187)	610** (300)	507* (297)
Black	2320*** (439)	2259*** (425)	1879*** (407)	2856*** (7730)	2662*** (768)
Hispanic	-281 (519)	-264 (494)	-395 (499)	230 (857)	141 (828)
Female	-894*** (172)	-846*** (166)	-818*** (163)	-722*** (272)	-724*** (269)
Mean of DepVar Spending Control	8,178 20 bins. Tot exp	8,178 Lin and 20 bins. Tot exp	8,178 Lin and 20 bins. Pt A/B	8,339 Lin Lst Yr	8,339 Lin Lst 2 Yrs
Sample	FFS Lst Yr	FFS Lst Yr	FFS Lst Yr	FFS Lst 2 Yrs	FFS Lst 2 Yrs
Observations	42,428	42,428	42,428	16,916	16,916

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Data used for this table is from the Medicare Current Beneficiary Survey (MCBS). The outcome variable is total spending, adjusted using the CPI to be 2007 dollars. Standard errors clustered at the individual level. Year FE are included in all specifications. Age FE include a dummy variable for each age, measured in years, of individuals in the data and are included in all regressions. Education FE include dummy variables for each educational attainment group (no schooling, nursery school to 8th grade, 9th to 12th grade, but no diploma, high school graduate, vocational, technical, business, etc., some college, but no degree, associate's degree, bachelor's degree, post graduate degree, as well as dummy variables for the question not being answered properly and are also included in each regression. Finally, each specification includes state of residence fixed effects. Health fixed effects include dummy variables for each self reported health status, as well as separate dummy variables for a missing or unknown variable. Models (1) - (3) restrict attention only to individuals who were in FFS in the previous year. (4) and (5) consider only individuals in FFS in the previous two years.

Table 6: Switcher Results: Early and Late Periods

	(1)	(2)	(3)	(4)	(5)
Fraction of Year on MA	1037**	1093**	1277**	1252**	1029
	(526)	(530)	(540)	(608)	(652)
Fraction of Year on MA * Year \geq 2004	1753**	1611**	1418*		1090
	(752)	(751)	(759)		(1649)
Lagged Part A Spending, \$1,000	65	69	76	322	309
	(76)	(75)	(76)	(549)	(548)
Lagged Part B Spending, \$1,000	880***	876***	863***	1262***	1265***
	(71)	(71)	(71)	(267)	(265)
Live in Metro Area?	637***	739***	879***	439	462
	(155)	(158)	(156)	(413)	(411)
Black		1815***	1718***	545*	538*
		(388)	(387)	(305)	(304)
Hispanic		-1137**	-1088**	333	325
		(468)	(474)	(379)	(377)
Mean of DepVar	8,178	8,178	8,178	8,935	8,935
Sample	FFS Lst Yr	FFS Lst Yr	FFS Lst Yr	MA Lst Yr	MA Lst Yr
Age FE?		X	X	X	X
Education FE?		X	X	X	X
Health FE?			X	X	X
Observations	42,428	42,428	42,428	7,107	7,107

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Data used for this table is from the Medicare Current Beneficiary Survey (MCBS). The outcome variable is total spending, adjusted using the CPI to be 2007 dollars. Standard errors clustered at the individual level. Year FE include a dummy variable for each year in the MCBS sample (1999 - 2006). Age FE include a dummy variable for each age, measured in years and are included in all regressions. Education FE include dummy variables for each educational attainment group (no schooling, nursery school to 8th grade, 9th to 12th grade, but no diploma, high school graduate, vocational, technical, business, etc., some college, but no degree, associate's degree, bachelor's degree, post graduate degree, as well as dummy variables for the question not being answered properly and are also included in each regression. Finally, each specification includes state of residence fixed effects. Health fixed effects include dummy variables for each self reported health status, as well as separate dummy variables for a missing or unknown variable. Models (1) - (3) restrict attention only to individuals who were in FFS in the previous year. Models (4) and (5) consider only individuals who were in MA the previous year.

Table 7: Relationship between Change in County-Level Per-Capita Medicare Expenditures and Change in Log Fraction of Recipients in FFS

	(1) Per Capita	(2) Adj Per Capita	(3) Part A Only	(4) Part B Only	(5) Per Capita	(6) Per Capita
Change Log(Fraction FFS)	-1191***	-914***	-648***	-543***	-12678***	-1118***
	(181)	(203)	(956)	(137)	(250)	(217)
Change Log(Fraction FFS) * Year \geq 2004					166	
					(285)	
Change Log(Fraction FFS) * Year \geq 2007						-398
						(397)
Mean of DepVar	236	231	101	135	236	236
R-squared	0.228	0.362	0.203	0.324	0.228	0.228
Observations	24,878	24,876	24,878	24,878	24,878	24,878

Standard errors in parentheses

Each column summarizes the results from a different specification in which the dependent variable (listed at the top of each column) is a measure of the change in a county's per-capita Medicare expenditures from one year to the next. Data includes all U.S. counties (with the exception of those in Alaska and a few others with missing data on Medicare enrollment in one or more years) in each year from 2000 through 2008 (and thus there are eight first-differences for each county). All specifications include 8 year effects, 3,110 county effects, and are weighted by each county's share of the U.S. Medicare population in each year. Dollar amounts are adjusted to 2007 dollars using the CPI-U and standard errors are clustered at the county level.

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 8: Impact of health indicators on probability of being in MA or joining MA

	Panel 1			Panel 2		
	Probability of being enrolled in MA			Probability of joining MA, conditional on being in FFS		
	Marginal probability	Standard error	Mean of variable	Marginal probability	Standard error	Mean of variable
Specification 1: Health conditions						
History in last year of given illness						
Cancer	-0.0004	0.0093	0.0296	-0.0009	0.0048	0.0289
Hip fracture	-0.0239	0.0148	0.0072	0.0000	0.0088	0.0071
Heart attack	-0.0027	0.0108	0.0248	0.0052	0.0062	0.0259
Stroke	-0.0139	0.0098	0.0249	-0.0033	0.0037	0.0271
History of given chronic illness						
Rheumatoid arthritis	0.0048	0.0066	0.1177	-0.0004	0.0023	0.1121
Angina	-0.0148 **	0.0058	0.1343	-0.0016	0.0020	0.1334
Diabetes	0.0108 **	0.0055	0.2018	-0.0011	0.0019	0.1968
Parkinson's	-0.0206	0.0159	0.0132	0.0012	0.0073	0.0127
Alzheimer's	-0.0116	0.0104	0.0311	-0.0056	0.0027	0.0268
Emphysema	-0.0089	0.0056	0.1596	-0.0008	0.0020	0.1595
<i>Enrolled in MA after 2003</i>						
History in last year of given illness						
Cancer	-0.0236	0.0142	0.0120	-0.0045	0.0039	0.0136
Hip fracture	0.0761 *	0.0405	0.0030	-0.0071	0.0033	0.0031
Heart attack	-0.0139	0.0163	0.0087	-0.0028	0.0041	0.0105
Stroke	0.0136	0.0192	0.0085	0.0007	0.0069	0.0112
History of given chronic illness						
Rheumatoid arthritis	-0.0029	0.0092	0.0500	-0.0005	0.0029	0.0550
Angina	0.0033	0.0096	0.0499	0.0031	0.0037	0.0579
Diabetes	0.0029	0.0077	0.0859	0.0020	0.0029	0.0952
Parkinson's	0.0326	0.0337	0.0053	-0.0025	0.0065	0.0060
Alzheimer's	0.0140	0.0177	0.0143	0.0106	0.0131	0.0137
Emphysema	0.0065	0.0089	0.0661	-0.0011	0.0023	0.0743
Specification 2: Self-reported health (Excellent omitted category)						
Very good	-0.0016	0.0052	0.2662	-0.0044 **	0.0020	0.2676
Good	-0.0071	0.0054	0.3193	-0.0042 *	0.0021	0.3204
Fair	-0.0194 **	0.0057	0.1881	-0.0028	0.0022	0.1909
Poor	-0.0262 **	0.0070	0.0848	-0.0014	0.0027	0.0793
Very good*>2003	0.0051	0.0084	0.1060	0.0054	0.0044	0.1242
Good*>2003	0.0221 **	0.0093	0.1251	0.0061 *	0.0043	0.1444
Fair*>2003	0.0325 **	0.0113	0.0722	0.0045	0.0045	0.0848
Poor*>2003	0.0100	0.0138	0.0319	0.0029	0.0050	0.0342
* $p < 0.10$, ** $p < 0.05$						
	Spec 1	Spec 2		Spec 1	Spec 2	
Mean of dependent variable	0.1646	0.1627		0.0175	0.0175	
Observations	78,975	83,100		36,736	36,928	

Data used for this table is from the Medicare Current Beneficiary Survey. Standard errors are clustered at the individual level. Controls for age, gender, and state and metro residence are included. Panel 1 includes the full sample of non-institutionalized respondents. Panel 2 restricts attention only to individuals observed in FFS in one year and followed the next.

Table 9: Effect of MA enrollment and health status on enrollee satisfaction

Dependent var: Satisfaction rating (1-4)	Obs.	OLS coefficient estimates (clustered SEs)	
		Enrolled in MA	MA x Health (demeaned)
Overall medical care	78,003	-0.0375*** (0.00974)	0.0396*** (0.00844)
Out-of-pocket costs	77,819	-0.0400*** (0.0115)	0.0478*** (0.00939)
Follow-up care	77,884	-0.0800*** (0.0144)	-0.00489 (0.0122)
Doctor's concern for your health	77,588	-0.0568*** (0.0107)	0.0230* (0.00946)
Information about your medical condition	77,864	-0.0283** (0.00950)	0.0249** (0.00848)
Access to specialists	77,769	-0.0612*** (0.0139)	0.00947 (0.0117)
Questions answered over phone	77,833	-0.0977*** (0.0189)	0.0345* (0.0154)
Availability of care nights and weekends	77,793	-0.120*** (0.0212)	0.00662 (0.0166)
Medicare care provided in same location	77,820	0.0953*** (0.0150)	0.0514*** (0.0118)
Overall medical care (fixed-effect estimate)	78,003	-0.00503 (0.0356)	0.0252 (0.0160)
Retains coverage type next year (probit coefficients)	44,469	-0.841*** (0.0356)	0.0625* (0.0270)

Notes: Each row represents a regression of the form: $satisfaction\ category_{it} = \beta_1 MA_{it} + \beta_2 MA_{it} \times Health_{it} + \gamma H_{it} + \lambda X_{it} + \epsilon_{it}$, where *satisfaction* takes values from one to four ("very dissatisfied," "dissatisfied," "satisfied," "very satisfied"), *MA* is a dummy variable for being enrolled in Medicare Advantage, *Health* is a (demeaned) linear measure of the five-category self-reported health variable, *H* is a vector of fixed effect for the five health categories, and *X* is a vector of basic controls: age, state-of-residence, and year fixed effects, and indicator variables for being female, disabled, or on Medicaid. As the *Health* variable is demeaned, the coefficient on the *MA* indicator variable represents the effect of being enrolled in MA for an enrollee with average health. A positive coefficient on *MA* x *Health* indicates that the relationship between satisfaction and health status for MA enrollees is greater ("more positive") than that for FFS enrollees.

Figure 1: % of Medicare Recipients in MA / M+C / Medicare Risk Plans: 1985-2010

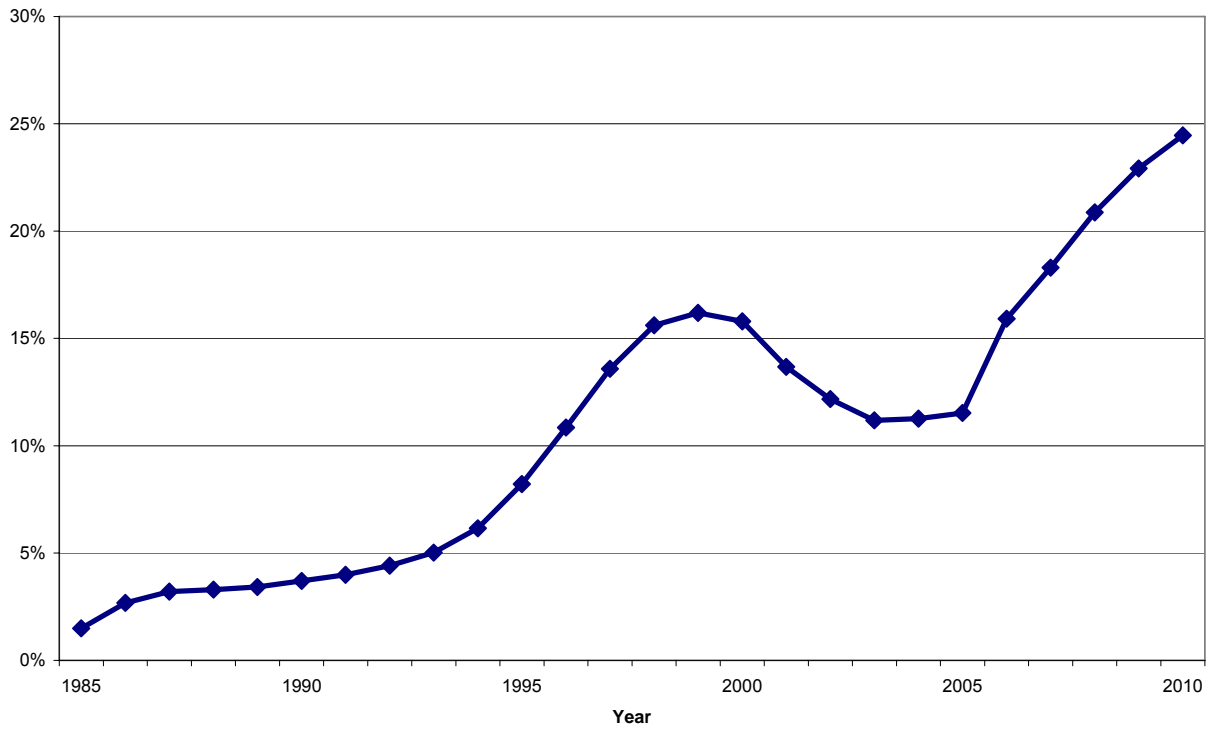
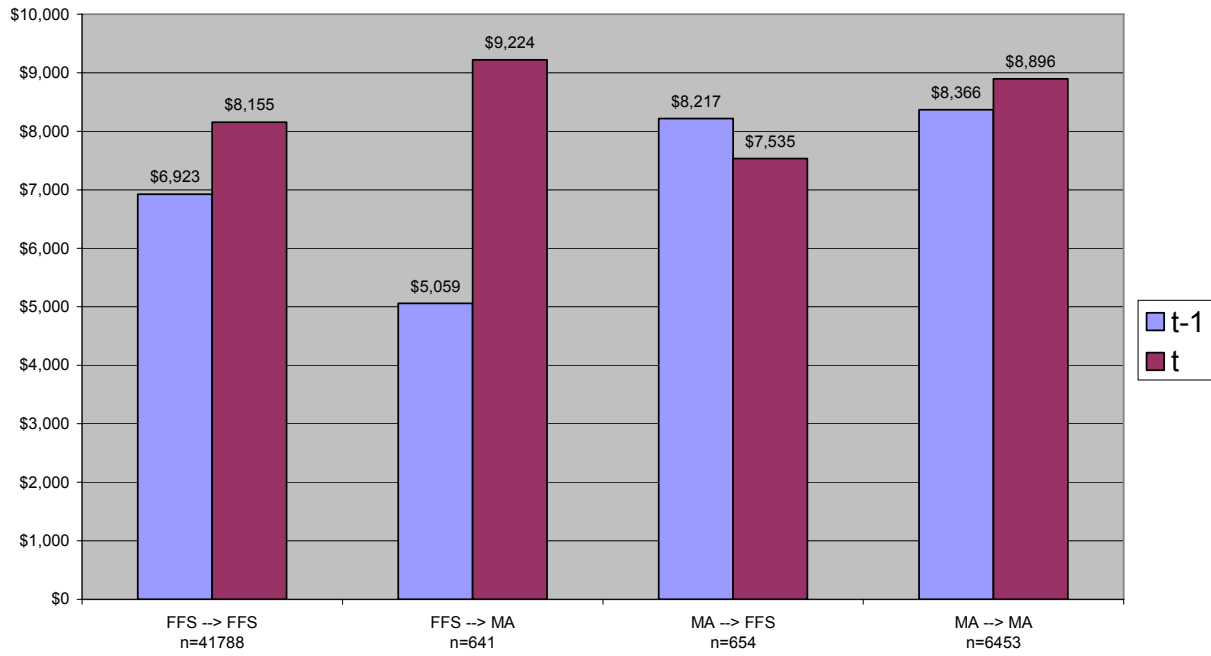
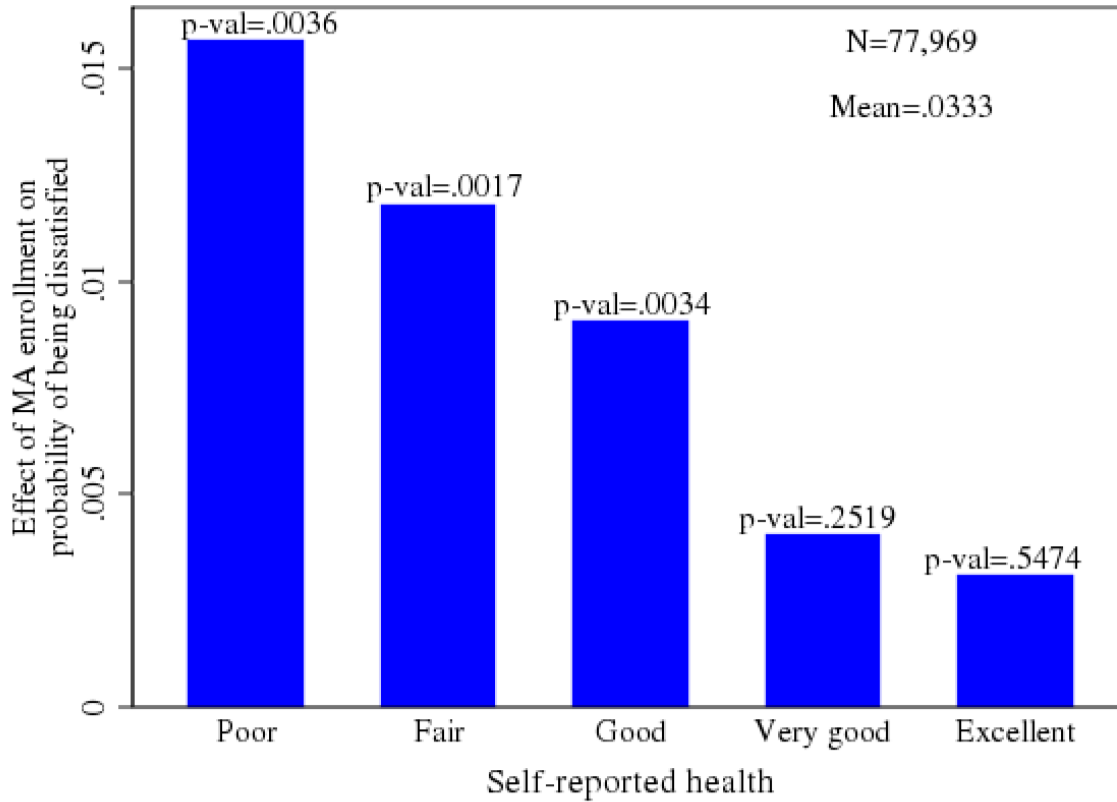


Figure 2: Average Total Cost



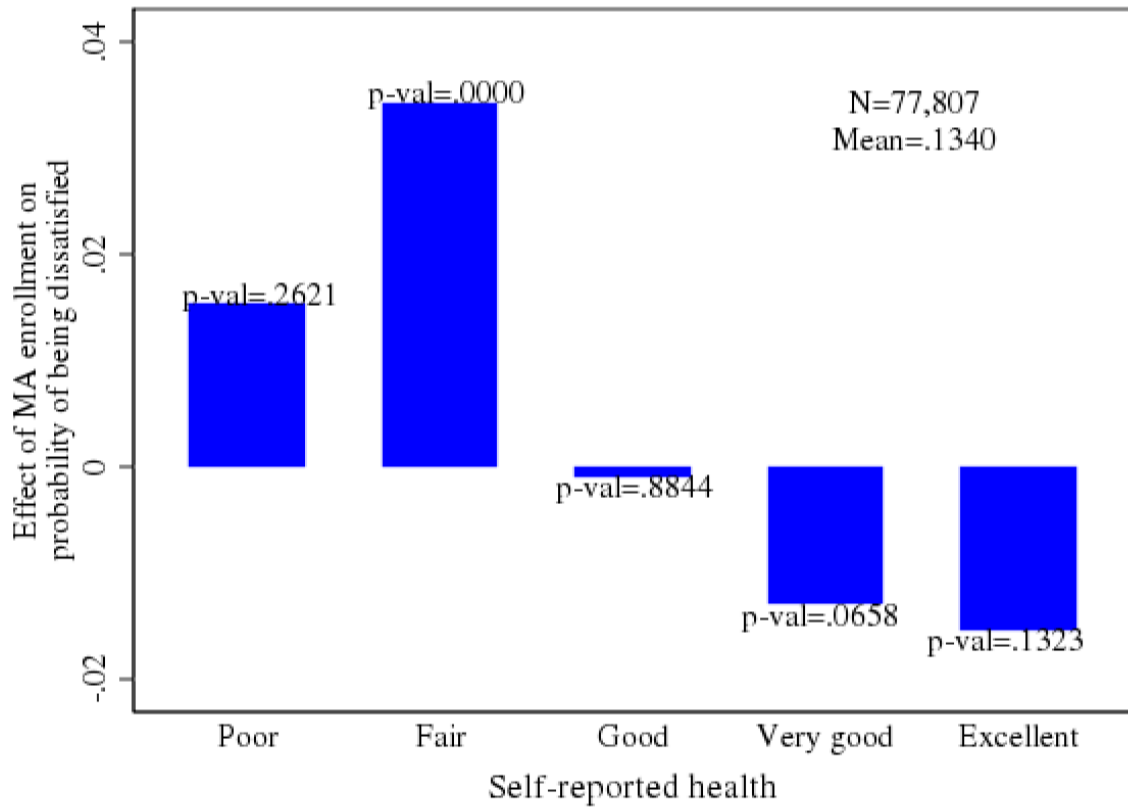
Note: Data is drawn from the Medicare Current Beneficiary Survey (MCBS). Individuals are labeled as MA if they are enrolled in MA for half or more of relevant months. All dollar amounts are adjusted using the CPI to be in 2007 dollars.

Figure 3: Overall dissatisfaction of MA relative to FFS enrollees



s: The graph plots the estimates (as marginal effects) of β_1, \dots, β_5 from the following probit regression: $\text{Pr}(dissatisfied = 1) = \beta_1 MA \times health_1 + \beta_2 MA \times health_2 + \beta_3 MA \times health_3 + \beta_4 MA \times health_4 + \beta_5 MA \times health_5 + \gamma \mathbf{H} + \lambda \mathbf{X} + \epsilon$, where *dissatisfied* is an indicator variable for being “very dissatisfied” or “dissatisfied,” *MA* is a dummy variable for being enrolled in Medicare Advantage, *health*₁, ..., *health*₅ are indicator variables for self-reported health being, respectively, “poor,” “fair,” “good,” “very good,” or “excellent,” \mathbf{H} is a vector of fixed effects for the five health categories, and \mathbf{X} is a vector of basic controls: age, indicator variables for being in Medicare Advantage, female, disabled, or on Medicaid, and fixed effects for year and state of residence.

Figure 4: Dissatisfaction of MA relative to FFS enrollees: out-of-pocket costs



Notes: See Figure 3.

Appendix Table 1. Cross Sectional Summary Statistics

	MA	FFS	Difference
64 and under years old	0.071 (0.257)	0.133 (0.340)	-0.062*** (0.003)
65 - 69 years old	0.182 (0.386)	0.155 (0.361)	0.027*** (0.003)
70 - 74 years old	0.257 (0.437)	0.222 (0.415)	0.036*** (0.004)
75 - 79 years old	0.225 (0.418)	0.200 (0.400)	0.025*** (0.004)
80 - 84 years old	0.147 (0.355)	0.150 (0.357)	-0.003 (0.003)
85 and over years old	0.117 (0.321)	0.140 (0.347)	-0.023*** (0.003)
College Graduate	0.133 (0.339)	0.149 (0.356)	-0.016*** (0.003)
Not a HS Graduate	0.303 (0.460)	0.319 (0.466)	-0.016*** (0.004)
Black	0.103 (0.304)	0.094 (0.292)	0.008*** (0.003)
Hispanic	0.037 (0.190)	0.019 (0.137)	0.018*** (0.001)
Female	0.577 (0.494)	0.569 (0.495)	0.008* (0.005)
Metro Area	0.965 (0.185)	0.718 (0.450)	0.247*** (0.004)
Health Poor or Fair	0.218 (0.413)	0.282 (0.450)	-0.065*** (0.004)
Health Very Good or Excellent	0.453 (0.498)	0.398 (0.489)	0.056*** (0.005)
Total Expenditure	\$8,591 (6516.973)	\$7,878 (17321.540)	712.744*** (147.340)
Sample Size	13,202	77,416	

Note: Unit of observation is a person in a year. Drops all person-year observations in which a person is eligible for only Part A or Part B for any part of the year or if the person is eligible for Medicare only because of having ESRD. An individual in a given year is classified as being on MA if she is on MA for more than half of the months for which she is Medicare eligible in that given year.

For summary statistics, means and standard deviations are reported. For the difference between the two columns, the coefficient and its standard error is reported.

* significant at the 10% level.

** significant at the 5% level.

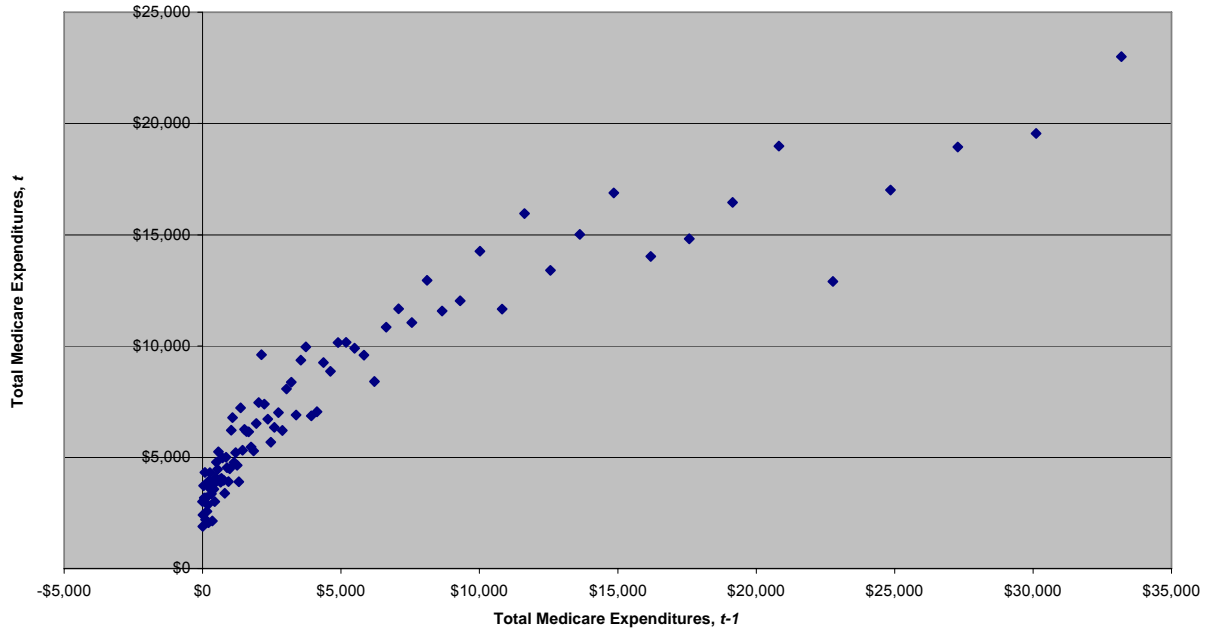
*** significant at the 1% level.

**Appendix Table 2:
Summary Statistics for the Annual Change in % of County's Medicare Recipients in MA Plans**

Year	μ	σ	5th	25th	50th	75th	95th
2001	-0.020	0.038	-0.094	-0.023	-0.008	0.000	0.007
2002	-0.018	0.033	-0.077	-0.026	-0.003	0.000	0.012
2003	-0.008	0.019	-0.033	-0.013	-0.002	0.000	0.010
2004	-0.001	0.011	-0.014	-0.005	0.000	0.002	0.016
2005	0.007	0.013	-0.005	0.000	0.003	0.013	0.031
2006	0.034	0.034	0.000	0.010	0.023	0.053	0.104
2007	0.023	0.026	-0.006	0.006	0.017	0.034	0.071
2008	0.029	0.022	0.007	0.013	0.024	0.039	0.074
Overall	0.006	0.033	-0.040	-0.003	0.002	0.018	0.059

Total number of observations is 3,110 in each year. Summary statistics are weighted by county Medicare enrollment in each year.

Appendix Figure 1: Spending in Years t and $t-1$



Note: Data is drawn from the Medicare Current Beneficiary Survey (MCBS). The sample includes individuals in FFS in both years $t-1$ and t . Individuals are placed into percentiles based on $t-1$ total Medicare expenditures. The mean total Medicare expenditures for individuals in each percentile is recorded on the y-axis. The figure excludes the top 5 percentiles.

Data Appendix: Sources for County-Level Analyses (to be completed)

Medicare county enrollment and MA enrollment: 1997-2005:

http://www.cms.gov/HealthPlanRepFileData/02_SC.asp#TopOfPage

Medicare county enrollment for 2004 - 2007:

<http://www.cms.gov/MedicareEnrpts/>

Medicare MA enrollment for 2006 and 2007:

<http://www.cms.gov/MCRAAdvPartDEnrolData/> (and then click monthly MA state-county-contract)

Medicare county enrollment and MA enrollment for 2008 forward:

<http://www.cms.gov/MCRAAdvPartDEnrolData/> (and then click MA state-county)

Medicare county-level per capita FFS expenditures: 2000-2008:

http://www.cms.gov/MedicareAdvtgSpecRateStats/05_FFS_Data.asp#TopOfPage