Does 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 healthy enrollees. To address this problem, starting in 2004, measures of enrollee health care utilization were used to determine payments to MA plans. Using both individual and aggregate county-level data, we find no evidence that this risk adjustment substantially reduced positive selection into MA plans. We do, however, find that Medicare recipients with mild cases of medical conditions reimbursed under risk adjustment became more likely to enroll in MA plans, suggesting that insurers responded to the policy-induced change in financial incentives. Our results further indicate that Medicare recipients who join MA plans would have cost Medicare $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 spending 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 Advantage[1] (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.[3] 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

[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).

[3] Selection for the Medicare population has been studied in other contexts as well. For example, Fang, Keane, and Silverman (2008) show that patients who enroll in Medigap insurance appear to tbe *positively* selected, with the sources of this selection being driven by income, education, longevity expectations, and financial planning horizons, as well as cognitive ability. Related studies have explored the existence of selection in other private insurance markets, including annuities (Finkelstein and Poterba, 2004) and auto insurance (Chiappori and Salanie, 2000)

total costs to the Medicare program ("Medicare spending").(Langwell and Hadley, 1989; PPRC, 1997; Batata, 2004).

Partly because of these findings, a comprehensive risk-adjustment model was introduced in 2004 that increased the capitation payment for individuals who had specific health conditions in the previous year. As risk-adjustment should better align capitation payments and actual medical costs had the enrollee remained in FFS, its supporters hoped that risk adjustment would 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.

However, the details of the risk adjustment formula potentially created a new avenue by which the capitation payments might poorly reflect actual medical costs. Specifically, the size of the increase in the capitation payment associated with a condition was set to approximately match the average cost of treating this condition for the FFS population. If beneficiaries with more mild versions of a disease enroll in MA plans, the capitation payment would exceed actual medical costs, potentially by a wide margin. Thus, the introduction of risk adjustment on the total costs of the Medicare program is ambiguous.

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 explore the selection patterns into MA plans. Using all Medicare claims for the FFS population in the MCBS, we find strong evidence that individuals who subsequently join MA plans are healthier than individuals who remain in FFS. MA joiners have approximately 30 percent fewer claims than FFS stayers. We find similar evidence of positive selection using an individual's self-reported health and his total Medicare spending while enrolled in FFS.

Our results in this section also highlight the potential challenges that the risk adjustment methodology faces in better aligning capitation payments and actual medical costs. Post risk adjustment, patients with the largest difference between their capitation payment and their actual medical costs are those who have a large number of medical conditions but who have mild cases of each condition. We find evidence that these individuals were more likely to join MA plans after risk adjustment. After the introduction of risk-adjustment, individuals who join MA plans experience a relative increase in the number of medical conditions that they have. However, conditional on having a medical condition, individuals who join MA plans appear to have more mild versions. Taken together, these results suggest that risk adjustment may not have been successful in reducing overpayments to MA plans.

Our second set of results uses the MCBS to directly 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 following individuals over time as they change their MA enrollment status and also controlling for both a rich set of demographic characteristics and 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.

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 Medicare expenditures for those remaining in Medicare FFS, our individual-level results will not capture this. To explore this possibility, in our third 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) and there were no spillovers to the FFS population. We use annual data from 2000 to 2008, which includes 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 to Medicare 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.

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.[4] However, we do find evidence that insurers responded to the policy-induced change in incentives by enrolling Medicare recipients with relatively mild cases of certain health conditions. 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 $110 billion in 2009. 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.

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. We also discuss the changes in the benchmark payments that occurred during this period. In section three we summarize our MCBS data used in our subsequent empirical analyses.

[4] 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.

Section four presents evidence of selection patterns into MA plans by health. In section five, we present our individual-level analyses for the effect of MA enrollment on Medicare expenditures. In section six 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. 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 expected 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.

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. 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 the inherent difficulties in out-of-sample prediction, 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, previous research 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.[5] However, as MA capitation payments were on average 95 percent of average FFS per capita spending during this period, plans were being compensated for insuring the average Medicare recipient, when they were typically insuring recipients far healthier and cheaper than average. 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 1.5 percent.

A more comprehensive risk-adjustment regime was introduced in 2004 based on the hierarchical condition categories (HCC) model. The HCC model uses claims data from the FFS population to predict spending, based on demographic characteristics and diagnosis history for the previous year. To measure diagnoses for individuals switching from FFS into an MA plan, the model uses the diagnosis codes (hereafter ICD-9 codes) on beneficiaries' fee-for-service claims. Each of the more than 15,000 possible ICD-9 codes is assigned to one of 189 condition groups that represent a specific medical condition. Hierarchies are imposed on certain related condition groups, so that a person would only be assigned to

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

the most severe one.[6] The model uses just 70 of these 189 condition groups in determining the capitation payment, as the predictive power of the remaining 119 was deemed insufficient to justify their inclusion. Factors are used for each of these 70 diagnosis groups that increase the capitation rate, and the magnitude of the factor does not depend on the number or severity of these claims.[7] For example, a person with one or more claims with a diagnosis falling in the "specified heart arrhythmias" group is assigned a factor of .368. With fully phased in risk adjustment, the plan would receive an additional 36.8 percent of the benchmark rate compared to an otherwise identical enrollee without the condition (Pope et al., 2004).

Calibrating the HCC model to the FFS population, Pope et al. (2004) show that implementing the risk adjustment model would reduce capitation payments for in the bottom fifth of expenditures by 54 percent and more than double capitation payments for individuals in the top fifth of expenditures. However, the authors also document that the HCC model still explains only 11.2 percent of cost variation among the FFS population.[8] As the share of the capitation payment based on the HCC model rose from 30 percent in 2004 to 50, 75 and 100 percent in 2005, 2006 and 2007, respectively, the actual explanatory power of the final risk score rose steadily from 1.5 percent in 2003 to 11.2 percent in 2007.

While previous literature has examined 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 exploring these same issues during the most recent decade. In the pages that follow, we aim to fill this gap, but before doing so, we discuss the key challenges for this or any similar risk-adjustment model.

One challenge in the Medicare Advantage setting is the need to estimate the HCC model using claims from only the FFS population. This reliance arises because MA plans do not submit claims 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

[6] For example, each ICD-9 code for ischemic heart disease is assigned to condition code 81, 82, 83, or 84. A person with ICD-9 codes that fall in two or more of these groups would only be assigned to the more severe one.

[7] CMS also uses six disease group interactions, reflecting the fact that individuals with, for example, both congestive heart failure and chronic obstructive pulmonary disease have greater expenditures on average than the two individual factors would predict.

[8] 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..

plans' cost data would undermine the guiding principle of avoiding marginal-cost compensation so as to reduce the incentive to provide needless or low-value care. A second challenge is that 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 substantially across individuals, and if private plans tend to enroll individuals with milder cases of the 70 conditions that are accounted for, then overpayments will persist. A final challenge for risk adjustment is that 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.[9]

III. Individual-Level Data and Summary Statistics

A. Individual-Level Data from the MCBS

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.[10] 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 pattern of MA enrollment and its impact on Medicare expenditures 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 data to determine MA plan enrollment in each month. During our study period from 1999 to 2006, there

[9] 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 ().

[10] 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.

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). 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 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 both high school dropouts and college graduates are significantly less likely.

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. Claims Data and Diagnosis Information from the MCBS

As discussed above, CMS currently uses FFS claims data to measure the presence of health conditions among those who subsequently enroll in MA plans. Appendix Table 1 presents summary statistics of the prevalence of Medicare claims in each of the 70 condition groups for a subset of the MCBS population that remains in the sample in the subsequent year. This table shows that nearly one-fourth of the population had at least 1 claim with a diagnosis in the most common category ("diabetes without complication"), and at least 2.0% of the population had at least one claim in 23 of the 70 categories. Many of the conditions, however, are extremely rare; in 31 categories, less than 1.0% of the population had any claims. Appendix Table 2 provides additional information on the average expenditures within each of these 70 HCC categories, restricting attention to only those claims with a primary (as opposed to one of up to eight secondary) diagnosis in the category.[11]

Given that MA plans do not submit claims data to CMS, it is not possible to compare utilization between those in FFS and their counterparts in MA. However, in Table 2, we compare summary statistics from FFS claims data for those who join MA in the next year with their counterparts who remain in FFS. Medicare recipients who subsequently joined MA were less likely to have any HCC-eligible claims, had fewer categories with one or more claims, had a lower overall average number of claims, and had lower average spending on HCC-eligible claims.[12] This table strongly suggests that the baseline utilization of these two groups is very different. We further investigate this issue in the next section.

IV. Evidence on Selection Patterns by Health Status

A. Empirical Strategy

In this section, we more formally investigate the extent to which individuals with lower utilization, and thus on average likely in better health, tend to enroll in MA plans. By differentially increasing the capitation payment for individuals with certain medical 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. Thus, we also explore whether the magnitude of this selection changed with the movement to risk adjustment described above.

To construct our primary measure of health, we rely on the universe of all Medicare claims for beneficiaries while they are in the MCBS. We then aggregate these 4.1 million (everyone) claims using the HCC model and construct a dataset in which the unit of observation is a beneficiary in a year for a condition category. We focus on the 70 condition categories used by the HCC model for two main

[11] These average measures include only claims with a primary diagnosis for the specific condition. In general, this will be substantially lower than the person's overall spending. Furthermore, most claims have more than one diagnosis. While this is accounted for in Appendix Table 3, in this table to avoid double counting we assign the funds only to the primary diagnosis.
[12] All of these differences are statistically significant at the 1% level.

reasons. First, because they form the basis of the risk adjustment formula and hence a plan's capitation payment, studying these condition categories allows us to test for changes in beneficiary characteristics after MA plans' incentives changed. Second, the claims data are administratively determined and are therefore not subject to measurement error arising from recall error on the part of the beneficiary. We estimate the following models on the population of individuals who are enrolled in FFS at time *t*-1:

$$(1) \qquad Y_{jct-1} = \alpha_\tau + \mu * MA_{jt} + \beta * X_{jt} + \rho * H_{jt} + \gamma * METRO_{jt} + \Sigma\lambda_c * I(\text{Condition Groups}_{jct-1} = c) + \varepsilon_{\varphi\tau}$$

in which Y_{jt-1} is an outcome variable for beneficiary *j* in year *t-1* for HCC category c and MA_{jt} is a dummy variable for being enrolled in MA in year *t*. X_{jt} and H_{jt} represent a set of controls for the person's demographic 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 higher average spending in metropolitan areas. Finally, we include fixed effects for each of the 70 condition groups.

To determine if individuals who join MA are healthier than those that remain in FFS, we consider two outcome variables: the total number of claims within a condition category and an indicator for having any positive claims with a condition category. We also consider the number of claims conditional on this number being strictly greater than zero. In each case, an estimate of μ that is less than 0 is consistent with selection of healthier individuals into Medicare Advantage. We also explore whether and to what extent these selection patterns changed over our sample period by interacting the dummy variable for being in MA with a dummy variable for being after the introduction of risk adjustment.

We estimate a companion set of specifications in which the outcome variable is a beneficiary's self-reported history of various chronic and acute illnesses or her self-reported subjective health. For these regressions, the unit of observation is a person in a year. Although these outcome variables are self-reported and do not correspond exactly to the risk adjustment methodology used by the Medicare program, these specifications have one potential advantage. While we only have accurate claims data for individuals while they are on FFS, all respondents in the MCBS were asked a set questions on their histories of several acute and chronic illnesses. We can therefore estimate (1) on the stock of individuals enrolled in MA at any time *t* rather than only individuals who recently joined.

B. Results from Claims-Level Analyses

Table 3 summarizes the main results from our claims-level regressions. Across all panels of the table, specifications (1) and (2) include condition category and year fixed effects, (3) and (4) include condition category by year interactions, and (5) and (6) include individual-level demographic controls. The results from Panel A suggest that individuals who join MA have roughly one third fewer claims than individuals who remain in FFS, providing evidence that healthier and lower cost individuals are more likely to join MA. The effect is highly statistically significant and is stable across specifications. In all six specifications, we cannot reject the null that this difference did not change after the introduction of risk adjustment.

Panel B explores the extent to which these results are robust to considering only condition categories in which an individual has some positive spending in this category in this year. Recall that a MA plan's capitation payment is based on the number and type of condition categories for which a beneficiary has any eligible claims; capitation payments are identical if an individual has one or thirty eligible claims within the same medical condition category. Given this payment system, MA plans have an incentive to attract individuals with a more mild case of a given disease, conditional on having positive claims within that condition category. Panel B presents evidence that is consistent with this prediction. Conditional on having strictly positive claims within a condition category, individuals who join MA have approximately 10% fewer claims than individuals who remain in FFS.

Panel C estimates regressions where the outcome variable is an indicator for having any claims within the condition category. As discussed above, the move to risk adjustment increased the incentive for MA plans to attract individuals with one or more claims in HCC condition categories, because they would receive additional reimbursement. If insurers responded to this incentive, the gap between MA joiners and FFS stayers should shrink after the introduction of risk adjustment. Panel C explores this issue. Consistent with MA joiners on average being healthier than FFS stayers, MA joiners are approximately 21 percent less likely to have one or more claims within a condition category. Strikingly, this gap shrinks by approximately half after the introduction of risk adjustment.

In a companion set of regressions, not reported here, we explore selection using self-reported history of chronic and acute illnesses and self-reported measures of overall health. We first consider if individuals who report having several acute and chronic illnesses are less likely to be enrolled in MA, and if these selection patterns changed after the introduction of risk adjustment.[13] To parallel (1), we estimate this model for the population of beneficiaries who were enrolled in FFS in a baseline year. We also explore the cross-sectional relationship between MA enrollment and health outcomes. In most of these specifications, while MA beneficiaries are less likely to have a condition than are FFS beneficiaries, we fail to reject the null that that this difference is equal to 0. In addition, we generally cannot reject the null that this relationship did not change after the introduction of risk adjustment. One possible explanation for this is that the baseline prevalence of each of these conditions is low, and thus our power to detect a change is low as well. To help circumvent these problems of statistical power, we also studied self-reported health as an outcome variable. These results suggest that even after controlling for a variety of demographic factors, individuals in MA are less likely to report being in fair or poor health. We find little evidence that these patterns changed after the introduction of risk adjustment.

Taken together, our results suggest that the average difference between MA joiners and FFS stayers did not change significantly following the move to risk adjustment. However, our results do suggest that MA plans responded to the change in incentives by enrolling relatively healthy individuals with one or more of the medical conditions reimbursed by the HCC model.

[13] 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.

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

In this section, we use the individual-level MCBS data to estimate the impact of MA enrollment on Medicare expenditures. As summarized in Section III, MA enrollees differ in several observable ways from their counterparts in traditional Medicare. Many of these factors will affect average Medicare spending among MA recipients, and thus we control for them in specifications of the following type:

$$(2) \quad S_{jt} = \alpha_t + \mu * MA_{jt} + \beta * X_{jt} + \rho * H_{jt} + \gamma * METRO_{jt} + \Sigma\lambda_k * I(State_{jt} = k) + \varepsilon_{jt}$$

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. 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. Other control variables are defined as in (1).

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*.

$$(3) \quad S_{jt} = \alpha_t + \mu * MA_{jt} + f(S_{j,t-1}) + \beta * X_{jt} + \rho * H_{jt} + \gamma * METRO_{jt} + \Sigma\lambda_k * I(State_{jt} = k) + \varepsilon_{jt}$$

in which f(.) 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. Second, this methodology estimates the effect of MA enrollment from individuals newly joining MA plans, and thus may 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.

A. 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.[14] 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.

To further explore the pattern of enrollment and disenrollment from MA, Table 4 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

[14] 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.

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. 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 5 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 excludes 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 populations. Consistent with this and as shown in the first panel of Table 5, 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.

The next four panels of Table 5 summarize this same information for four distinct groups: those in traditional Medicare in consecutive years, those who switch from FFS 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 percent in the early period and 15 percent in the late period. The increases for those transitioning into MA plans were much larger at 54 percent and 93 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, average Medicare spending *declines* when individuals switch out of MA plans, by 6 and 15 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 exit 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 5 percent during the early period and 8 percent during the late period for those who remain in MA.

Taken as a whole, these summary statistics suggest that the effect of MA enrollment on Medicare expenditures may actually have increased between the early and late periods. In the section that follows, we explore this issue more rigorously.

B. Regression Results

The difference between average expenditures for beneficiaries on MA and FFS in our sample is \$713, or 9% higher than the average spending for Medicare FFS beneficiaries in our sample. However, there are many important observable differences between the MA and FFS populations. Model (1) in Table 6 shows that controlling for some observable differences between the MA and FFS populations 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 average impact of MA enrollment on Medicare expenditures. For example and as shown in Table 1, MA recipients are much less likely than FFS recipients to report being in fair or poor 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 to \$842. 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 6 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. In this specification, 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 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 mimics 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.[15] Failing to account for this

[15] 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.

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 7.

In Table 7, we summarize the baseline estimates from the switcher regressions and show that, if anything, the estimate in model (5) of Table 6 understates the causal impact of MA enrollment on 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 6] 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 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 6.

Model (3) allows the effect of baseline spending on spending in the subsequent year to differ across Part A (hospital inpatient care) and Part B (physician services) of Medicare. Including this flexibility is potentially important because Part B spending is substantially more persistent that Part A spending and because the ratio of Part A to Part B spending is lower for MA joiners than for FFS stayers.[16] In Model (3), individuals are placed into 21 bins based on Part A spending 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

[16] 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.

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.[17]

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 years *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 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, it is possible that individuals tend to 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 data has little impact on the coefficient of interest. While changing the sample to include only those 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

[17] 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.

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.[18]

We next explore how the estimate has evolved over time with the shift to greater risk-adjustment. Models (1) – (3) of Table 8 present estimates in which we control 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 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 more than twice as large as in the early period.

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 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

[18] 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.

$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.

C. 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 he 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 5 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.[19] In addition, comparing models (2) and (3) of Table 8 shows that controlling for self reported health has little effect on the estimate of interest.

A closely related concern is that even if an individual's health does not deteriorate the year that he joins MA, if an individual endogenously retimes his obtaining of medical services because of the decision to enroll in MA, our estimates will be biased. In particular, suppose that an individual defers care until after he joins MA. Year *t* costs will be artificially low, while year *t+1* costs will be artificially high. Such behavior would bias our results toward 0. Alternatively, if individuals defer care until they join MA plans, our estimates will be biased upward. Although the strategic timing of claims is important in some contexts (Cabral, 2010), three pieces of evidence suggest this is not biasing our estimates. First, unless individuals are retiming care from year *t-1* to year *t+1*, the fact that controlling for *t-1* spending hardly changes our estimates suggests that retiming from *t* to *t+1* may not be important. Second, our satisfaction results (in Section VII below) show that most features of MA plans, including cost sharing,

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

are less attractive to individuals in poor health, suggesting that individuals are unlikely to delay care until after they join MA. Finally, the MCBS contains self-reported measures of health care utilization for services that plausibly might be retained. When we replicate the analysis in Table 7, using changes in utilization of these services, we find no evidence that individuals defer care until they join MA.[20]

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 8 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.

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

To address certain limitations of this individual level analysis, in this section we utilize annual county-level data on per-capita Medicare spending among those in FFS.[21] We focus on elderly Medicare recipients, who account for approximately 85 percent of all program beneficiaries, and consider data for

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[20]Specifically, we determine if individuals have seen a doctor in the previous year, use a hearing aid, and (for men) if they have had a prostate surgery. While all three medical treatments are potentially subject to retiming, we are especially interested in the first two, as vision and dental benefits are frequently major selling points for MA plans (see http://www.ahip.org/content/default.aspx?bc=39|341|321). For this analysis, we create a dummy variable for (1) having an eye exam in the past year, (2) using a hearing aid, and (3) having prostate surgery in the past year. We then create an outcome variable that is the change in this dummy variable from year *t+1* and *t*. We the restrict attention to individuals who are in FFS in year *t-1*. We run 6 separate regressions: two for each outcome variable with and without our full set of controls. The coefficient of interest (fraction of the year on MA) is not statistically significant for the regressions for (2) or (3). For (1), however, the estimated coefficient is small and *negative*, suggesting that individuals are slightly less likely to obtain an eye exam the year they join MA than they were the year they were on FFS.

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

the 2000 through 2008 period. We adjust 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 is \$233. 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. 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. As shown in Appendix Table 4, there is considerable variation across counties in each year with respect to the change in the fraction of Medicare recipients in MA plans.

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:

$$(4)\ \Delta AC_{FFSj,t} = \beta_0 + \beta_1 * \Delta \ln(\%FFS_{j,t}) + \tau_t + \delta_j + \varepsilon_{j,t}$$

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.

The specifications summarized in Table 9 weight each county-year observation by its share of the U.S. Medicare population in that year. Standard errors are clustered at the county level to account for possible serial correlation in the error term. 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 6.6, 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 somewhat 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 early and late part of our study period by interacting our key explanatory variable with a POST-2004 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 that suggested 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.

Furthermore, 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. 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 taking into account, there is no evidence that lagged changes in MA penetration influences average FFS costs.[22]

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.[23]

A. Empirical Strategy

In every year of our sample, the MCBS asks a subsample of enrollees in the survey to rate different aspects of their medical care.[24] We investigate how satisfaction varies by MA enrollment and health in the following equation:

$$(5) \qquad \text{Satisfaction}_{it} = \beta MA_{it} * H_{it} + \gamma MA_{it} + \delta H_{it} + \theta X_{it} + \varepsilon_{it}$$

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

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

[23] 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).

[24] 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.

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 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.

B. Main results

The first nine rows of Table 10 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 10 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.

C. Additional results

One difficulty in interpreting the results in Table 10 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 as specified in the following equation:

(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*x*Health-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 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 10.

These two figures as well as the results in Table 10 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 10 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 little evidence 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 appreciably since the introduction of the most comprehensive model of risk adjustment in 2004.

Using detailed claims data, we also find that individuals who join MA have fewer claims and fewer disease categories than individuals who remain in FFS. This first result does not change significantly following the introduction of risk adjustment. However, we do find evidence suggesting that MA plans responded to the change in financial incentives caused by risk adjustment by enrolling individuals with more reimbursable conditions.

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

	Stay FFS	Join MA
Any HCC-eligible Claims?	0.95	0.88
	(0.22)	(0.32)
Total Number of HCC categories with Claims	2.96	2.45
	(2.50)	(2.18)
Total Number of Claims	60.86	45.04
	(87.69)	(61.04)
Sum of Spending	6,661	4,554
	(15,154)	(9,745)
Sum of Spending in Phy, Inp, and Otp	5,560	3,607
	(13,108)	(7,642)
Sum of Phy Spending	1,945	1,360
	(3,471)	(2,231)
N	41,597	523

Note: the unit of observation is a person X year. The sample is limited to individuals who not in MA for any months in their baseline year and who remain in the MCBS in the following year. These restrictions leave a total of 42,120 person X year observations. Dollar amounts are adjusted to 2007 dollars using the CPI-U.

Table 3: Claims Level Regressions

	Year and HCC FE		Year X HCC FE		Year X HCC FE & Demographic Controls	
	(1)	(2)	(3)	(4)	(5)	(6)
			Panel A: Total Number of Claims			
Join MA?	-0.095***	-0.143***	-0.095***	-0.143***	-0.106***	-0.157***
	(0.0214)	(0.0503)	(0.0214)	(0.0504)	(0.0216)	(0.0494)
Join MA * post		0.061		0.061		0.065
		(0.0556)		(0.0556)		(0.0545)
N			2,948,400			
Mean Dep Var			0.316			
			Panel B: Total Number of Claims | Claims >0			
Join MA?	-1.115	-1.091	-1.023	-1.087	-1.348*	-1.71
	(0.6785)	(1.8584)	(0.6718)	(1.8287)	(0.6919)	(1.7339)
Join MA * post		-0.03		0.08		0.449
		(1.9888)		(1.9611)		(1.8876)
N			87,538			
Mean Dep Var			10.635			
			Panel C: Dummy for Greater than 0 Claims			
Join MA?	-0.006***	-0.011***	-0.006***	-0.011***	-0.006***	-0.011***
	(0.0014)	(0.0027)	(0.0014)	(0.0027)	(0.0014)	(0.0026)
Join MA * post		0.006**		0.006**		0.006*
		(0.0031)		(0.0031)		(0.0030)
N			2,948,400			
Mean Dep Var			0.03			

Note: the unit of observation is a person X year X HCC cat. The universe is people who remain in the sample in the following year and are in FFS in the baseline year.

Note that the “number of claims” and the “total number of claims” is restricted to claims that “count” for risk adjustment: inp, otp, and phy.

The control variables included in the regression are: year and hcc category fixed effects (1) – (2), year X hcc category fixed effects (3) – (4), and year X hcc category fixed effects and demographic controls (5) – (6).

Join MA is a dummy variable for joining MA next year. Join MA * Post is a dummy variable for joining MA * dummy for year>= 2003.

* p<0.10. ** p<0.05. *** p<0.01.

Table 4: Frequency Distribution of Transitions

Baseline Year	Next Year	FFS → FFS	FFS → MA	MA → FFS	MA → MA	Total Remain in Sample	Leave Sample	Total in Sample
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 w hich a person is eligible for only Part A or Part B for any part of the year of 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 w hich she is Medicare eligible in that given year.

Table 5: Summary Statistics by Transition Type

Panel A: FFS → FFS

	Early Years			Late Years		
	t	t+1	Change	t	t+1	Change
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		
	t	t+1	Change	t	t+1	Change
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		
	t	t+1	Change	t	t+1	Change
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		
	t	t+1	Change	t	t+1	Change
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 6: 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 Lst Yr	FFS Lst Yr	FFS Lst 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 7: Switcher Results: Full Period

	(1)	(2)	(3)	(4)	(5)
Fraction of Year on MA	2082***	2186***	2590***	2763***	2740***
	(469)	(463)	(464)	(953)	(940)
Lagged total expenses, $1,000		276***		377***	324***
		(56)		(25)	(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***	817***	849***	610**	507*
	(193)	(189)	(187)	(300)	(297)
Black	2320***	2259***	1879***	2856***	2662***
	(439)	(425)	(407)	(7730	(768)
Hispanic	-281	-264	-395	230	141
	(519)	(494)	(499)	(857)	(828)
Female	-894***	-846***	-818***	-722***	-724***
	(172)	(166)	(163)	(272)	(269)
Mean of DepVar	8,178	8, 178	8,178	8,339	8,339
Spending Control	20 bins. Tot exp	Lin and 20 bins. Tot exp	Lin and 20 bins. Pt A/B	Lin Lst Yr	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 8: 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>=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 9: Relationship between Change in County-Level Per-Capita Medicare Expenditures and Change in Log Fraction of Recipients in FFS

	(1)	(2)	(3)	(4)	(5)	(6)
	Per Capita	Adj Per Capita	Part A Only	Part B Only	Per Capita	Per Capita
Change Log(Fraction FFS)	-1191***	-914***	-648***	-543***	-1267***	-1118***
	(181)	(203)	(956)	(137)	(250)	(217)
Change Log(Fraction FFS) * Year>=2004					166	
					(285)	
Change Log(Fraction FFS) * Year>=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 9: Effect of MA enrollment and health status on enrollee satisfaction

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

Notes: Each row represents a regression of the form: $satisfaction\,category_i = \beta_1 MA_i + \beta_2 MA_i \times Health_i + \gamma \mathbf{H}_i + \lambda \mathbf{X}_i + \epsilon_i$, where $satisfaction$ takes values from one to four ("very dissatisfied," "dissatsifed," "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, $\mathbf{H}$ is a vector of fixed effect for the five health categories, and $\mathbf{X}$ 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 $MA \times Health$ 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

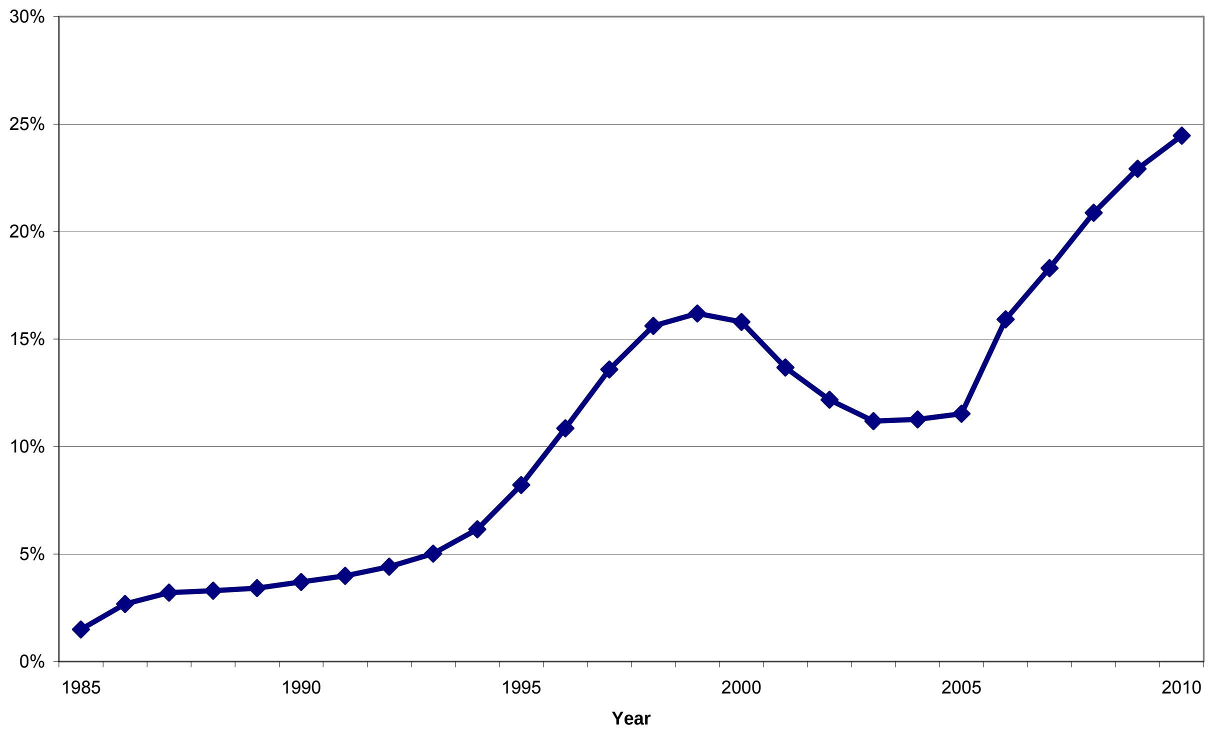
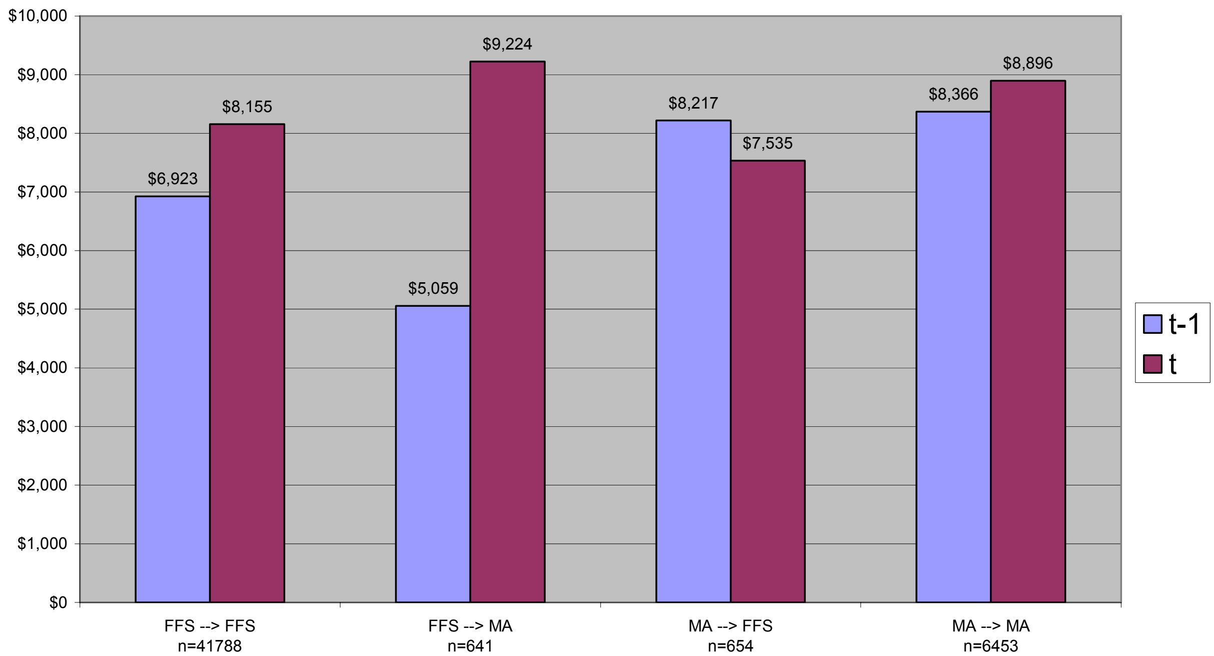
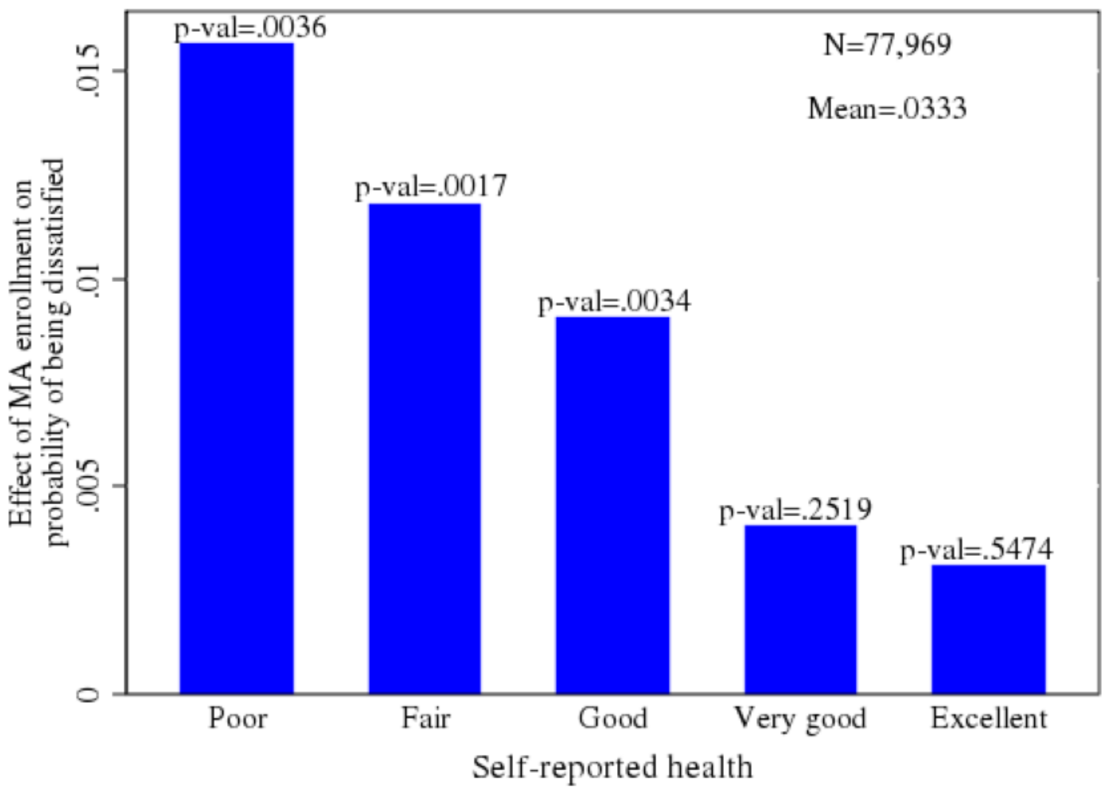


Figure 2: Average Total Cost



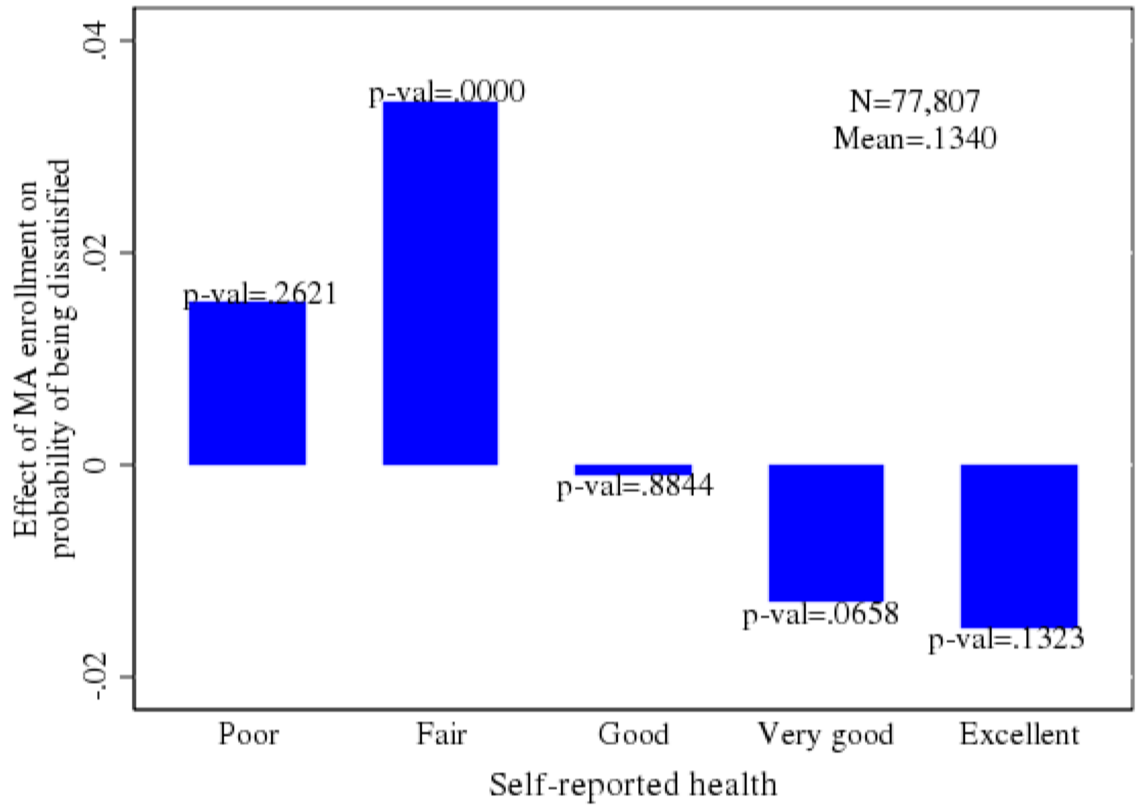
Note: Data is drawn from the Medicare Current Beneficiary Survey (MCBS). Individuals are labeled as MA if they are enrolled in MA for half or more of relavant months. All dollar amounts are adusted using the CPI to be in 2007 dollars.

Figure 3: Overall dissatisfaction of MA relative to FFS enrollees



Notes: The graph plots the estimates (as marginal effects) of $\beta_1, ...\beta_5$ from the following probit regression $Pr(dissatisfied = 1) = \beta_1 MA \times health_1 + \beta_2 MA \times health_2 + \beta_3 MA \times health_3 + \beta_4 MA \times health_4 + \beta_5 MA \times health_5 + \gamma \mathbf{H} + \lambda \mathbf{X} + \epsilon$, 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 fo self-reported health being, respectively, "poor," "fair," "good," "very good," or "excellent," $\mathbf{H}$ is a vector of fixe effects for the five health categories, and $\mathbf{X}$ is a vector of basic controls: age, indicator variables for being in Medicar 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: Summary Statistics For Person X Year X HCC Category: Number of Claims

HCC Category	Percent of People with 1+ Claims	Average Number of Claims Given 1+ Claims	HCC Category	Percent of People with 1+ Claims	Average Number of Claims Given 1+ Claims
Diabetes w.out Complication	24.61%	15.44	*Continued.*		
Chronic Obstructive Pulmonary Disease	15.93%	8.99	Severe Hematological Disorders	1.17%	14.13
Congestive Heart Failure	15.66%	11.69	Bone/Joint/Muscle Infections/Necrosis	1.12%	10.38
Vascular Disease	14.68%	6.45	Disorders of Immunity	1.00%	16.37
Specified Heart Arrhythmias	13.07%	14.51	Aspiration & Specified Bacterial Pneumonias	0.96%	6.48
Breast, Prostate, Colorectal Circ. Manifestation	9.87%	16.49	Artificial Openings for Feeding or Elimination	0.91%	4.09
Angina Pectoris/Old Myocardial Infarction	9.20%	5.45	Major Head Injury	0.89%	4.61
Rheumatoid Arthritis & Inf. Conn. Tissue Disease	5.94%	11.60	Spinal Cord Disorders/Injuries	0.88%	5.22
Ischemic or Unspecified Stroke	5.54%	8.64	Drug/Alcohol Dependence	0.82%	5.76
Renal Failure	4.75%	23.97	Proliferative Diab. Retinopathy & Vitreous Hem.	0.78%	4.66
Unstable Angina & Other Acute Ischemic Hrt Dis.	4.36%	7.54	Inflammatory Bowel Disease	0.72%	6.03
Major Depressive, Bipolar, & Paranoid Disorders	4.33%	9.89	Cirrhosis of Liver	0.70%	10.42
Seizure Disorders & Convulsions	4.24%	9.66	Pneumococcal Pneu., Empyema, Lung Abscess	0.67%	4.38
Polyneuropathy	4.22%	5.29	Diabetes w. Acute Complications	0.61%	8.42
Cardio-Respiratory Failure & Shock	3.91%	8.93	Dialysis Status	0.59%	21.12
Diabetes w. Neurologic/Specified Manifestation	3.77%	6.01	Chronic Hepatitis	0.58%	7.02
Chronic Ulcer of Skin, Except Decubitus	3.55%	8.03	Cerebral Palsy & Other Paralytic Syndromes	0.56%	5.32
Major Complications of Medical Care & Trauma	3.49%	4.96	Cerebral Hemorrhage	0.51%	7.15
Diabetes w. Ophthalmologic/Unspe. Manifestation	3.15%	5.29	Multiple Sclerosis	0.49%	13.25
Diabetes w. Renal or Peripheral Circ. Manifestation	2.93%	8.28	Amputation Status, Lower Limb/Amp. Compli.	0.46%	4.36
Lymphatic, Head & Neck, Brain, & Other	2.25%	26.66	Drug/Alcohol Psychosis	0.43%	3.15
Vascular Disease w. Complications	2.24%	8.28	Quadriplegia, Other Extensive Paralysis	0.35%	6.15
Intestinal Obstruction/Perforation	2.06%	6.23	End-Stage Liver Disease	0.33%	6.14
Hip Fracture/Dislocation	1.78%	14.02	Coma, Brain Compression/Anoxic Damage	0.32%	4.18
Parkinson’s & Huntington’s Diseases	1.77%	9.40	Opportunistic Infections	0.27%	3.40
Schizophrenia	1.71%	13.84	Traumatic Amputation	0.27%	9.14
Hemiplegia/Hemiparesis	1.60%	9.43	HIV/AIDS	0.26%	22.44
Vertebral Fractures w.o Spinal Cord Injury	1.53%	5.74	Paraplegia	0.20%	7.59
Septicemia/Shock	1.52%	9.39	Major Organ Transplant Status	0.19%	13.37
Decubitus Ulcer of Skin	1.51%	7.11	Respirator Dependence/Tracheostomy Status	0.15%	5.14
Lung, U. Digestive Tract, & Other Severe Cancers	1.45%	26.86	Respiratory Arrest	0.14%	6.05
Acute Myocardial Infarction	1.35%	8.74	Muscular Dystrophy	0.08%	7.02
Protein-Calorie Malnutrition	1.35%	6.48	Cystic Fibrosis	0.04%	5.50
Metastatic Cancer & Acute Leukemia	1.28%	17.19	Severe Head Injury	0.01%	4.51
Nephritis	1.21%	4.66	Extensive Third-Degree Burns	0.00%	3.52
Pancreatic Disease	1.18%	8.20			

Note: the unit of observation is a person X HCC category X year. The sample is limited to individuals who not in MA for any months in their baseline year and who remain in the MCBS in the following year. These restrictions leave a total of 42,120 person X year observations. Statistics include all claims (primary, secondary, etc).

Appendix Table 2: Summary Statistics For Person X Year X HCC Category: Claim Amount

HCC Category	Avg. Tot. Spending	Avg. Tot. Spending Given Positive Spending	HCC Category	Avg. Tot. Spending	Avg. Tot. Spending Given Positive Spending
Congestive Heart Failure	246.1	2,475	*Continued.*		
Renal Failure	239.8	8,056	Cerebral Hemorrhage	17.0	4,718
Breast, Prostate, Colorectal Circ. Manifestation	235.9	2,846	Hemiplegia/Hemiparesis	16.2	2,337
Chronic Obstructive Pulmonary Disease	202.5	1,927	Parkinson's & Huntington's Diseases	15.1	1,252
Hip Fracture/Dislocation	152.9	9,437	Artificial Openings for Feeding or Elimination	14.8	1,506
Acute Myocardial Infarction	137.5	12,591	Major Head Injury	14.2	2,177
Specified Heart Arrhythmias	132.9	1,473	Diabetes w. Ophthalmologic/Unspe. Manifestation	10.3	525
Ischemic or Unspecified Stroke	117.0	2,798	Cirrhosis of Liver	9.5	2,635
Diabetes w.out Complication	106.6	560	Pneumococcal Pneu., Empyema, Lung Abscess	9.4	1,942
Major Complications of Medical Care & Trauma	103.1	4,668	HIV/AIDS	9.2	4,261
Vascular Disease	100.2	1,303	Diabetes w. Acute Complications	8.2	1,703
Lymphatic, Head & Neck, Brain, & Other	79.4	4,494	Disorders of Immunity	7.3	1,458
Cardio-Respiratory Failure & Shock	70.8	3,131	Multiple Sclerosis	7.1	1,868
Lung, U. Digestive Tract, & Other Severe Cancers	64.0	5,869	Amputation Status, Lower Limb/Amp. Compli.	6.5	5,110
Major Depressive, Bipolar, & Paranoid Disorders	59.5	1,786	Inflammatory Bowel Disease	5.7	1,178
Vascular Disease w. Complications	55.0	3,649	Spinal Cord Disorders/Injuries	4.5	794
Septicemia/Shock	52.8	5,433	Drug/Alcohol Psychosis	4.3	3,067
Intestinal Obstruction/Perforation	49.2	4,100	Quadriplegia, Other Extensive Paralysis	3.8	1,822
Aspiration & Specified Bacterial Pneumonias	43.6	7,425	Proliferative Diab. Retinopathy & Vitreous Hem.	3.7	732
Schizophrenia	39.8	3,158	Major Organ Transplant Status	3.5	3,412
Decubitus Ulcer of Skin	31.8	2,974	Opportunistic Infections	3.3	2,286
Dialysis Status	28.8	7,788	Drug/Alcohol Dependence	3.1	1,236
Metastatic Cancer & Acute Leukemia	28.6	3,909	Protein-Calorie Malnutrition	3.1	854
Pancreatic Disease	28.2	4,317	End-Stage Liver Disease	2.1	1,386
Diabetes w. Neurologic/Specified Manifestation	27.2	1,287	Cerebral Palsy & Other Paralytic Syndromes	2.1	828
Seizure Disorders & Convulsions	26.9	1,106	Chronic Hepatitis	1.3	461
Rheumatoid Arthritis & Inf. Conn. Tissue Disease	26.6	740	Respirator Dependence/Tracheostomy Status	1.2	1,779
Unstable Angina & Other Acute Ischemic Hrt Dis.	26.3	878	Coma, Brain Compression/Anoxic Damage	1.2	991
Bone/Joint/Muscle Infections/Necrosis	25.0	3,126	Paraplegia	1.2	870
Vertebral Fractures w.o Spinal Cord Injury	24.8	2,705	Nephritis	1.1	342
Chronic Ulcer of Skin, Except Decubitus	21.7	925	Muscular Dystrophy	0.6	1,255
Diabetes w. Renal or Peripheral Circ. Manifestation	21.4	1,249	Respiratory Arrest	0.3	355
Traumatic Amputation	18.9	4,773	Severe Head Injury	0.0	390
Angina Pectoris/Old Myocardial Infarction	18.7	482	Cystic Fibrosis	0.0	88
Severe Hematological Disorders	18.4	3,068	Extensive Third-Degree Burns	0.0	36
Polyneuropathy	18.0	1,071			

Note: the unit of observation is a person X HCC category X year. The sample is limited to individuals who not in MA for any months in their baseline year and who remain in the MCBS in the following year. These restrictions leave a total of 42,120 person X year observations. Statistics include all claims associated with a primary diagnosis. Dollar amounts are

Appendix Table 3. Cross Sectional Summary Statistics

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

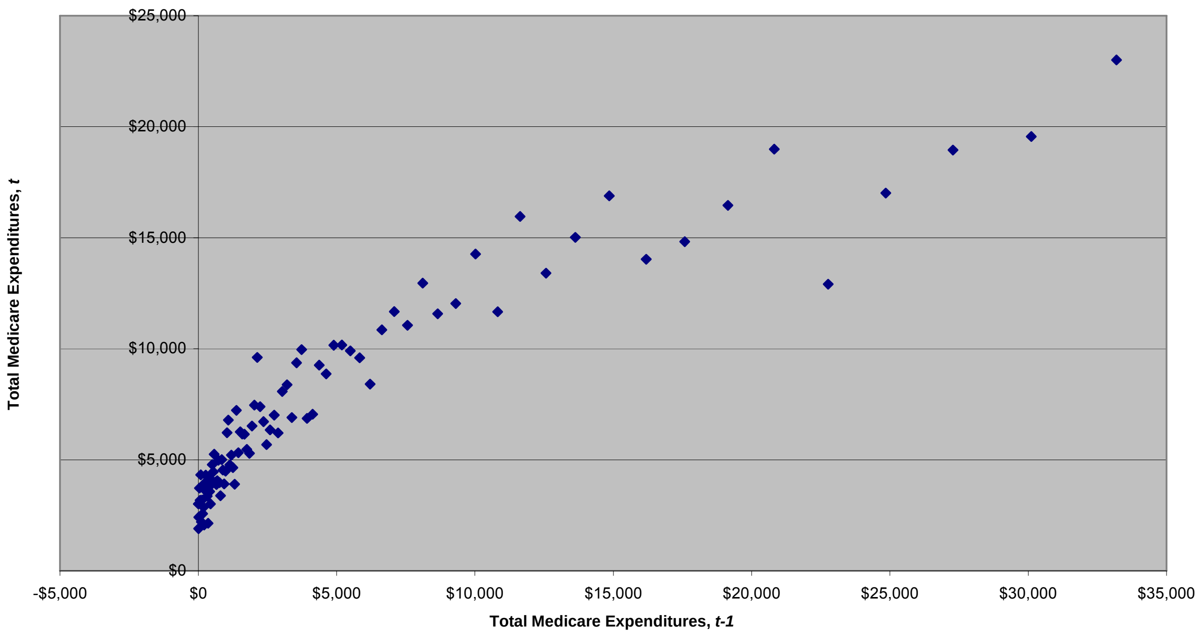
Note: Unit of observation is a person in a year. Drops all person-year observations in w hich a person is eligible for only Part A or Part B for any part of the year of if the person is eligible for Medicare only because of having ESRD. An individual in a given year is classified as being on MA if she is on MA for more than half of the months for w hich she is Medicare eligible in that given year.
For summary statistics, means and standard deviations are reported. For the difference betw een the tw o 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 4:
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/MCRAdvPartDEnrolData/ (and then click monthly MA state-county-contract)

Medicare county enrollment and MA enrollment for 2008 forward:
http://www.cms.gov/MCRAdvPartDEnrolData/ (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