NBER WORKING PAPER SERIES

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Working Paper 24038
http://www.nber.org/papers/w24038

NATIONAL BUREAU OF ECONOMIC RESEARCH
1050 Massachusetts Avenue
Cambridge, MA 02138
November 2017

We would like to thank Bianca Frogner, Doug Conrad, Paul Fishman, and participants of the 2017 AcademyHealth conference and the Program in Health Economics and Outcomes Methodology (PHEnOM) at the University of Washington for their comments. We also thank Sara Walter and Andrew Chu at CMS for assisting us with the access to and clarification of the PBP datasets. We gratefully acknowledge financial support from the Department of Health Services at the University of Washington and the National Institute of Health (R01 AG049815-01A1). The views expressed herein are those of the authors and do not necessarily reflect the views of the National Bureau of Economic Research.

At least one co-author has disclosed a financial relationship of potential relevance for this research. Further information is available online at http://www.nber.org/papers/w24038.ack

Service-level Selection: Strategic Risk Selection in Medicare Advantage in Response to Risk Adjustment
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NBER Working Paper No. 24038
November 2017
JEL No. G22,H51,I13,I18

ABSTRACT

The Centers for Medicare and Medicaid Services (CMS) has phased in the Hierarchical Condition Categories (HCC) risk adjustment model during 2004-2006 to more accurately estimate capitated payments to Medicare Advantage (MA) plans to reflect each beneficiary's health status. However, it is debatable whether the CMS-HCC model has led to strategic evolutions of risk selection. We examine the competing claims and analyze the risk selection behavior of MA plans in response to the CMS-HCC model. We find that the CMS-HCC model reduced the phenomenon that MA plans avoid high-cost beneficiaries in traditional Medicare plans, whereas it led to increased disenrollment of high-cost beneficiaries, conditional on illness severity, from MA plans. We explain this phenomenon in relation to service-level selection. First, we show that MA plans have incentives to effectuate risk selection via service-level selection, by lowering coverage levels for services that are more likely to be used by beneficiaries who could be unprofitable under the CMS-HCC model. Then, we empirically test our theoretical prediction that compared to the pre-implementation period (2001-2003), MA plans have raised copayments disproportionately more for services needed by unprofitable beneficiaries than for other services in the post-implementation period (2007-2009). This induced unprofitable beneficiaries to voluntarily dis-enroll from their MA plans. Further evidence supporting this selection mechanism is that those dissatisfied with out-of-pocket costs were more likely to dis-enroll from MA plans. We estimate that such strategic behavior led MA plans to save $5.2 billion by transferring the costs to the federal government.

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1. Introduction

Recent health care reforms have facilitated the transition from volume- to value-based payment models in hopes of achieving cost control and enhancing the quality of care. One such example is that the Centers for Medicare and Medicaid Services (CMS) reimburses Medicare Advantage (MA) plans with a capitated amount per beneficiary to encourage coordinated care in managed care settings. It has been shown that MA plans save money without sacrificing quality. However, as an unintended consequence, MA plans have selectively enrolled healthier people to receive overpayments, known as favorable selection. It remains inconclusive whether cost savings are attributable to cost-effective care management or to risk selection. To reduce risk selection, CMS has adjusted payments to MA plans to reflect the health status of their enrollees, a process known as risk adjustment (Pope et al. 2000). To more accurately estimate capitated payments, in 2004, CMS introduced a new risk adjustment model—the CMS-Hierarchical Condition Categories (HCC) model—which uses extensive inpatient and outpatient diagnostic information from the prior year to generate risk scores (Pope et al. 2004).

It is debatable whether the CMS-HCC model has been effective in reducing risk selection (Newhouse et al. 2015) or whether it has led to strategic evolutions of risk selection (Brown et al. 2014). On one hand, it has been shown that the CMS-HCC model considerably reduced the phenomenon of avoiding sicker beneficiaries (i.e., those with high-risk scores) in traditional Medicare (TM) plans (McWilliams, Hsu, and Newhouse 2012, Newhouse et al. 2015, Morrisey et al. 2013, Newhouse et al. 2012). As risk adjustment leads to neutral payments for beneficiaries with conditions included in the risk adjustment formula, MA plans no longer have incentive to avoid those with high-risk scores if their conditions are included in the CMS-HCC model.

On the other hand, there is suggestive evidence showing that MA plans could strategically respond to the CMS-HCC model. Brown et al. (2014) argue that the HCC model merely shifted the profitable population from healthy people (i.e., those with low-risk scores) to sick ones who are over-compensated, within their risk-score. This can be achieved because, first, there is considerable variability in actual expenditures of beneficiaries around their risk-adjusted payments. For all beneficiaries with a given health condition, the CMS-HCC model is designed to adjust payments to MA plans by the same rate. However, the severity of the condition and thus the cost of treating it can vary within a given condition (Medicare Payment Advisory

Commission 2012).[1] Second, the variability of the within-risk-score expenditures is larger for those with higher risk scores.[2] Because the CMS-HCC model only accounts for about 100 major conditions, this generates underpayments for those whose conditions are not accurately measured by the model, who tend to have multiple chronic conditions (Frogner et al. 2011).[3]

We examine these competing claims and analyze the risk selection behavior of MA plans in response to the CMS-HCC model. Specifically, we hypothesize that MA plans engage in *service-level selection*, in which they provide relatively lower coverage levels for some services to discourage enrollment of certain beneficiaries. While the CMS-HCC model encouraged MA plans to accept TM beneficiaries with high-risk scores, we claim that MA plans strategically behave to avoid beneficiaries who could be unprofitable under the CMS-HCC model (i.e., those with higher expenditures than their risk-adjusted payments). Although there is a large literature on investigating service-level selection as a risk selection strategy (McGuire et al. 2014, Ellis and McGuire 2007, Ellis, Jiang, and Kuo 2013, Cao and McGuire 2003, Eggleston and Bir 2009, Newhouse et al. 2013, Frank, Glazer, and McGuire 2000), to the best of our knowledge, there is no research explaining mechanisms through which MA plans could engage in service-level selection as a strategic risk selection behavior in response to risk adjustment.

We find that the CMS-HCC model achieved the goal of reducing favorable selection based on health; however, we also find that it led to increased disenrollment of unprofitable beneficiaries from MA plans, leading MA plans to save costs of $5.2 billion in 2007-2009. We explain this phenomenon via service-level selection. Building upon Ellis and McGuire (2007), we theoretically show that MA plans have incentives to effectuate risk selection through service-level selection, as unprofitable beneficiaries are more likely to use services that are expensive and are thus more vulnerable to under provision by MA plans. Specifically, we show that those with higher expenditures than their risk-adjusted payments are more likely to use services that health plans would ration more tightly (i.e., services with higher service-level selection index). This phenomenon is more likely to be pronounced for those with higher risk scores. Then, we

[1] For example, the coefficient for breast, prostate, colorectal and other cancers (HCC10) was estimated to be 0.187 by the CMC-HCC model (Pope et al. 2004). This means that CMS pays 1.187 times higher rates for Medicare beneficiaries with breast, prostate, colorectal and other cancers than those without the condition. Depending on cancer stage, however, their actual expenditures would vary. Specifically, those with stage 4 cancer are more likely to incur higher expenditures than the rate set by CMS, whereas those with stage 1 cancer are more likely to incur lower expenditures than the rate set by CMS.

[2] The variability of total health care expenditures for Medicare beneficiaries with high-risk scores is larger than that for Medicare beneficiaries with low-risk scores because the mean total health care expenditures for the former is higher than the latter.

[3] Frogner et al. (2011) showed that the estimate for breast, prostate, colorectal, and other cancers (HCC 10) was only $1,835 despite its seriousness of the illness. Moreover, it was found that the interaction between chronic kidney failure and congestive heart failure had a statistically significant negative estimate, thereby reducing reimbursements for beneficiaries with these two diseases by $614. However, this is unlikely because having multiple chronic diseases requires more complex care.

find evidence supporting our theoretical prediction that MA plans have actually raised copayments disproportionately more for services with higher service-level selection index (i.e., ambulance, home health service, partial hospitalization, and inpatient hospital service) than services with lower service-level selection index (i.e., outpatient substance abuse services, outpatient X-rays, and outpatient hospital services). Such disproportionate increases in copayment induced unprofitable beneficiaries to voluntarily disenroll from MA plans. In additional analyses, we find evidence supporting this selection mechanism that those with dissatisfaction with out-of-pocket costs were more likely to disenroll from MA plans. Consequently, the variation of total Medicare expenditures for MA enrollees with high-risk scores reduced over time. These findings indicate that service-level selection allowed MA plans to avoid the risk of enrolling unprofitable beneficiaries.

2. The Medicare Advantage Program

Medicare beneficiaries can choose to either enroll in a TM plan or an MA plan. Under the Balanced Budget Act of 1997, private plans have contracted with CMS to provide the elderly Medicare Parts A and B benefits.[4] When individuals become eligible for Medicare, they are assigned to TM by default. Then, they can choose to stay in TM or switch to a MA plan, depending on their preferences and needs. Because MA plans offer more generous benefits and lower cost-sharing than TM plans, beneficiaries may prefer enrolling in the MA plan. In contrast, as MA plans have limited providers' networks (Jacobson et al. 2016), those with complex diseases may prefer TM's freedom of provider choice. To encourage beneficiaries to choose the plan that efficiently provides them care while accounting for their individual preferences, CMS adopted the "payment neutrality" approach, which sets MA payments equal to the average Medicare expenditures of TM beneficiaries in the MA enrollee's county (Medicare Payment Advisory Commission 2014).

Over several decades, policy-makers have promoted managed care in Medicare as a way to improve quality of care while containing costs. As this approach creates incentives to encourage preventive care and better care coordination, it is especially helpful in caring for Medicare beneficiaries, 68.4 percent of whom had two or more chronic conditions and 36.4

[4] There are two rationales for privatization of managed care. First, capitated payments to private plans would incentivize them to actively manage their enrollees' care, leading to more efficient care provision. Also, privatization could lead to competition among private plans as well as TM plans, possibly lowering health care costs while improving quality of care.

percent had four or more chronic conditions (Lochner and Cox 2013). In line with the encouragement of managed care, the benchmark levels have been increased to encourage plans to enter in the MA market (Medicare Payment Advisory Commission 2014).[5] Consequently, CMS paid $170 billion to MA plans on behalf of 16 million beneficiaries, reaching a historic high of MA's penetration of almost 31 percent of the Medicare population (Kaiser Family Foundation 2015).

However, benefit designs by MA plans may affect whether beneficiaries choose TM or MA plans'. TM plans under the fee-for-service (FFS) payment system are paid for each test and procedure, and thus theoretically they have no incentive to selectively accept beneficiaries with certain characteristics. In contrast, MA plans under the capitated payment system are paid a fixed amount per beneficiary, which would create an incentive to selectively accept healthier people and avoid sicker ones. To effectuate favorable selection, MA plans could vary benefit designs. For example, MA plans could increase cost-sharing for certain services to avoid unprofitable beneficiaries. This is plausible because while MA plans must provide the same services covered by TM plans (i.e., Medicare Parts A and B benefits), and the actuarial value of the total benefits package must at least be equivalent to TM's benefits, cost-sharing for any particular service can vary between MA plans and TM. Moreover, MA plans could restrict physician networks to distract unprofitable beneficiaries (Jacobson et al. 2016) or offer additional services (e.g., dental care and vision care services) to attract profitable beneficiaries.

To mitigate favorable selection of MA plans, CMS has used a risk-adjusted payment methodology to estimate capitated payments to MA plans. Payment rates to MA plans are determined by enrollee's risk scores, county-level benchmarks set by CMS, and plan bids.[6] However, the fundamental goal of risk adjustment is to adjust payments to accurately reflect the health status of each enrollee. CMS adjusts payments to MA plans, using risk scores estimated based on each beneficiary's demographics and diagnoses in a prior year. However, its ultimate goal is not accuracy per se, but rather improved incentives (Glazer and McGuire 2000, Van de

[5] Prior to the Balanced Budget Act of 1997, MA plans were paid based on 95 percent of a county's average Medicare expenditures of TM beneficiaries. In the Balanced Budget Act, Congress increased benchmark levels to encourage plans to enter in the MA market. Consequently, as of 2015, MA plans are paid by 102 percent of TM costs.
[6] Since 2006, CMS has implemented a competitive bidding system to determine payments to MA plans. The payment is based on bids submitted by MA plans, and then it is risk-adjusted by the CMS-HCC model. The plan bids for Parts A and B services are compared to the county-level benchmark. If the plan bid is less than the benchmark, the plan receives its bid. In addition, CMS retains 25 percent of the difference between its payment benchmark and bid. The remaining 75 percent of the difference must be returned to enrollees in the form of additional benefits or lower premium. If the plan' bid is higher than the benchmark, enrollees pay the difference in the form of a monthly premium in addition to the Medicare Part B premium.

Ven and Ellis 2000). As such, risk adjustment intends to disincentivize health plans from selectively enrolling and caring for healthy beneficiaries, and furthermore incentivize the plans to compete based on providing high-value care. Up until 2000, the risk adjustment process only accounted for age, gender, Medicaid eligibility, institutional status, and county of residence. Starting in 2000, CMS began to use information on inpatient diagnoses to adjust payments to MA plans through the Principal Inpatient Diagnostic Cost Group (PIP-DCG) model (Pope et al. 2000).[7] To more accurately estimate capitated payments, in 2004, CMS introduced a new risk adjustment model—the CMS-HCC model—which uses extensive inpatient and outpatient diagnostic information from the prior year to generate risk scores (Pope et al. 2004).[8]

3. Previous Literature

The literature on risk selection in the MA program has mainly focused on one aspect of risk selection: whether TM beneficiaries with high-risk scores were less likely to enroll in MA plans. Some studies found that such selection behavior was greatly reduced after the full phase-in of the CMS-HCC model starting January 1, 2007. Using a 20 percent random sample of Medicare claims in 2003-2008, Newhouse et al. (2012) found that differences in predicted expenditure between TM-to-MA switchers (i.e., Medicare beneficiaries who enrolled in TM and switched from TM to MA next year) and TM stayers (i.e., those who enrolled in TM and remained in TM next year) declined between 2004 and 2008 by a factor of three. Also, differences in adjusted mortality rates between these two groups narrowed between 1998 and 2008 by a factor of two. Using the 2001-2007 Medicare Current Beneficiary Survey (MCBS), McWilliams, Hsu, and Newhouse (2012) observed that differences in health care use and self-reported health between all TM and MA beneficiaries were narrowed from 2001-2003 to 2006-2007. They also found that differences between TM-to-MA switchers and TM stayers were narrowed. Using a five percent random sample of Medicare claims in 1999-2008, Morrisey et al. (2013) showed that the implementation of the CMS-HCC model led to increase the number of

[7] The PIP-DCG model accounts for 24 age/sex cells, interactions of Medicaid status and age/sex cells, interactions of originally disabled status and age/sex cells, working-aged status, and the 16 PIP-DCG diagnostic categories (Pope et al. 2000).
[8] The CMS-HCC model accounts for 24 age/sex cells, interactions of Medicaid status and sex and age/disabled entitlement status, interactions of originally disabled status and sex, 70 HCC diagnostic categories, interactions of diagnostic categories with entitlement by disability, and six disease interactions (Pope et al. 2004). For HCC diagnostic categories, tens of thousands of the *International Classification of Diseases* (ICD)-9 codes are grouped into a small number of organized categories to generate a diagnostic profile of each person. Thus, each HCC diagnostic category includes conditions that are clinically related to each other and have similar cost implications.

new MA enrollees and decrease the number of MA disenrollees (i.e., those who enrolled in MA and switched from MA to TM next year), resulting in increased MA enrollment.

Another strand of the literature has examined on another aspect of risk selection: whether MA enrollees with high-risk scores were more likely to disenroll from MA plans. There is evidence showing that MA plans might change the targeted population for risk selection in response to the CMS-HCC model. McWilliams, Hsu, and Newhouse (2012) found that compared to MA stayers (i.e., those who enrolled in MA and remained in MA next year) or TM stayers, MA-to-TM switchers (i.e., those who enrolled in MA and switched from MA to TM next year) self-reported poorer health and used more health care after the full phase-in of the CMS-HCC model. Morrisey et al. (2013) observed that after the full implementation, disenrollment from MA plans was more pronounced among the high-expenditure beneficiaries. Using a five percent sample of Medicare claims between 2006-2011, Jacobson, Neuman, and Damico (2015) found that Medicaid-eligible beneficiaries and beneficiaries younger than 65 years with disabilities were more likely to disenroll from MA plans. On the other hand, the low-expenditure beneficiaries were more likely to stay in MA plans and relatively younger beneficiaries aged 65 to 69 years were more likely to switch from TM plans to MA plans.

However, very little is known about the mechanism in which MA plans might engage in risk selection beyond risk scores. To the best our knowledge, there is one paper that argued that MA plans might strategically behave in response to the CMS-HCC model by risk-selecting based on expenditures conditional on risk scores (Brown et al. 2014). Specifically, Brown et al. (2014) found that the CMS-HCC model reduced the phenomenon of avoiding high-risk score beneficiaries whose conditions are included in the CMS-HCC model because it increases payments for them by accounting for additional information from outpatient claims. However, the CMS-HCC model still generates underpayments for some of those with high-risk scores as their complex health status cannot be accurately captured by the model (Frogner et al. 2011). As such, Brown et al. (2014) found that in response to the CMS-HCC model, MA plans have selectively enrolled beneficiaries with high-risk scores but low expenditures conditional on their risk scores. This indicates that the CMS-HCC model cannot reflect the health status of beneficiaries beyond health status captured through claims data (i.e., risk scores), creating an incentive for MA plans to avoid those whose health status could be worse than estimated.

Newhouse et al. (2015) re-examined the question of how MA plans have responded to the CMS-HCC model and found that the combination of the CMS-HCC model and a lock-in policy[9] largely reduced favorable selection after 2007, the year in which the CMS-HCC model was fully phased in. Since the full implementation of the CMS-HCC model coincided with the introduction of the lock-in policy, they acknowledged that it would be hard to distinguish the effect of the CMS-HCC model from the effect of the lock-in policy on reducing risk selection. It is worthwhile to note that Newhouse et al. (2015) intended to validate findings from a prior study of Brown et al. (2014). Using the 1994-2006 MCBS, Brown et al. (2014) found evidence showing that the CMS-HCC model did not reduce favorable selection due to MA plan's strategic response to the CMS-HCC model.[10] However, Newhouse et al. (2015) used a 20 percent random sample of Medicare claims between 2001-2011 and found that the combination of the CMS-HCC model and the lock-in policy reduced overpayments attributable to selection by roughly a factor of five (from $1,984 in 2001-2002 to $320 in 2007-2011), thereby rebutting Brown et al. (2014)'s claim.[11] However, it remains unanswered whether MA plans have changed risk selection strategies in response to the CMS-HCC model to selectively disenroll those with high-risk scores but high expenditures conditional on their risk scores. Determining this aspect is critical to comprehensively understand the strategic risk selection behaviors of MA plans in response to the CMS-HCC model.

4. Service-level Selection

By law, MA plans are not allowed to deny coverage based on beneficiaries' health status. However, MA plans might practice risk selection in subtle ways so that unprofitable

[9] Beginning in 2006, CMS imposed a partial enrollment lock-in to prevent MA enrollees from switching from MA plans to TM plans monthly, limiting temporarily switches to TM plans when more generous coverage or freedom of provider choice is desired.

[10] Brown et al. (2014) concluded that favorable selection was not decreased on net because the decrease in selection along dimensions included in the formula of the CMS-HCC model was more than offset by the increase in selection conditional on the risk score. Specifically, they found that there were smaller differences in risk scores between TM-to-MA switchers and TM stayers after its initial phase-in starting January 1, 2004. This result is consistent with those of the studies showing the effect of the CMS-HCC model on reducing risk selection (McWilliams, Hsu, and Newhouse 2012, Newhouse et al. 2012, Morrisey et al. 2013). However, they showed that compared to TM stayers, actual expenditures conditional on the risk score of TM-to-MA switchers substantially fell after the initial phase-in period. This suggests that MA plans might strategically behave to enroll those with expenditures lower than what is predicted by the CMS-HCC model.

[11] Such contradictory results are attributable to data and methodology differences. First, due to a relatively small sample in the MCBS, Brown et al. (2014) had to pool the years from 1994 to 2002 and then compared selection in those years with selection during the pooled years from 2004 to 2006. This might be problematic because MA reimbursement policy changed markedly during the period. For example, the Balanced Budget Act of 1997 established floors on reimbursement to MA plans for low-paying areas and restricted annual increases in reimbursement to MA plans for high-paying areas to two percent. In contrast, Newhouse et al. (2015) used the sample for the pre-implementation period from 2001 to 2003 and compared selection during the pre-implementation period with selection during the post-implementation period. Moreover, using a large sample size, the study estimated the degree of selection in each year, allowing them to control for various Medicare payment polices across years. Also, Brown et al. (2014) included all MA enrollees from the MCBS. However, starting in 2004, CMS allowed MA plans to create plans for enrollees with special needs (e.g., institutionalized or Medicaid-eligible enrollees), many of whom are non-elderly. As comparing those groups before and after 2004 is problematic, Newhouse et al. (2015) limited to MA enrollees who were elderly, who were not institutionalized, and who were not eligible for Medicaid.

beneficiaries voluntarily disenroll from MA plans. To achieve such risk selection, MA plans could risk-select through collecting additional data or advertising. However, because such selection mechanisms would lead to substantial screening costs, MA plans are likely to seek screening approaches with lowest costs. One such approach that generates relatively low screening costs is service-level selection because MA plans do not need to predict each beneficiary's expenditures but rather only need to predict services more likely used by beneficiaries who could be unprofitable under the CMS-HCC model.

Service-level selection is one type of risk selection, which is based on the phenomenon that unprofitable individuals are more likely to use services that are expensive to health plans subject to capitated payments and are thus more vulnerable to under-provision by health plans. As with risk selection, service-level selection occurs due to asymmetric information between two parties, in which health plans do not know individuals' private information about health status and preferences for health care.[12] The health plan only knows the probability of using the service at the population level, while the individual knows her need, or probability of need, for each health care service and chooses the best health plan that can satisfy her need (Frank, Glazer, and McGuire 2000, Ellis and McGuire 2007). Since rational individuals respond to health plan design when selecting plans, reducing coverage levels for services related to financial losses (i.e., services more likely used by unprofitable individuals) would induce unprofitable individuals to voluntarily disenroll from the plan. In this way, service-level selection would allow health plans to reduce the scope of enrolling those who could be costly to them. Although unprofitable individuals enroll in the plan, service-level selection would also enable health plans to reduce their financial loss as they shift the costs to the individual.

This paper builds on the existing theoretical and empirical work on service-level selection. Frank, Glazer, and McGuire (2000) characterized plans' rationing as a shadow price on access to various types of care, and then showed how a health plan chooses the profit-maximizing shadow price for each service. A shadow price is regarded as a device to capture various rationing strategies by a plan, which determines access to care. For example, the shadow price reflects plan decisions about capacity in various service areas as well as the makeup of

[12] Even with symmetric information, health plans can have incentives for risk selection if they are not allowed to use the private information to set premiums or benefit features (Van de Ven and Ellis 2000). In the MA program, for example, CMS reimburses MA plans based on the costs predicted by the CMS-HCC model that only partially accounts for clinically significant medical conditions with significant costs. Because the model does not incorporate all diagnoses, payments to MA plans are too low for sicker beneficiaries and too high for healthier beneficiaries. Consequently, the imperfect risk-adjustment model can create incentives for MA plans to engage in inefficient sorting of individuals across health plans and distortion of plan benefits through service-level selection.

networks or payment to providers. Building upon Frank, Glazer, and McGuire (2000), Ellis and McGuire (2007) derived a service-level selection index measuring the plan's incentives to ration. Using Medicare claims data for 1996-1997, they measured the relative magnitude of potential selection across various types of services. For instance, hospice, home health care, durable medical equipment, provider specialties of pulmonary care, oncology ambulance, and psychiatry were shown to have potential for under-coverage by managed care plans. On the other hand, eye procedures, magnetic resonance imaging (MRI), and provider specialties such as chiropractic and gynecology were found to be candidates for over-coverage. Similar patterns of health plans incentives for service-level selection were found in other studies, for example, Cao and McGuire (2003) in Medicare, Eggleston and Bir (2009) in the state employee insurance program, and Ellis, Jiang, and Kuo (2013) in commercial health plans.

There is suggestive evidence showing that service-level selection occurs in the MA program. Newhouse et al. (2013) estimated margins (i.e., the ratio of average revenue to average cost) across 48 HCCs and unique combinations of HCCs from data on the cost of care from two MA-health maintenance organization (HMO) plans. Despite no evidence of selection across HCCs, they showed that margins in the two plans varied greatly across HCCs. Two additional studies examine switching behavior for those with need for costly services such as nursing home and home health care.[13] Rahman et al. (2015) found that a high proportion of MA enrollees with need for nursing home or home health care disenrolled from MA plans the next year. Similarly, Goldberg et al. (2016) showed that the switching rate for MA and TM beneficiaries without a nursing home stay was the same, but those who required nursing home services in the prior year were more likely to disenroll from MA plans. This phenomenon was more prominent for those with the most costly, longest nursing home stays.

However, there is no examination of the mechanisms through which MA plans could engage in service-level selection against risk adjustment. This study focuses on cost-sharing as a way to effectuate service-level selection. While cost-sharing is designed to protect people against financial risk, it also affects incentives to use more or less health care services. In the presence of low cost-sharing, an individual may use health care services more because she pays less for care than it costs, which is known as moral hazard in health insurance. Health plans may exploit the

[13] Implications from Rahman et al. (2015) and Goldberg et al. (2016) may be limited due to different characteristics of the study population. Compared to Medicare beneficiaries who are eligible only for Medicare, those with need for nursing home can use Medicaid-covered services in addition to Medicare-covered services and can enroll in or exit MA plans at any time.

mechanism designed to prevent moral hazard in order to engage in risk selection. Implementing service-level selection through cost-sharing would likely be effective, as quantitative studies found that cost was an important consideration most MA enrollees switching to TM. Specifically, Government Accountability Office (2017) showed that cost-related concerns were a leading reason for disenrollment of those with poor health as well as those with better health. Moreover, McCormack et al. (2005) found that cost-related concerns were combined with other reasons, amplifying the likelihood of disenrollment from MA plans.

In this study, we present how MA plans effectuate service-level selection as a strategic risk selection mechanism in response to the CMS-HCC model. First, we demonstrate that Medicare beneficiaries who are expected to incur higher expenditures than their risk-adjusted payments estimated by the CMS-HCC model are likely to use services that are expensive, which are more vulnerable to under-provision by MA plans. If this pattern is widespread across CMS-HCC-levels, then MA plans have strong incentives to engage in service-level selection. To avoid those with significantly higher expenditures than their risk-adjusted payments, then MA plans are likely to increase enrollees' cost-sharing more for services that appeal to them than other services after the full phase-in of the CMS-HCC model. If such disproportionate increases in cost-sharing are large enough to affect individuals' plan choice, it would induce those with significantly higher expenditures than their risk-adjusted payments to voluntarily disenroll from MA plans. Consequently, the variation of total health care expenditures for MA enrollees would decrease after the full phase-in period. This effect would be more pronounced for those with high-risk scores than those with low-risk scores.

5. Theoretical Predictions

5.1. Existing Model on Service-level Selection

We build upon the two prior theoretical models on service-level selection: Frank, Glazer, and McGuire (2000) and Ellis and McGuire (2007). A health plan's profit is revenue less costs. Health plans' revenues from individual i, rev_i, typically comprise a risk-adjusted (capitated) payment. Following Ellis and McGuire (2007), we assume that a premium that the plan charges has been predetermined and thus does not influence plans' strategies to effectuate risk selection. On the other hand, the plan incurs costs for providing services. Frank, Glazer, and McGuire (2000) characterized plans' rationing as a shadow price on access to various types of care. From the perspective of an individual, this can be interpreted as a threshold of clinical need or benefit that the individual must exceed to receive services. As such, a higher shadow price means tighter rationing. For service s, the plan sets a shadow price to ration the service. Let $q = \{q_s\}$ be a vector of shadow prices determined by the plan to ration services and $m_i(q) = \{m_{is}(q_s)\}$ be the vector of expenditure on service s that individual i spends as a function of the service-specific shadow price. The level of expenditure that individual i spends on service s, $m_{is}(q_s)$, is determined by the point at which the marginal benefit of expenditure for that individual is equal to the shadow price q_s. Therefore, the plan's profit for individual i can be expressed as $rev_i - \sum_s m_{is}(q_s)$.

The plan's total profit depends on who joins. Whether an individual joins the plan is determined by her expectation of what she would receive in the plan. Let $\widehat{m}_{is}(q_s)$ be the services that individual i expects to receive in a plan that rations using service-specific shadow prices q_s. From the perspective of the plan, individual i enrolls in the plan with a probability $n_i(\widehat{m}_{is}(q_s))$ as a function of shadow prices. Therefore, the plan's total profit can be expressed as:

$$\pi(q) = \sum_i n_i(\widehat{m}_{is}(q_s)) \left[rev_i - \sum_s m_{is}(q_s) \right] \tag{1}$$

The plan chooses each q_s to maximize expected profits in the equation (1). To find profit-maximizing values of each q_s, the equation (1) is differentiated with respect to q_s.

By differentiating the equation (1) from Frank, Glazer, and McGuire (2000), Ellis and McGuire (2007) derived the service-level selection index, which measures the plans' incentives to ration care tightly across services, I_s:

$$I_s = \sigma_\pi \phi \eta_s \left[\frac{\sigma_{\widehat{m}_s}}{\overline{m}_s} \rho_{\widehat{m}_s,\pi} - C \right] \tag{2}$$

where π is the plan's net profits, σ_π is the standard deviation of π, ϕ is a uniform enrollment function and is constant across service s, η_s is the demand elasticity for service s (a negative number), $\widehat{m}_s$ is the individual's expected expenditure on service s, $\sigma_{\widehat{m}_s}$ is the standard deviation of $\widehat{m}_s$, $\overline{m}_s$ is the mean level of expected spending on service s, $\rho_{\widehat{m}_s,\pi}$ is the correlation between $\widehat{m}_s$ and π, and C is a numeric constant to capture terms that do not depend on service s.

The service-level selection index, I_s, measures the relative magnitude of selection incentives across services. It consists of three components: 1) the coefficient of variation of the expected expenditure ($\widehat{m}_s$) on service s (*predictability*), $\frac{\sigma_{\widehat{m}_s}}{\bar{m}_s}$, 2) the correlation between the expected expenditure on service s ($\widehat{m}_s$) and net profits (π) (*predictiveness*), $\rho_{\widehat{m}_s,\pi}$, and 3) the demand elasticity for service s, η_s. In this study, we focus on the first two components.[14]

First, *predictability* represents how well individuals can predict service-level use. If individuals cannot predict service-level use well, service-level selection would have little to no effect on enrollment or plan profits. If individuals cannot predict service-level use at all (i.e., everyone expects themselves to be average users), predictability is zero. Selective rationing of health plans would not affect individual's plan choices, and no distortion occurs. When individuals can predict their service-level use, expected expenditures ($\widehat{m}_s$) would vary, and predictability increases. In this case, selective rationing would be effective in attracting or deterring certain types of individuals. In other words, the better the information that individuals have about their future health care use, the larger the distortion caused by the plan's selective rationing to avoid unprofitable individuals.

Second, *predictiveness* represents how use of a service is correlated with net profit per individual. This indicates whether a service is more likely to be used by those with financial gains or losses for the plan. When use of a service is negatively correlated with profits (π), the plan would want to ration the service to avoid those individuals associated with financial losses. When use of a service is positively correlated with profits, however, the plan would not want to ration the service to attract those with financial gains.

To summarize, Ellis and McGuire (2007) presented that the plan's incentives to ration at the service level are proportional to the product of *predictability* and *predictiveness*. For services that are either not predictable or does not correlate with net profit, a health plan has no incentive to ration. For services with a large positive value of I_s (i.e., services that are highly predictable and negatively correlated with net profit), however, the plan has incentives to ration these services tightly. For services with a large negative value of I_s (i.e., services that are highly predictable and positively correlated with net profit), the plan has incentives to not ration these services.

[14] Although the demand elasticity affects the magnitude of the index, it is unlikely to affect the order of the index. As shown in a literature review of Ringel et al. (2002), the price elasticity of demand for health care services is in general low. Almost services were estimated to be relatively less price-sensitive (in the range of -0.1 to -0.2 for inpatient, outpatient, and mental health services). Thus, we do not consider the demand elasticity in further analysis.

5.2. Our Extended Model on Service-level Selection in Response to Risk Adjustment

Building upon Ellis and McGuire (2007), we extend the model to show that MA plans are likely to employ service-level selection as a strategic behavior in response to the CMS-HCC model.

MA plans must accept all Medicare beneficiaries who wish to join and offer at least the same benefits as TM plans (i.e., services covered under Parts A and B).[15] CMS pays MA plans a fixed capitated payment to cover the costs for services covered under TM. Using the CMS-HCC model, CMS calculates payments to MA plans separately for each enrollee in the plan, multiplying the plan's payment rate by the enrollee's risk score r. CMS uses the prior-year's TM data to estimate risk scores for current MA enrollees.[16] The capitation payment for an MA enrollee is based on the estimated Parts A and B payments had a TM plan covered her directly. Following Ellis and McGuire (2007), let M_i denote total annual Medicare expenditure that individual i spends, and define $M_i = \sum_s m_{is}(q_s)$. Let $C(r_i)$ denote the risk-adjusted (capitated) payment that an MA plan receives from CMS for individual i with a risk score r.[17] The MA plan's profit for individual i is expressed as:

(3) $$\pi_i = C(r_i) - M_i$$

In an ideal risk-adjusted payment system, MA plans will have no incentive to select Medicare beneficiaries. Under an imperfect risk adjustment model, however, MA plans have incentives to discourage enrollment of those with predictably higher expenditures than their risk-adjusted payments, $\pi_i < 0$. The incentives would be stronger for those with high-risk scores than low-risk scores, as the CMS-HCC model underpredicts expenditures for those with the most severe health status (Medicare Payment Advisory Commission 2012). This can be expressed as:

(4) $$var(\pi_{i_h}) > var(\pi_{i_l})$$

where $var(\,)$ indicates the variance of a variable. i_h and i_l indicate those with high-risk scores and low-risk scores, respectively.

Given varying selection incentives across services, as shown in Ellis and McGuire (2007), MA plans would be interested in figuring out the relationship between the magnitude of

[15] Since risk-adjusted payments are based on services in Parts A and B, we focus on services covered by both TM and MA plans.
[16] The data includes TM beneficiaries entitled by age or disability with continuous 12-month enrollment in TM, and thus those entitled by end stage renal disease or those without 12 months base year Medicare enrollment are excluded. For each of them, a separate risk adjustment model is used to predict their next year expenditures.
[17] Since 2006, CMS has implemented a competitive bidding system to determine payments to MA plans. However, as a rebate must be returned to enrollees as a reduction in premiums or additional benefits, the bidding system is unlikely to affect $C(r_i)$.

the incentive to ration services used by beneficiaries with a given risk score and the degree to how far their total expenditures are from the mean of their conditional expenditure distribution. If a service that is more likely to be used by those with substantially higher expenditures than their risk-adjusted payment is the one that MA plans want to ration more tightly, then MA plans would ration care by the order of the service-level selection index estimated from Ellis and McGuire (2007). We assume that such rationing behaviors occur across MA plans competing for beneficiaries. This suggests that switching to another MA plan is unlikely, since in a competitive market MA plans behave similarly.

We demonstrate how the probability of using a service by beneficiaries with expenditures higher than their risk-adjusted payment is related to the Ellis and McGuire (2007)'s service-level selection index. Define u_{is} as individual i's actual use of services s. u_{is} takes the value 1 if individual i used service s and zero otherwise. Let $\hat{P}(u_{is}) \in \{0,1\}$ denote the probability that individual i expects to use service s. To effectuate service-level selection, MA plans do not need to forecast service use at the individual level but rather focus on forecasting at the population level. Thus, define the population-level expected probabilities of using service s given that an individual i's total actual expenditure is higher than her risk-adjusted payment as follows:

$$(5) \qquad \frac{\sum_i \left(\hat{P}(u_{is}, \pi_i < 0)\right)}{\sum_i \hat{P}(\pi_i < 0)} = \sum_i \left(\frac{\hat{P}(u_{is}) \hat{P}(\pi_i < 0 | u_{is})}{\hat{P}(\pi_i < 0)} \right)$$

where the last equality of the equation (5) is drawn from Bayes' theorem.

Then, we re-express as follows:

$$(6) \qquad \sum_i \hat{P}(u_{is}) \sum_i \left(\frac{\hat{P}(\pi_i < 0 | u_{\square s})}{\hat{P}(\pi_i < 0)} \right) - \sum_i \sum_{j, j \neq i} \left(\frac{\hat{P}(u_{is}) \hat{P}(\pi_i < 0 | u_{is})}{\hat{P}(\pi_i < 0)} \right)$$

The equation (6) can be written as:

$$(7) \qquad NN \left[\frac{1}{N} \sum_i \hat{P}(u_{is}) \right] \left[\frac{1}{N} \sum_i \left(\frac{\hat{P}(\pi_i < 0 | u_{is})}{\hat{P}(\pi_i < 0)} \right) \right] - \frac{1}{N} \frac{1}{N} \sum_i \sum_{j, j \neq i} \left(\frac{\hat{P}(u_{is}) \hat{P}(\pi_i < 0 | u_{is})}{\hat{P}(\pi_i < 0)} \right)$$

The equation (7) indicates that the probability of using a service by unprofitable individuals is directly proportional to the second component of the first term, $\left[\frac{1}{N} \sum_i \left(\frac{\hat{P}(\pi_i < 0 | u_{is})}{\hat{P}(\pi_i < 0)} \right) \right]$, which represents the average ratio of the probability that individual i is expected to incur a net loss given use of service s to the probability of a net loss for the individual i. This is another way to measure *predictiveness* of the service-level selection index, $(\rho_{\hat{m}_s, \pi})$, which measures the correlation of use of service with profitability to the plan. If $\left[\frac{1}{N}\sum_i\left(\frac{\hat{P}(\pi_i<0|u_{is})}{\hat{P}(\pi_i<0)}\right)\right]$ is high, it implies that service s is more likely to be used by those with financial losses. If $\left[\frac{1}{N}\sum_i\left(\frac{\hat{P}(\pi_i<0|u_{is})}{\hat{P}(\pi_i<0)}\right)\right]$ is low, on the other hand, it implies that service s is more likely to be used by those with financial profits. Theoretically, *predictiveness* can have either negative or positive value. However, Ellis and McGuire (2007) showed that all services (except for chiropractic services) were estimated to have positive values.

To sum up, where Ellis and McGuire (2007) stopped at showing the incentives for health plans to ration care tightly across services, we go further and demonstrate the relationship between the plans' incentive to ration a service tightly and the probability of using the service by unprofitable individuals across services.

(8)
$$\frac{\sum_i\left(\hat{P}(u_{is},\pi_i<0)\right)}{\sum_i\hat{P}(\pi_i<0)} \propto \rho_{\hat{m}_s,\pi} \propto I_s$$

The equation (8) presents that unprofitable beneficiaries under the CMS-HCC model are more likely to use services that MA plans want to ration more tightly. This suggests that MA plans have incentives to effectuate risk selection via service-level selection. This phenomenon is more pronounced for those with high-risk scores, as the CMS-HCC model systematically underpredicts expenditures for those with high-risk scores (Medicare Payment Advisory Commission 2012), which indicates that net losses increase with an increase in risk score. Hence, as risk scores increase, $\left[\frac{1}{N}\sum_i\left(\frac{\hat{P}(\pi_i<0|u_{is})}{\hat{P}(\pi_i<0)}\right)\right]$ is likely to increase.

6. Data

We use two data sets: the 2001-2009 Plan Benefit Package (PBP) and the 2001-2009 Medicare Current Beneficiary Survey (MCBS).

6.1. Plan-level Data: PBP

The PBP provides information on the set of benefits that an MA plan offers (e.g., premiums, cost-sharing, and additional benefits by service). The data are submitted to CMS for benefit analysis, marketing, and beneficiary communication purposes. Recently, CMS has used the data to review and approve all benefits yearly to ensure that MA plans do not discriminate against beneficiaries with poor health or those who incur financial losses.

We identify MA plans with complete information on cost-sharing for all services covered under Parts A and B. We exclude private fee-for-service (PFFS) plans because their characteristics are similar to TM plans despite that PFFS plans are classified as and paid like an MA plan. We also exclude cost-based, demonstration, special needs, Medicare Savings Account, and employer-sponsored plans because they are available only to small numbers of Medicare beneficiaries. Thus, we limit analysis to HMO and preferred provider organization (PPO) plans.

Table 1 shows summary statistics on MA plans by the three implementation periods of the CMS-HCC model. The mean numbers of MA plans were 481 (SD = 72) and 2,843 (SD = 348) in the pre- and post-implementation periods, respectively. The shares of HMO plans were 98.75 percent and 71.26 percent in the pre- and post-implementation periods, respectively.

6.2. Individual-level Data: MCBS

The MCBS is a longitudinal survey of a nationally representative sample of the Medicare population. CMS annually surveys a nationally representative sample of roughly 11,000 Medicare beneficiaries each year, and link with Medicare claims data. In each MCBS dataset, three rounds of interviews per year are conducted to collect detailed information on access to and satisfaction of care, functional status, medical conditions, health care expenditures, health insurance, and other health-related topics through the four-years.

The MCBS is particularly well suited for studying service-level selection in MA plans. First, it provides a nationally representative sample of the Medicare population with four-year follow-up. This allows us to track the switching behavior between TM and MA plans over time. Furthermore, the MCBS offers comprehensive information on health status and health care utilization for both TM and MA enrollees. While Medicare claims data offers complete information from Medicare-covered services for all TM beneficiaries in the sample, the claims data for MA enrollees is not publicly available. However, the MCBS obtains information on health status and health care utilization for all MA enrollees through survey. This enables us to capture comprehensive information for all TM and MA enrollees in the sample over time. Lastly, the MCBS offers comprehensive information on self-reported health outcomes and satisfaction of care. It allows us to examine the reason for plan switching.

We first identify a sample of Medicare beneficiaries who were eligible for both Medicare Parts A and B coverage in the two consecutive years (t and $t + 1$ years) during the study period. We exclude the following types of beneficiaries from the sample: beneficiaries whose original eligibility was attributable to disability or end-stage renal disease, newly eligible beneficiaries (since no prior claims information is available), those who died, dual-eligible beneficiaries, those who switched into Special Needs Plans, those who did not have 12 months of continuous enrollment in Medicare (both Parts A and B benefits) in year t, and those not enrolled in Medicare in January of year $t + 1$. Medicare beneficiaries are classified as TM enrollees if enrolled in a TM plan for all 12 months of the calendar year, and classified as MA enrollees if enrolled in an MA plan for at least one month of the year and enrolled in any Medicare plan in every month of the year. Finally, we construct two comparison groups. To examine whether the CMS-HCC model reduced the phenomenon that MA plans selectively avoid TM beneficiaries with high-risk scores, we compare TM stayers (those enrolled in TM during year t and remained in TM during year $t + 1$) and TM-to-MA switchers (those enrolled in TM during year t, but switched from TM to MA during year $t + 1$) (Panel A). To examine the MA plans' strategic risk selection behaviors in response to the CMS-HCC model, we compare MA stayers (those enrolled in MA during year t and remained in MA during year $t + 1$) and MA-to-TM switchers (those enrolled in MA during year t, but switched from MA to TM during year $t + 1$) (Panel B).

Table 2 shows baseline characteristics for the MCBS population by the three implementation periods. We find that the number and proportion of TM-to-MA switchers increased [76 (0.50 percent) and 470 (3.22 percent) for the pre- and post-implementation periods, respectively], whereas the number and proportion of MA-to-TM switchers decreased after the full phase-in period [272 (1.78 percent) and 131 (0.90 percent) for the pre- and post-implementation periods, respectively]. Moreover, we find that the difference in total Medicare expenditures between TM stayers and TM-to-MA switchers decreased after the full implementation period ($3,377 and $555 for the pre- and post-implementation periods, respectively). However, the difference in total Medicare expenditures between MA-to-TM switchers and MA stayers increased after the period ($1,431 and $5,269 for the pre- and post-implementation periods, respectively).

7. Did MA Plans Raise Copays More for Services Needed by Unprofitable Beneficiaries?

7.1. Empirical Strategy

To test whether MA plans employed service-level selection in response to the CMS-HCC model, we compare changes in weighted average service-specific copayments between the pre- and post-implementation periods. We calculate service-specific copayments of 33 services covered under Medicare Parts A and B benefits (Table 3). Most MA plans use copayments, but others use coinsurance rates. Assuming that there are marginal variations in service prices between TM and MA plans[18] and across MA plans, we convert coinsurance rates to copayments based on mean allowed charges or charges per TM beneficiaries for each service and year, which was estimated from the MCBS. To map each claim or line item into the PBP services, we use the service categories used in the Medicare Options Compare Out-of-Pocket Cost (OOPC) Estimates Methodology (Centers for Medicare and Medicaid Services 2008).[19] For inpatient hospital, skilled nursing facility, mental health specialty services, psychiatric services, and outpatient substance abuse services, MA plans can set up varying cost-sharing by a length of stay or number of visit. For these services, we calculate a copayment based on a typical length of stay or number of visit (Government Accountability Office 2010). For other services, a copayment is calculated per visit. All service-specific copayments are adjusted to 2009 US dollars by the equivalent service-specific price index (Agency for Healthcare Research and Quality). To account for varying numbers of MA plans across years, we adjust by weighting the number of MA plans in each year. Then, we plot the changes with respect to the service-level selection index estimated from Ellis and McGuire (2007), and compare the plotted relations between the pre-and post-implementation periods.

7.2. Results

Figure 1 presents evidence of service-level selection in the MA program after the CMS-HCC model was fully phased in. Both of the fitted lines for the pre- and post- implementation periods show an upward trend with respect to the service-level selection index, but the fitted line for the post-implementation period is tilted upward more than the pre-implementation period.

[18] Trish et al. (2017) found that physician reimbursement in MA plans was similar to or slightly less than TM rates. For a standard mid-level office visit with an established patient, the mean MA price was 96.9 percent of TM. For physician services, mean MA reimbursement ranged from 91.3 percent of TM for cataract removal in an ambulatory surgery center to 100.2 percent of TM for the professional fee for interpretation of a computed tomographic scan in an emergency department.

[19] The mapping identification for each service is conducted based on the Berenson-Eggers Type of Services (BETOS) codes, physician specialty codes, service type, place of service, bill type code, and revenue center code. The service-specific mapping identification is described in Centers for Medicare and Medicaid Services (2008). We use the 2009 OOPC Methodology, which is the model close to the year in which the CMS-HCC model was fully phased-in. Although the way of identifying each PBP service differs across years, dramatic changes are unlikely.

This indicates that MA plans increased enrollees' copayments disproportionately more for services with higher service-level selection index than services with lower service-level selection index after the full phase-in period. Specifically, MA plans increased copayments more for services with higher service-level selection index [e.g., ambulance (the ratio of weighted average copayments in the post-implementation period to the pre-implementation period: 3.38), home health services (2.08), partial hospitalization (1.98), inpatient hospital—psychiatric (1.37), and inpatient hospital—acute (1.36)] than services with lower service-level selection index [e.g., outpatient substance abuse services (0.92), outpatient X-rays (0.85), and outpatient hospital services (0.85)].

8. Did the CMS-HCC Model Reduce Risk Selection or Induce A Strategic Behavior?

8.1. Empirical Strategy

To examine whether the CMS-HCC model reduced the phenomenon of selectively avoiding TM enrollees with high-risk scores, we replicate analyses from Newhouse et al. (2015), which examined selection patterns at different implementation timings (i.e., after the initial and full phase-in of the CMS-HCC model, respectively). We perform this analysis with two purposes. The first is to examine whether a relatively small sample from the MCBS provides consistent results with a larger sample from Medicare claims. Following Newhouse et al. (2015), we compare selection during the pre-implementation period with selection during the post-implementation period. If our findings are consistent with those from Newhouse et al. (2015), then this indicates that our analysis with the MCBS provides generalizable results and insights. The second is to examine whether the effectiveness of the CMS-HCC model was larger after the full phase-in of the CMS-HCC model than the initial phase-in. For those who enrolled in TM plans during year t, risk scores are estimated based on the risk adjustment methodology.[20] For those who enrolled in MA plans during year t, since the claims data for MA enrollees is not

[20] The way of estimating their risk scores changed over the time period. Risk scores for the pre-implementation period (2001-2003) are estimated based on the PIP-DCG model (Pope et al. 2000). The coefficients estimated from Pope et al. (2000) are used. Risk scores for the post-implementation period (2007-2009) are estimated based on the CMS-HCC model (the HCC 2007 version 12 model). The coefficients estimated from Pope et al. (2004) are used, which is also available at the National Bureau of Economic Research website (http://www.nber.org/data/cms-risk-adjustment.html). Risk scores for the implementation period (2004-2006) are estimated by putting varying weights between the PIP-DCG and CMS-HCC models across years. In 2004, the CMS-HCC model had 30 percent weight in determining payment, in 2005, 50 percent weight, and in 2006, 75 percent weight. From 2004 to 2006, the remaining weight was on the PIP-DCG model.

publicly available, we follow the risk score estimation method from McWilliams, Hsu, and Newhouse (2012).[21]

To test the hypotheses, we conduct the following difference-in-difference analysis via ordinary least squares (OLS).

(9) $$Risk\ score_{it} = \alpha_0 + \alpha_1 Share\ of\ Year\ in\ MA_{i,t+1} + \alpha_2 Share\ of\ Year\ in\ MA_{i,t+1} \times After\ 2002_t + \alpha_3 Year\ Dummies + \epsilon_{it}$$

where $Risk\ score_{it}$ is beneficiary i's risk score at year t, $Share\ of\ Year\ in\ MA_{i,t+1}$ measures the share of the beneficiary's Medicare-eligible months that she stayed in MA plans in year $t+1$. $After\ 2002_t$ takes the value one for the years after 2003 and takes zero otherwise. We conduct the regression on those enrolled in TM plans all 12 months of the baseline years 2001-2005. Also, we perform the same regression with $After\ 2005_t$ on those enrolled in TM plans all months of the baseline years 2001-2002 and 2006-2008. We include year fixed effects, and use sample weights provided by the MCBS. Since we use repeated observations on individuals, standard errors are clustered at the individual level.

In the above equation, the key coefficients are those for $Share\ of\ Year\ in\ MA_{i,t+1}$ and $Share\ of\ Year\ in\ MA_{i,t+1} \times After\ 2002(or\ After\ 2005)_t$. Following Brown et al. (2014) and Newhouse et al. (2015), we interpret the results in a way that one simply assumes that TM-to-MA switchers spent the entire next year in MA plans so that the share of the next year spent in MA plans is one for TM-to-MA switchers and zero for TM stayers. Then, the predicted risk score for TM-to-MA switchers in 2002 is $\alpha_0 + \alpha_1$, whereas the predicted risk score for TM stayers in that year is just α_0. For subsequent years, one simply adds the coefficient of the interaction term. For those who switched beginning in 2004 or 2007, their predicted risk scores are $\alpha_0 + \alpha_1 + \alpha_2$. As such, we interpret the values of $\alpha_1 + \alpha_2$ as how much risk selection decreased or increased after adopting the CMS-HCC model. To check the robustness of our results, we perform additional specifications. We conduct the analysis excluding outliers (i.e., those with risk scores above the 95th percentile in each year), and estimate with quantile regressions instead of OLS.

We next examine whether MA plans selectively accepted TM enrollees with lower Medicare expenditures conditional on their risk scores after the CMS-HCC model. For those who

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[21] Enrollee-specific capitated payments to MA plans are calculated by multiplying county-specific benchmark rates by enrollee's demographic factors and individual HCC risk scores, modified somewhat by plan bids relative to benchmark rates. To obtain risk scores for MA enrollee each year, we divide capitated payments by county benchmark rates available from CMS.

enrolled in TM plans during year t, total Medicare expenditures are calculated by summing any Parts A and B expenditures. For those who enrolled in MA plans during year t, total Medicare expenditures for MA enrollees are estimated by summing any Parts A and B expenditures (if enrolled in TM plans) and the self-reported MA expenditures. We evaluate selection pattern after each of the two implementation points. To test the hypothesis, we conduct the following difference-in-difference analysis via OLS:

$$(10)\quad Expenditure_{it} = \beta_0 + \beta_1 Share\ of\ Year\ in\ MA_{i,t+1} + \beta_2 Share\ of\ Year\ in\ MA_{i,t+1} \times After\ 2002_t + \beta_3 Year\ Dummies + \beta_4 Risk\ score_{it} + \epsilon_{it}$$

where $Expenditure_{it}$ is beneficiary i's total Medicare expenditure at year t, and all other notation are the same as in the previous analysis. As with the above regression, we perform the regression on those who enrolled in TM plans all 12 months of the baseline years 2001-2005. Also, we perform the same regression with $After\ 2005_t$ on those who enrolled in TM plans all months of the baseline years 2001-2002 and 2006-2008.

To examine strategic risk selection behaviors of MA plans, we also perform the same analyses for those who enrolled in MA plans in year t (i.e., MA-to-TM switchers and MA stayers). Previous studies, including Brown et al. (2014) and Newhouse et al. (2015), have focused on examining whether the CMS-HCC model reduced the phenomenon of avoiding TM beneficiaries with high-risk scores, mainly due to lack of data for MA enrollees. Since the MCBS provides the data for MA enrollees, we can examine whether MA plans responded strategically to the CMS-HCC model to induce voluntary disenrollment of unprofitable MA enrollees. As with the above regression, we perform two regressions on those who enrolled in MA plans for at least one month of the baseline years 2001-2005 as well as 2001-2002 and 2006-2008, respectively. Using these results, we estimate cost savings attributable to such MA plans' strategic behavior in 2007-2009. Assuming that the switching rate of MA-to-TM switchers is generalizable to the entire MA population, we estimate the number of MA-to-TM switchers in the entire MA population in 2007-2009 (Kaiser Family Foundation 2015), and then multiply it by the average excess expenditures of MA-to-TM switchers beyond their risk-adjusted payments (i.e., the values of $\beta_1 + \beta_2$).

8.2. Results

Table 4 displays the results from re-estimating equations from Newhouse et al. (2015) (Panel A) and our own analyses (Panel B). We find that the phenomenon of avoiding TM beneficiaries with high-risk scores reduced after adopting the CMS-HCC model. For those who enrolled in TM plans during year t, column (1) shows that while TM-to-MA switchers had average risk scores roughly 0.14 points lower than TM stayers in the pre-implementation period, the switchers' risk scores increased by 0.1 after the initial phase-in, assuming they spent full year in MA plans. As shown in column (2), the the risk score difference was almost identical in magnitude after the full phase-in. When the outliers were excluded (columns 2 and 5) or the equation was estimated via quantile regression (columns 3 and 6), we find similar results. On the other hand, as shown in column (7)-(8), in the pre-implementation period, TM-to-MA switchers had baseline expenditures roughly $2,400 lower than TM stayers, assuming they spent full year in MA plans. The amount of favorable selection decreased during the initial phase-in period (by $1,910), and dramatically decreased after the full phase-in period (by $84).

However, we find evidence of strategic risk selection of MA plans in response to the CMS-HCC model. For those who enrolled in MA plans during year t, column (1) shows that while MA stayers had average risk scores roughly 0.01 points higher than MA-to-TM switchers in the pre-implementation period, the stayers' risk scores decreased by 0.15 after the initial phase-in, assuming the stayers spent full year in MA plans and the switchers spent full year in TM plans. After the full phase-in, as shown in column (2), the risk score difference slightly increased. We also find robust results when the outliers were excluded or the equation was estimated via quantile regression. Moreover, as shown in column (7)-(8), in the pre-implementation period MA stayers had baseline expenditures roughly $1,400 lower than TM-to-MA switchers. However, the amount of selection increased during the initial phase-in period (by $3,022), and increased even more after the full phase-in period (by $4,996). Based on these results, it is estimated that such strategic behavior led MA plans to save costs of $5.2 billion (=$4,996 × 3.7% × 28.6 million) in 2007-2009.

9. Did Service-level Selection Induce Voluntary Disenrollment of Unprofitable Beneficiaries?

9.1. Empirical Strategy

To test whether service-level selection affected the individuals' plan switching behavior after the full phase-in period of the CMS-HCC model, we compare service-specific use between switchers and stayers in the pre-implementation period with that in the post-implementation period. We measure health care utilization by type of service. For those who enrolled in TM during year t, we use claims to create a total of 29 types of services categories (Table 5). Part A claims are classified into the following five types of service (hospital inpatient visit, hospital outpatient visit, home health care, hospice, and other facility services). Part B claims are classified into 24 categories by the Berenson-Eggers Type of Services (BETOS) code, which is used to create clinically meaningful groupings of procedures and services to analyze Medicare expenditures by type of service. For those who enrolled in MA plans during year t, we estimate service-level use through the survey. In the survey, participants reported their use of health care by the following 7 types of service: inpatient hospitalizations, outpatient department visits, home health care, skilled nursing facility, medical provider events, office visits, and durable medical equipment supplier use. Thus, we create a total of 7 types of services categories (Table5). To examine whether TM enrollees who used services with lower service-level selection index were more likely to switch to MA plans than those who used services with higher service-level selection index after the full phase-in ("intensive-margin selection"), we estimate the ratio of the proportion of TM-to-MA switchers with use of a particular type of service to TM stayers with use of the service in the pre- and post-implementation periods, respectively. To examine whether TM enrollees who used more services with lower service-level selection index were more likely to switch to MA plans than those who used more services with higher service-level selection index after the full phase-in ("extensive-margin selection"), we also estimate the ratio of average number of services per enrollee of TM-to-MA switchers to TM stayers in the pre- and post-implementation periods, respectively. Then, we plot the ratios with respect to the service-level selection index estimated from Ellis and McGuire (2007), and compare the plotted relations between the pre-and post-implementation periods. We perform the same analysis for those who enrolled in MA plans in year t.

To test whether service-level selection allowed MA plans to reduce the scope of enrolling those with higher expenditures than their risk-adjusted payments, especially for those with high-risk scores, we estimate the coefficients of variance of total Medicare expenditures for TM and MA enrollees, respectively, over time by risk score. The coefficient of variance of total Medicare expenditures is estimated as the ratio of the standard deviation of total Medicare expenditures to its mean. We divide the study population into enrollees with high-risk scores and other risk scores. High-risk scores indicate above the 90^{th} percentile of the risk score distribution in each year.

9.2. Results

The above two figures in Figure 2 presents changes in service use of TM-to-MA switchers to TM stayers between the pre- and post-implementation periods. In both figures showing results for intensive and extensive margin selection, respectively, we observe that for most services, the ratios of service use of TM-to-MA switchers to TM stayers are lower than one. Also, the fitted lines for the pre- and post-implementation periods show a downward trend with respect to the service-level selection index, with an almost same slope and are below the ratio of one. However, the intercept of the post-implementation period is higher than the intercept of the post-implementation period. We also find that after the full phase-in period, TM-to-MA switchers systematically used more services across all services compared to TM switchers (not shown). However, even after the full implementation period, TM enrollees who used services with higher service-level selection index were more likely to switch to MA plans than those who used services with lower service-level selection index. When the outlier service with the highest value of the service-level selection index (hospice) was excluded, we find similar findings (Appendix Figure 1). However, we observe that the fitted line for the post-implementation period is tilted upward more than the pre-implementation period. This indicates that TM enrollees who used services with higher service-level selection index in the full phase-in period were more likely to switch to MA plans than the equivalent population in the pre-implementation period.

On the other hand, the below two figures show changes in service use of MA-to-TM switchers to MA stayers between the pre- and post-implementation periods. In the left figure showing results for intensive margin selection, we find that the fitted line for the post-implementation period is below the fitted line for the pre-implementation period. Furthermore, the fitted line for the post-implementation period is tilted downward more than the pre-implementation period. This indicates that the ratio of the proportion of using a service between MA-to-TM switchers and MA stayers reduced after the full implementation period. In the right figure showing results for extensive margin selection, however, we observe the opposite findings. The fitted line for the post-implementation period shows an upward trend with respect to the service-level selection index, whereas the fitted line for the pre-implementation period shows an almost flat trend. This indicates that MA enrollees who used services with higher service-level selection index after the full implementation period were more likely to disenroll from MA plans than the equivalent population in the pre-implementation period.

The left and right figures in Figure 3 present the coefficients of variance of total Medicare expenditures for TM enrollees and MA enrollees by risk scores, respectively. For TM enrollees with both high-risk scores and other risk scores, the coefficients of variance of total Medicare expenditures decreased over time. However, we observe different patterns for MA enrollees by risk scores. The coefficients of variance of total Medicare expenditures for MA enrollees with high-risk scores show a downward trend over time, whereas those for MA enrollees with other risk scores show a slightly upward trend over time.

10. Why Did MA Enrollees Disenroll from MA Plans?

10.1. Empirical Strategy

To examine the reasons of disenrollment from MA plans, we conduct various analyses. To test whether high-risk score enrollees who stayed in MA plans longer were more likely to disenroll from MA plans than those who stayed in MA plans shorter, we estimate the coefficient of variance of total Medicare expenditures for MA enrollees with high-risk scores by MA enrollment periods. Due to a relatively small sample size of those with high-risk scores, we categorize the population into those with 12 months enrollment in MA plans and those with less than 12 months enrollment in MA plans. Moreover, we examine whether disenrollment from MA plans was related to lower satisfaction on care costs, quality of care, or access to care. Satisfaction is measured by four levels: very dissatisfied, dissatisfied, satisfied, and very satisfied. To test this hypothesis, we conduct the following difference-in-difference analysis via OLS:

$$(11)\quad Satisfaction_{it} = \gamma_0 + \gamma_1 Share\ of\ Year\ in\ MA_{i,t+1} + \gamma_2 Share\ of\ Year\ in\ MA_{i,t+1} \times After\ 2005_t + \gamma_3 Year\ Dummies + \gamma_4 Risk\ score_{it} + \gamma_5 Health_{it} + \gamma_6 Demograhics_{it} + \epsilon_{it}$$

where $Satisfaction_{it}$ is beneficiary i's reported satisfaction in year t. $Health$ measures the five-category self-reported health variable (one "poor" up to five "excellent") and

$Demographics$ includes age, race, female, and disabled status.[22] All other notations are the same as in the previous analysis. We perform the regression on those who enrolled in MA plans for at least one month of the baseline years 2001-2002 and 2006-2008.

10.2. Results

Figure 4 displays the coefficients of variance of total Medicare expenditures for MA enrollees with high-risk scores by MA enrollment periods. We find that the coefficients of variance of total Medicare expenditures for those with 12 months enrollment in MA decreased more steeply than those with less than 12 months enrollment.

Table 6 shows the results from examining the relation of MA disenrollment and satisfaction of care after the fully phase-in period. Column (1) shows that after the full phase-in period, the relation of MA disenrollment and satisfaction on care costs was the most pronounced among satisfaction measures considered in this study. Specifically, it was shown that relative to MA-to-TM switchers, MA stayers were less satisfied with out-of-pocket costs by 0.004 points in the pre-implementation period, assuming the stayers spent full year in MA plans and the switchers spent full year in TM plans. However, MA stayers were more satisfied with out-of-pocket costs by 0.176 points than MA-to-TM switchers after the full implementation of the CMS-HCC model. As shown in column (2), the similar trend for satisfaction on overall quality of care. Compared to MA-to-TM switchers, MA stayers were less satisfied with overall care quality by 0.025 points in the pre- phase-in period, assuming the stayers spent full year in MA plans and the switchers spent full year in TM plans. However, MA stayers were more satisfied with overall care quality by 0.110 points than MA-to-TM switchers in the full phase-in period.

11. Discussion and Conclusion

The goal of this paper is to shed light on the competing claims on the effectiveness of the CMS-HCC model and to comprehensively understand strategic risk selection behaviors of MA plans. We find that the CMS-HCC model reduced the phenomenon that MA plans avoid beneficiaries with high-risk scores in TM plans, whereas it led to increased disenrollment of high-cost beneficiaries, conditional on risk score, in MA plans. We explain this phenomenon

[22] We adjust for self-reported health status because individuals with poor health are likely to report negative feelings toward one's health care. We also control for age, race, female, and disabled status because different demographic groups are likely to evaluate their health and health care differently.

through service-level selection. Through theoretical and empirical analysis, we show that after the full phase-in period of the CMS-HCC model, MA plans have the incentive to and did increase copayments disproportionately more for services that appeal to beneficiaries who could be unprofitable under the CMS-HCC model than other services. The disproportionate changes in copayments led to voluntary disenrollment of beneficiaries with need for these services, who tend to incur higher expenditures than their risk-adjusted payments. We also find evidence supporting our hypothesis that those who were less satisfied with out-of-pocket costs were more likely to disenroll from MA plans. Such strategic behavior led to MA plans to save $5.2 billion in 2007-2009 by simply transferring the costs to the federal government, thereby placing significant financial burdens on the federal government.

Our study shows evidence of the intended effect of the CMS-HCC model on reducing risk selection for TM beneficiaries with high-risk scores, consistent with findings from Newhouse et al. (2015). Specifically, the differences in risk scores between TM-to-MA switchers and TM stayers reduced by a factor of three. Also, the differences in total Medicare expenditures between them decreased after the full phase-in period to $84. This shows the intended consequences of the implementation of the CMS-HCC model. As risk adjustment leads to neutral payments for those with conditions included in the risk adjustment formula, MA plans no longer have incentive to avoid them. These findings add to earlier studies that found that a more clinically detailed risk-adjustment model strengthens the incentives for MA plans to retain sick enrollees. However, the risk selection did not disappear, but rather changed in form.

The main contribution of this study is to show that MA plans magnified service-level selection in response to the CMS-HCC model to avoid beneficiaries who could be costly under the CMS-HCC model. Our theoretical analysis demonstrates that MA plans have incentives to effectuate risk selection via service-level selection, as unprofitable beneficiaries under the CMS-HCC model are more likely to use services that are expensive and are thus more vulnerable to under-provision by MA plans. It informs why MA plans more engaged in service-level selection after adopting the CMS-HCC model. Then, our empirical analysis provides supporting evidence showing that MA plans increased copayments disproportionately more for services that are more likely to be used by them. This validates our theoretical analysis and contributes to providing evidence-based policy implications. With the theoretical foundation and empirical evidence based closely on the theoretical foundation, this study adds to the body of literature regarding

MA plans' strategic response to the policy-induced change in financial incentives as an important contributor to service-level selection.

It is worthwhile to note that relatively large increases in copayments were found in two types of services. First, increases in copayments for home health services and inpatient psychiatric hospital services were likely targeted for disenrollment of beneficiaries with multiple chronic conditions. This is because their expenditures were systematically underpredicted by the CMS-HCC model. For example, the model, on average, underestimated expenditures for those with more than six chronic conditions by $608 (Government Accountability Office 2011). Also, increases in copayments for ambulance and acute inpatient hospital services were likely intended to encourage disenrollment of those with potentially high risks because of poor health behaviors. Consequently, such service-level selection would likely lead to voluntary disenrollment of those who currently need these services as well as those who potentially have the need for these services.

Our study also shows the pronounced effect of service-level selection on MA enrollees after the full phase-in of the CMS-HCC model. Specifically, we find that MA enrollees who used services with high service-level selection index in the previous year were more likely to disenroll from MA plans in the following years than those who used services with low service-level selection index in the previous year. This phenomenon was observed during the initial phase-in period, and it was magnified with the full phase-in, showing that the amount of selection increased by $3,033 and $4,996, respectively. We also show that MA enrollees who were less satisfied with their out-of-pocket costs were more likely to disenroll from MA plans. These findings suggest that MA enrollees were more likely to disenroll from MA plans due to increased burdens on out-of-pocket costs as a result of service-level selection. The effects were more pronounced for those who stayed in MA plans longer, as they were more likely to be exposed to high out-of-pocket costs. Service-level selection could result in poor health status, because it is likely to lead to delayed care during the MA enrollment period, and inefficient or uncoordinated cares following enrollment switching to TM plans. This is especially true for those with multiple chronic conditions, which requires more integrated and coordinated care due to complex conditions and treatment.

Although the size of MA-to-TM switchers was only about one percent of the entire Medicare population, risk selection in this population cannot be considered trivial with the following three reasons. First, the size of the population increases over time. From our data, 3.7 percent of MA enrollees left their MA plans between 2007 and 2009. However, in 2014, nearly 12 percent left their MA plans (Government Accountability Office 2017). Also, the cost implications for this population would be significant given that the top five percent of the US population accounts for about 50 percent of total health care expenditures (Cohen 2014). If MA-to-TM switchers keep experiencing delayed care or receiving fragmented care in non-managed care settings, this would incur even higher treatment costs. Moreover, as MA payments are partly determined by the average expenditures of TM beneficiaries at the county level, switching of high-cost MA enrollees to TM plans could lead to MA payment increases, possibly placing significant financial burdens on the federal government.

Our findings provide key implications for CMS in developing a better risk adjustment model. To ensure that MA plans' benefit package designs do not discriminate against beneficiaries in poor health with high health care expenditures, since 2010, CMS has reviewed all benefit packages yearly (Government Accountability Office 2010). In addition to the review process for MA plans' benefit structures, developing a better risk adjustment model is inevitable as MA plans would continue to engage in risk selection if a risk adjustment model does not estimate capitation payments as sufficiently close to the actual expenditures. The new risk adjustment model needs to be designed to generate economic forces to prevent MA plans' strategic behaviors in engaging in service-level selection. Specifically, developing a risk adjustment model that not only conditions on each beneficiary's risk scores but also reflect each beneficiary's potential service-level use may contribute to reducing service-level selection. This approach enables to provide overpayments for services that are more likely used by unprofitable beneficiaries and underpayments for services that are more likely used by profitable beneficiaries, thereby equalizing incentives in rationing all services (Glazer and McGuire 2000). By regarding risk adjustment as a tax/subsidy scheme, overpaying for services in high demand by unprofitable beneficiaries and underpaying for services in high demand by profitable beneficiaries would redistribute health care costs away from profitable beneficiaries and toward unprofitable beneficiaries. As the payment for profitable services is much smaller than the payment for unprofitable services, this would penalize MA plans for attracting only those in need of profitable services. Thus, the new risk adjustment model could reduce the potential for MA plans to use service-level selection.

This study has several limitations. First, we assumed that the magnitude of the incentives to ration care tightly at the service level is consistent across time. However, it might not be true because reimbursement policies and rates change over time, possibly affecting the magnitude of the incentive across years. Following Ellis and McGuire (2007), we also assumed that all individuals share the same elasticity of demand for a certain service. However, the demand elasticity might differ across services (Manning et al. 1987) as well as individuals. Moreover, Ellis, Martins, and Zhu (Forthcoming) further developed the service-level selection index by accounting for variation in cost-sharing, risk-adjusted profits, and demand elasticities across services. However, estimates of the new service-level selection index were empirically calculated based on commercial claims data. Considered differences in demographic profiles and health care utilization patterns between the Medicare population and the commercial insured, the estimates are unlikely to be applicable to our study. Furthermore, we assumed small variations in service prices and service utilization between TM and MA plans and across MA plans. Thus, we converted service-specific coinsurance rates to copayments using mean allowed charges per TM beneficiaries for each year and year. However, prices in MA plans are not equivalent to those in TM plans (Trish et al. 2017, Baker et al. 2016). Also, there is substantial heterogeneity in cost and market power across MA plans (Glazer and McGuire 2016). Moreover, we used self-reported data to estimate utilization and Medicare expenditures for MA enrollees, which is significantly underreported (Eppig and Chulis 1997). However, it is unlikely that such reporting errors have systematically changed over the study period (McWilliams, Hsu, and Newhouse 2012). Finally, disproportionate changes in copayments could have induced to prevent disenrollment of TM beneficiaries in high need of these services. Individuals generally consider various aspects of plan benefits in changing a health plan (McCormack et al. 2005, Government Accountability Office 2017) . However, this study focuses only on cost-sharing structures. Therefore, there is the possibility that MA plans lessened other strategies of distorting service offerings such as access to specialist or provision of additional benefits in order to accept sicker TM beneficiaries who were no longer unprofitable under the CMS-HCC model, which is beyond the scope of this study.

Findings from this study indicate that the CMS-HCC model reduced the MA plans' risk selection of avoiding TM beneficiaries with high-risk scores, whereas it induced MA plans to strategically behave in response to the CMS-HCC model via service-level selection. MA plans have raised copayments disproportionately more for services needed by high-need beneficiaries than for other services, thereby inducing unprofitable beneficiaries to voluntarily disenroll from their MA plans, mainly due to increased out-of-pocket costs. This allows MA plans to avoid the risk of enrolling unprofitable beneficiaries. Our results provide key policy implications for CMS in moving towards a better risk adjustment model that accounts for the enrollees' predicted risk scores while generating economic incentives for MA plans that discourage service-level selection.

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Table 1. Summary Statistics on MA Plans by the Implementation Periods of the CMS-HCC Model

	Implementation periods of the CMS-HCC model		
	Pre-implementation period (2001-2003)	Implementation period (2004-2006)	Post-implementation period (2007-2009)
Types of MA plans, Weighted Mean (SD)			
Health Maintenance Organization (HMO)	475 (73)	1,136 (502)	2,026 (175)
Preferred Provider Organization (PPO)	6 (5)	344 (311)	817 (178)
Total	481 (72)	1,480 (813)	2,843 (348)

Notes: Plans with complete information on cost-sharing for all services covered under Medicare Parts A and B were included. Other types of MA plans such as Medicare Savings Account, and Private Fee-For-Service plans were excluded from our analysis. To account for varying numbers of MA plans across years, we adjust by weighting the number of MA plans in each year.

Table 2. Baseline Characteristics for the MCBS Sample by the Implementation Periods of the CMS-HCC Model

	Implementation periods of the CMS-HCC model (baseline year t equals)		
	Pre-implementation period (2001-2002)	Implementation period (2003-2005)	Post-implementation period (2006-2008)
Transition frequencies, N (percent)			
TM (year t) → TM (year $t+1$)	13,312 (86.88)	12,454 (81.94)	10,588 (72.61)
TM (year t) → MA (year $t+1$)	76 (0.5)	338 (2.22)	470 (3.22)
MA (year t) → TM (year $t+1$)	272 (1.78)	100 (0.66)	131 (0.9)
MA (year t) → MA (year $t+1$)	1,662 (10.85)	2,307 (15.18)	3,393 (23.27)
Total	15,322 (100.00)	15,199 (100.00)	14,582 (100.00)
Total Medicare expenditures at year t, Mean (SD)			
TM (year t) → TM (year $t+1$)	5,461 (12560)	6,111 (14,440)	7,003 (13,497)
TM (year t) → MA (year $t+1$)	2,084 (4141)	4,313 (9,252)	6,448 (12,492)
MA (year t) → TM (year $t+1$)	5,595 (12031)	6,015 (10,853)	8,995 (16,007)
MA (year t) → MA (year $t+1$)	4,164 (9305)	3,617 (10,273)	3,726 (9,380)
Weighted average for all beneficiaries	5,306	5,692	6,240
Risk scores at year t, Mean (SD)			
TM (year t) → TM (year $t+1$)	1.00 (0.32)	1.00 (0.32)	1.00 (0.41)
TM (year t) → MA (year $t+1$)	0.92 (0.33)	0.96 (0.29)	0.96 (0.37)
MA (year t) → TM (year $t+1$)	1.01 (0.27)	1.15 (0.43)	1.07 (0.97)
MA (year t) → MA (year $t+1$)	1.04 (0.34)	1.06 (0.35)	1.09 (0.76)
Weighted average for all beneficiaries	1.00	1.01	1.02

Notes: Medicare enrollees were classified as TM enrollees if enrolled in TM plans for all 12 months of the calendar year, and classified as MA enrollees if enrolled in an MA plan for at least one month of the year and enrolled in any Medicare plan in every month of the year. Total Medicare expenditures for TM enrollees were estimated by summing any Part A and Part B expenditures, and total Medicare expenditures for MA enrollees were estimated by summing any Part A and Part B expenditures (if enrolled in TM) and the self-reported MA expenditures. All expenditures were adjusted to 2009 dollars using the CPI-U. Risk scores for TM enrollees were estimated from Medicare claims and risk scores for MA enrollees were estimated by dividing the reported capitation payments by county-level benchmark rates. The way of estimating the risk scores for TM enrollees varied by the implementation periods, and thus risk scores cannot be directly comparable across the three periods. Sample weights provided by the MCBS were used.

Table 3. Type of Services from the PBP Data

Type of service	Unit of analysis
Inpatient hospital—acute	6 days
Inpatient hospital—psychiatric	21 days
Skilled nursing facility	35 days
Comprehensive outpatient rehabilitation	1 visit
Emergency care	1 visit
Urgent care	1 visit
Partial hospitalization	1 visit
Home health services	1 visit
Primary care physician services	1 visit
Chiropractic services	1 visit
Occupational therapy services	1 visit
Physician specialist services	1 visit
Mental health specialty services—individual session	50 sessions
Mental health specialty services—group session	50 sessions
Podiatrist services	1 visit
Other health care professional services	1 visit
Psychiatric services—individual session	50 sessions
Psychiatric services—group session	50 sessions
Physical therapy and speech/language pathology services	1 visit
Diagnostic services	1 visit
Radiation therapy services	1 visit
Outpatient X-Rays	1 visit
Outpatient hospital services	1 visit
Ambulatory Surgical Center (ASC) services	1 visit
Outpatient substance abuse services—individual session	50 sessions
Outpatient substance abuse services—group session	50 sessions
Cardiac rehabilitation services	1 visit
Ambulance	1 visit
Durable medical equipment	1 visit
Medical supplies	1 visit
Prosthetics	1 visit
Diabetes monitoring supplies	1 visit
Drug prescription	1 visit

Notes: MA plans can set up varying copayments by a length of stay or number of visit, for example, inpatient hospital, skilled nursing facility, mental health specialty services, psychiatric services, and outpatient substance abuse services. For these services, a copayment was estimated on a basis of a typical length of stay or number of visit, according to Government Accountability Office (2010).

Table 4. Changes in Risk Selection Patterns after Adopting the CMS-HCC Model

	Dependent variable: risk score at *year t* or total Medicare expenditure at year t							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Risk score	Risk score	Risk score	Risk score	Risk score	Risk score	Expenditure	Expenditure
Panel A (Those enrolled in TM at year t)								
Share of year in MA	-0.14	-0.20	-0.20	-0.14	-0.20	-0.20	-2300.12	-2401.84
	(0.07)	(0.04)	(0.09)	(0.07)	(0.04)	(0.09)	(1162.01)	(1135.36)
Share of year in MA × after 2002	0.10	0.18	0.18				389.5	
	(0.07)	(0.04)	(0.09)				(1295.64)	
Share of year in MA × after 2005				0.09	0.18	0.17		2317.51
				(0.07)	(0.04)	(0.09)		(1319.32)
Risk score							14020.01	13306.74
							(735.75)	(524.18)
Mean outcome variable	1.00	0.94	1.00	1.00	0.93	1.00	5,756.85	6,190.07
Estimated method	OLS	OLS	Quantile	OLS	OLS	Quantile	OLS	OLS
Evaluation period	After 2003	After 2003	After 2003	After 2006	After 2006	After 2006	After 2003	After 2006
Outliers trimmed	No	Yes	No	No	Yes	No	No	No
Observations	26,180	23,973	26,180	24,446	22,380	24,446	26,180	24,446
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Risk score	Risk score	Risk score	Risk score	Risk score	Risk score	Expenditure	Expenditure
Panel B (Those enrolled in MA at year t)								
Share of year in MA	0.01	0.00	-0.01	0.01	0.00	-0.01	-1400.90	-1418.53
	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(741.82)	(741.59)
Share of year in MA × after 2002	-0.16	-0.05	-0.10				-1621.68	
	(0.05)	(0.03)	(0.05)				(1434.73)	
Share of year in MA × after 2005				-0.18	-0.04	-0.07		-3577.63
				(0.09)	(0.05)	(0.09)		(1393.99)
Risk score				-0.18	-0.04	-0.07	155.42	1576.19
				(0.09)	(0.05)	(0.09)	(360.43)	(282.19)
Mean outcome variable	1.05	1.07	1.05	1.08	0.98	1.08	3,999.28	4,057.00
Estimated method	OLS	OLS	Quantile	OLS	OLS	Quantile	OLS	OLS
Evaluation period	After 2003	After 2003	After 2003	After 2006	After 2006	After 2006	After 2003	After 2006
Outliers trimmed	No	Yes	No	No	Yes	No	No	No
Observations	4,340	4,120	4,340	5,458	5,182	5,458	4,340	5,458

Notes: "Outliers trimmed" means exclusion of individuals with risk scores above the 95^{th} percentile in each year. Year fixed effects were included in all regressions. Sample weights provided by the MCBS were used. Standard errors, in parentheses, were clustered by the individual.

Table 5. Type of Services Used to Examine Service-level Selection in the MCBS Sample

Type of service[a]	Service-level selection index[b]	TM enrollees[c]	MA enrollees[d]
Hospice	2.578	Yes	No
Home health	0.875	Yes	Yes
Durable medical equipment	0.703	Yes	Yes
Hospital inpatient visit	0.592	Yes	Yes
Other	0.495	Yes	No
Hospital visit	0.356	Yes	No
Home visit	0.348	Yes	No
ER visit	0.265	Yes	No
Consultation	0.219	Yes	No
Other facility services	0.172	Yes	Yes
Hospital outpatient visit	0.170	Yes	Yes
Advanced imaging—CAT	0.169	Yes	No
Oncology	0.159	Yes	No
Lab tests	0.144	Yes	No
Other tests	0.134	Yes	No
Standard imaging	0.119	Yes	No
Specialist	0.114	Yes	No
Echography	0.113	Yes	No
Ambulatory procedures	0.105	Yes	No
Imaging procedure	0.102	Yes	No
Office visit	0.096	Yes	Yes
Major procedure—cardiovascular	0.096	Yes	No
Minor procedure	0.095	Yes	No
Anesthesia	0.092	Yes	No
Endoscopy	0.087	Yes	No
Major procedure	0.087	Yes	No
Major procedure—orthopedic	0.083	Yes	No
Advanced imaging—MRI	0.083	Yes	No
Eye procedure	0.045	Yes	No

Notes: Services covered under Medicare Parts A and B were classified into 29 services. Specifically, Part A claims were classified into the following five types of service (hospital inpatient visit, hospital outpatient visit, home health, hospice, and other facility service). Part B claims were classified into 24 categories by the Berenson-Eggers Type of Services (BETOS) codes. Service-level selection index estimated from Ellis and McGuire (2007) was used and type of service was presented by the order of the service-level selection index. Service-level use for TM beneficiaries was estimated from MCBS claims data. Service-level use for MA enrollees was estimated from self-reported data.

Table 6. Relation of MA Disenrollment and Satisfaction after Adopting the CMS-HCC Model

	Dependent variable: satisfaction rating at *year t*				
	(1) Out-of-pocket costs	(2) Quality of care	(3) Access to specialist	(4) Ease of access to care from residence	(5) Care provided in the same location
Panel B (Those enrolled in MA at year t)					
Share of year in MA	-0.004	-0.025	0.012	-0.011	-0.013
	(0.05)	(0.04)	(0.04)	(0.05)	(0.04)
Share of year in MA	0.180	0.135	0.000	0.103	0.047
× after 2005	(0.08)	(0.06)	(0.06)	(0.08)	(0.06)
Risk score	0.011	0.011	0.025	-0.002	0.006
	(0.02)	(0.01)	(0.01)	(0.02)	(0.01)
Mean outcome variable	2.99	3.29	3.17	3.11	3.12
Observations	5,144	4,956	4,719	2,579	4,576
	(6) Availability of care nights and weekends	(7) Follow-up care	(8) Questions answered over phone	(9) Doctor's concern for your health	(10) Information about your medical condition
Panel B (Those enrolled in MA at year t)					
Share of year in MA	0.029	0.001	0.055	0.024	0.074
	(0.04)	(0.04)	(0.05)	(0.04)	(0.04)
Share of year in MA	0.054	0.003	-0.018	-0.072	-0.085
× after 2005	(0.05)	(0.06)	(0.07)	(0.06)	(0.06)
Risk score	0.004	0.026	0.012	0.015	0.019
	(0.01)	(0.01)	(0.02)	(0.01)	(0.01)
Mean outcome variable	3.19	3.18	3.04	3.16	3.13
Observations	5,188	4,524	3,745	5,047	5,110

Notes: Each satisfaction measure took values from one to four ("very dissatisfied", "dissatisfied", "satisfied", "very satisfied"). Self-reported health, age, race, female, and disabled status were adjusted. Year fixed effects were included in all regressions. Sample weights provided by the MCBS were used. Standard errors, in parentheses, were clustered by the individual.

Figure 1. Disproportionate Changes in Service-specific Copayments between the Pre-and Post-implementation Periods of the CMS-HCC Model

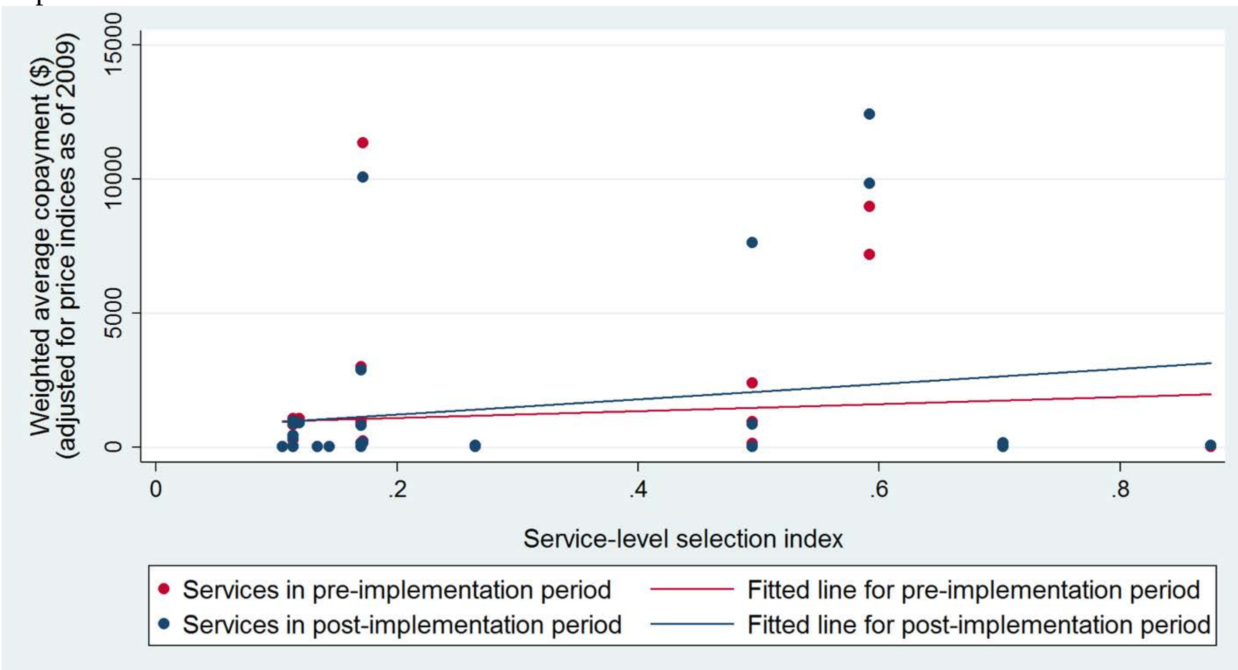


Figure 2. Changes in Service-specific Use of Switchers to Stayers between the Pre- and Post-implementation Periods of the CMS-HCC Model

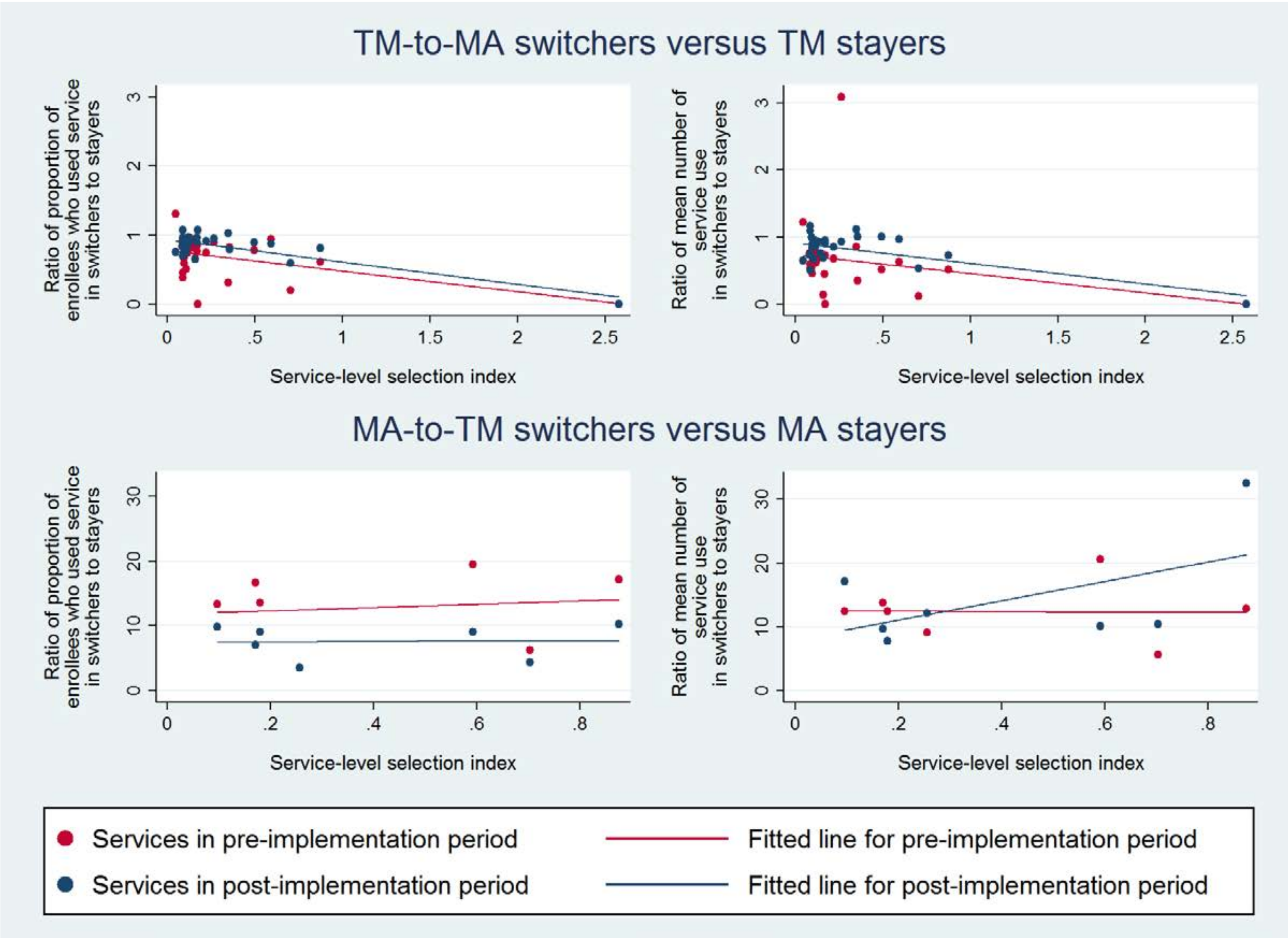
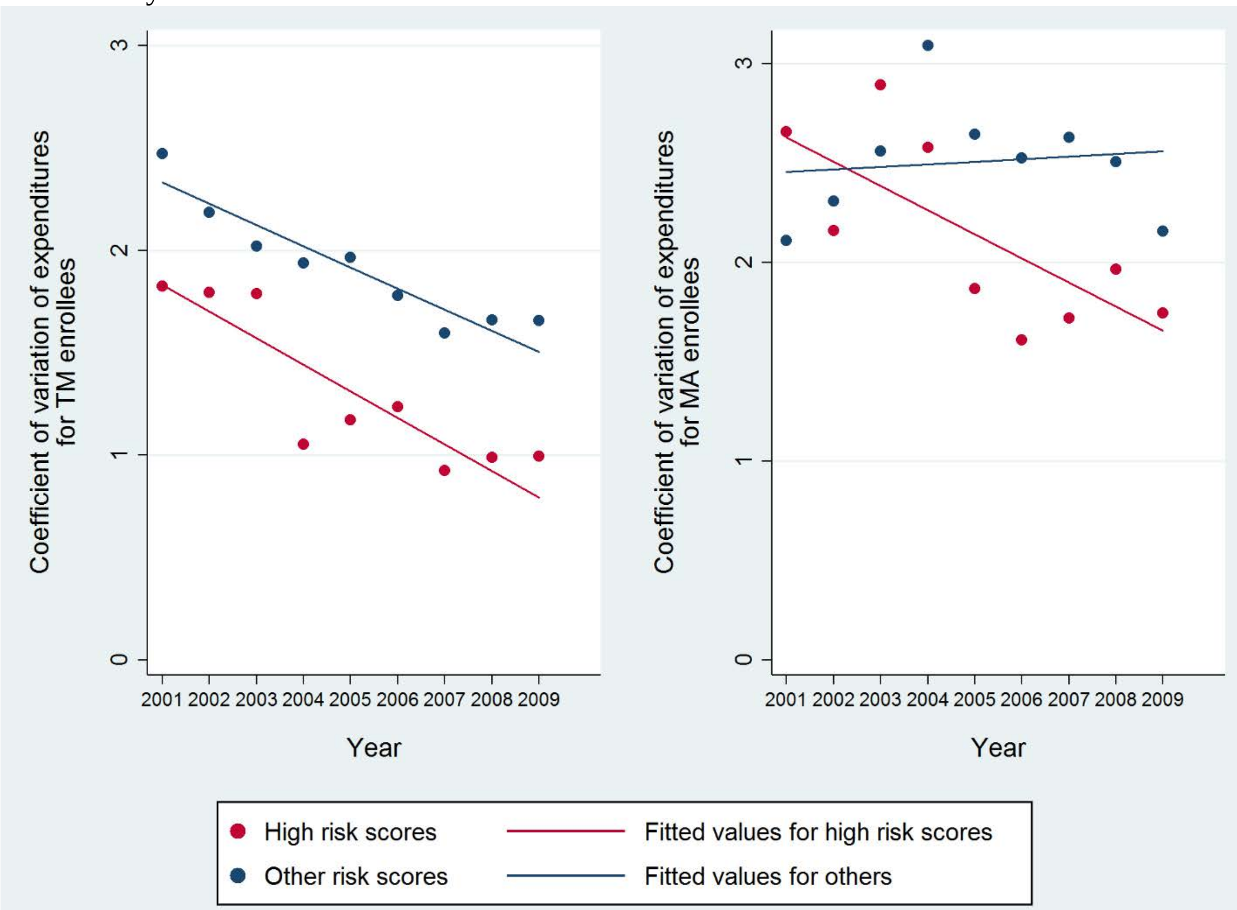
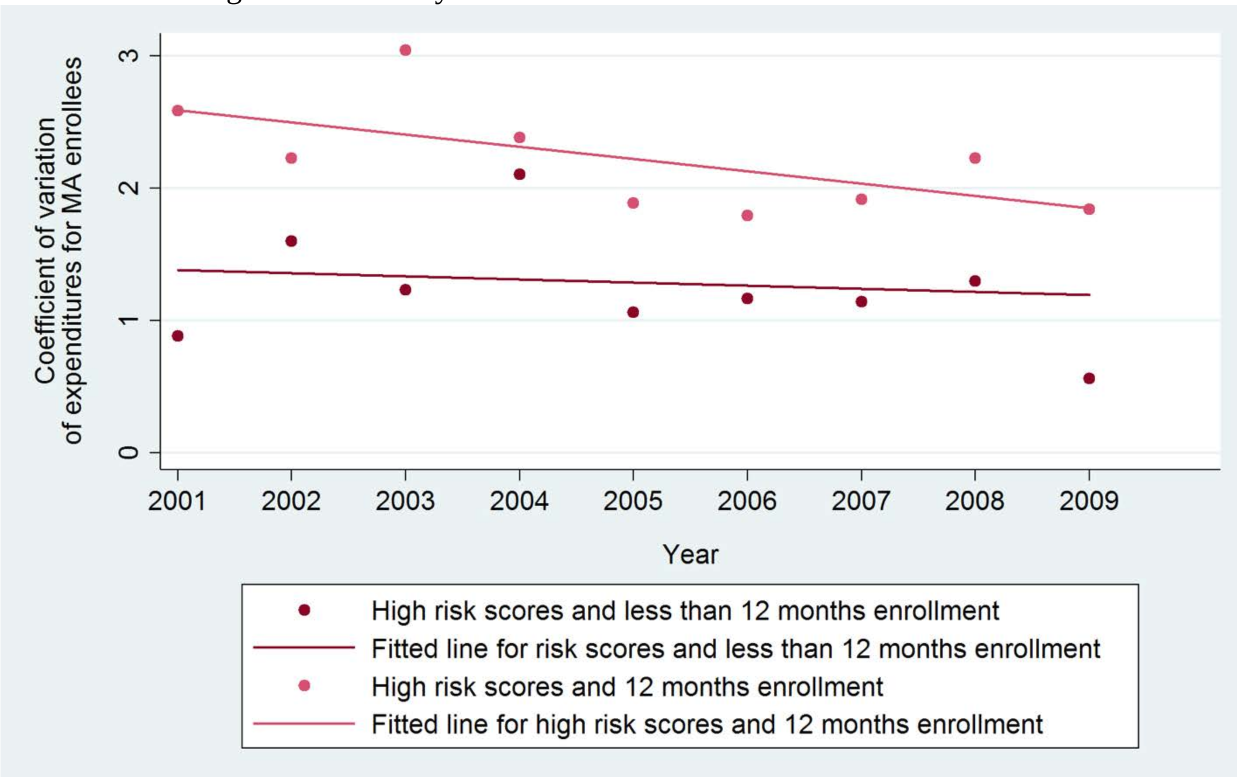


Figure 3. Coefficient of Variance of Total Medicare Expenditures for TM Enrollees and MA Enrollees by Risk Scores



Notes: The coefficient of variance was estimated as the ratio of the standard deviation to the mean. High-risk scores indicate above the 90th percentile of the risk score distribution.

Figure 4. Coefficient of Variance of Total Medicare Expenditures for TM Enrollees and MA Enrollees with High-risk Scores by Enrollment Periods



Notes: The coefficient of variance was estimated as the ratio of the standard deviation to the mean. High risk scores indicate above the 90th percentile of the risk score distribution.

APPENDIX:

Figure 1. Changes in Service-specific Use of Switchers to Stayers between the Pre- and Post-implementation Periods of the CMS-HCC Model (without Hospice with the Highest Service-level Selection Index)