

# Externalities and Taxation of Supplemental Insurance: A Study of Medicare and Medigap\*

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## Abstract

Most health insurance policies use cost-sharing to reduce excess utilization. The purchase of supplemental insurance can blunt the impact of this cost-sharing, potentially increasing utilization and exerting a negative externality on the primary insurance provider. This paper estimates the effect of private Medigap supplemental insurance on public Medicare spending using Medigap premium discontinuities in local medical markets that span state boundaries. Using administrative data on the universe of Medicare beneficiaries, we estimate that Medigap increases an individual's Medicare spending by 22.2%. We find that the take-up of Medigap is price sensitive with an estimated demand elasticity of -1.8. Using these estimates, we calculate that a 15% tax on Medigap premiums would generate combined tax revenue and cost savings of \$12.9 billion annually. A Pigouvian tax would generate combined annual savings of \$31.6 billion.

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# 1 Introduction

Health insurance policies typically include cost-sharing features such as coinsurance, copayments, and deductibles. By partially exposing beneficiaries to the marginal price of care, this cost-sharing strikes a balance between the risk-smoothing benefits of insurance and excess utilization from moral hazard (Zeckhauser, 1970). In some settings, individuals can purchase supplemental insurance, reducing their exposure to this cost-sharing and potentially exerting a negative externality on the primary insurance provider.

A leading example is the interaction between public Medicare insurance and private Medigap supplemental insurance. Most elderly Americans have health insurance through Medicare, which controls utilization with a deductible of approximately \$1,000 for each hospital admission and coinsurance of 20% for physician office visits.<sup>1</sup> In addition to these features, Medicare has no annual or lifetime out-of-pocket maximum, meaning that Medicare coverage leaves beneficiaries exposed to substantial out-of-pocket risk. Although most private insurance prohibits the purchase of supplemental insurance policies, Medicare allows its beneficiaries to purchase private supplemental insurance called Medigap. This supplemental insurance covers essentially all of the cost-sharing features of Medicare, potentially leading to excess utilization and exerting a negative externality on Medicare.<sup>2</sup> Taxing the purchase of Medigap to account for this externality may be a promising avenue for controlling rising Medicare costs—and increasing overall efficiency.

Researchers have long been aware that supplemental insurance may impose a fiscal externality on Medicare—and policymakers have issued a number of proposals to tax or regulate Medigap.<sup>3</sup> Yet despite this policy interest, considerable uncertainty remains about the effects of such reforms. Estimating the causal impact of Medigap is difficult because supplemental insurance coverage may be correlated with unobserved determinants of medical utilization. Previous studies that have examined this relationship with regressions of medical spending on Medigap status admit that adverse or advantageous selection may bias the results.

The primary objectives of this paper are (i) to provide a causal estimate of the externality

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<sup>1</sup>All dollar values are inflation-adjusted to 2005 values using the CPI-U. The Part A deductible was \$912 in 2005 and has been raised by \$27 nominal dollars on average per year since 2000.

<sup>2</sup>Because Medicare pays for a large fraction of the care done on the margin, if beneficiaries increase spending due to Medigap enrollment, then Medicare pays for a large fraction of this excess care.

<sup>3</sup>For example, President Obama's 2013 budget plan proposes levying a 15% tax on Medigap premiums.

Medigap imposes on the Medicare system and (ii) to estimate how a corrective tax on Medigap would impact Medicare costs and welfare. Our empirical strategy leverages Medigap premium discontinuities that occur at state boundaries. Medical costs, and thus the costs financed through supplemental Medigap insurance, exhibit considerable within-state variation due to geographic variation in factors ranging from household incomes to local physician practice styles to the supply of medical resources. Yet despite this local variation in the determinants of health care spending, within-state variation in Medigap premiums is very limited (Maestas, Schroeder and Goldman, 2009). This means that on opposite sides of state boundaries, otherwise identical individuals who belong to the same local medical market can face very different Medigap premiums solely due to state-level risk pooling in the Medigap market.

We focus our analysis on premium discontinuities within Hospital Service Areas (HSAs) that straddle state borders. Defined by the Dartmouth Atlas, HSAs are hospital catchment areas defined as sets of adjacent ZIP codes within which individuals go to the same hospitals for medical care. Approximately 250 of the 3,436 HSAs cross state lines, accounting for 11% of the individuals in our sample. We show that individuals who live on different sides of these cross-border HSAs are demographically alike and see the same providers for medical care. Since Medigap premiums tend to vary at the state level, individuals on different sides of these HSAs can face substantially different Medigap premiums and enroll in Medigap at sharply different rates.

An example is the HSA centered on Bennington, Vermont, which spans the border between southwest Vermont and upstate New York. On the Vermont side of the border, Medigap premiums are \$1,058 per year. On the New York side of the border, premiums are \$1,504 per year or about 40% higher—largely due to the high cost of medical care in New York City hundreds of miles to the south. More generally, our identification strategy exploits the fact that otherwise identical individuals in tightly defined HSAs can face sharply different Medigap premiums solely due to Medicare costs elsewhere in their state. We isolate this variation with a “leave-out costs” instrumental variable, which we define as the average uncovered Medicare spending for all Medicare beneficiaries outside an individual’s HSA but within his state of residence.<sup>4</sup> Leave-out costs differ by at least \$64 in 50% of cross-border HSAs and by at least \$166 in 20% of the cross-border HSAs

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<sup>4</sup>Throughout the paper, “uncovered Medicare spending” refers to the portion of Medicare-eligible spending that is the responsibility of the beneficiary and is paid either by the beneficiary or the beneficiary’s supplemental insurer.

in our sample. Our first stage regression of premiums on leave-out costs and HSA fixed effects is highly predictive with an R-squared ranging between 0.84 and 0.93 across the specifications and a p-value on the instrument of less than 0.01.

We use this variation in premiums to estimate the price sensitivity of Medigap demand. Our preferred instrumental variable estimates indicate a demand elasticity of -1.5 to -1.8. These estimates are stable across alternative specifications and different approaches to measuring Medigap coverage in our data. Our empirical strategy also allows us to examine potential substitution into alternative forms of coverage, and we find no evidence of substitution into Medicare Advantage or Medicaid based on our variation in premiums.

Using administrative data on the universe of Medicare beneficiaries, we use this same instrumental variables strategy to examine the impact of Medigap on medical utilization and Medicare costs. Our estimates can be interpreted as local average treatment effects for individuals who are marginal to variation in premiums—the same individuals who would respond to a tax on premiums. We find that Medigap increases Part B physician claims by 33.7% and Part A hospital stays by 23.9%. Summing across all categories of spending, we find that Medigap increases overall Medicare costs by \$1,396 per year on a base of \$6,290 or by 22.2%. This effect averages over individuals with higher spending due to moral hazard and any individuals with potentially lower spending due to increased use of preventative care (Chandra, Gruber and McKnight, 2010). We show our results are robust to alternative specifications, and we conduct several falsification tests using individuals and procedures that should be unaffected by the variation in premiums.

In the final part of the paper, we combine our demand and cost estimates to calculate the impact of taxing Medigap.<sup>5</sup> Our estimates indicate that a 15% tax on Medigap premiums, with full pass-through, would decrease Medigap coverage by 13 percentage points on a base of 48% and reduce net government costs by 4.3% per Medicare beneficiary, with a standard error of 1.7 percentage points. About 35% of this savings would come from tax revenue while the remainder would come from lower Medigap enrollment. A tax equal to the full \$1,396 externality requires us to extrapolate outside the premium variation in the data. To a first order approximation, our estimates indicate that such a tax would eliminate the Medigap market and decrease Medicare

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<sup>5</sup>Because our instrument affects Medigap enrollment through premiums, the Medigap externality that we calculate by combining our demand and cost estimates is precisely the policy-relevant externality needed for evaluating the effect of a tax.

costs by 10.7% per beneficiary. We conclude by discussing optimal Medigap taxation and welfare.

Our paper builds on an older literature that assesses the impact of Medigap with regressions of medical spending on an indicator for Medigap enrollment, controlling for selection into Medigap with available covariates. These papers—such as [Ettner \(1997\)](#), [Wolfe and Goddeeris \(1991\)](#), [Khandker and McCormack \(1999\)](#), and [Hurd and McGarry \(1997\)](#)—find that Medigap increases Medicare costs by about 25%, but admit that selection may bias these estimates.<sup>6</sup> Our paper is also related to [Chandra, Gruber and McKnight \(2010\)](#), which studies the effects of a change in the generosity of the retiree supplemental insurance provided to California state employees through the CalPERs system. The authors’ main finding is that CalPERs drug coverage can reduce hospitalizations among the chronically ill. These results do not have direct relevance to this setting because Medigap does not typically include drug coverage.<sup>7,8</sup>

Our paper contributes to the literature in a number of ways. First, to the best of our knowledge, our paper is the first to estimate the fiscal externality from Medigap using a quasi-experimental source of variation. Second, by using this variation to also estimate the price sensitivity of demand, we are able to quantify the cost savings and welfare effects of taxing Medigap. Third, many public insurance programs throughout the world allow policyholders to purchase private supplemental insurance.<sup>9</sup> Thus, we think that our approach can be applied to studying how to reduce costs and increase surplus from public insurance in a broad range of settings.

The remainder of the paper proceeds as follows. Section 2 describes the relevant institutional details of Medicare and Medigap insurance. The empirical strategy is outlined in Section 3. Sec-

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<sup>6</sup>[Lemieux, Chovan and Heath \(2008\)](#) argue that selection is probably adverse, leading these studies to overstate the impact of Medigap. [Finkelstein \(2004\)](#) finds evidence consistent with adverse selection in the Medigap market. [Fang, Keane and Silverman \(2008\)](#) find evidence of advantageous selection into Medigap, though this advantageous selection disappears once they condition on a wider set of covariates.

<sup>7</sup>Medigap policies currently sold exclude drug coverage. During our sample period (prior to the introduction of Part D), some Medigap policies included some prescription drug coverage, though nearly 90% of Medigap purchasers opted for policies that had no drug coverage (according to self-reported plan choice in the Medicare Current Beneficiary Survey; see the Appendix).

<sup>8</sup>As argued in [Goldman and Philipson \(2007\)](#), the degree of complementarity or substitutability between medical care and prescription drugs should influence optimal insurance design across the two domains. In the context of Medicare and supplemental coverage provided by employers (which often includes drug coverage), such complementarities may lead one to consider a more nuanced tax/subsidy policy to address any positive or negative externalities imposed by various features of supplemental insurance policies.

<sup>9</sup>In France, more than 92% of the population holds private supplemental insurance to protect against the substantial coinsurance payments (10% to 40%) of the universal public health insurance system. In Austria, about a third of the population has a supplemental private insurance plan that covers additional charges not covered under the basic health insurance benefits. About 30% of Belgians carry private supplemental health insurance policies. Approximately 30% of the population of Denmark purchases Voluntary Health Insurance (VHI) in order to cover the costs of statutory copayments of the universal health care coverage package. See [KFF \(2008\)](#) and [Cato \(2008\)](#) for more details.

tion 4 describes the data and the identifying variation. Section 5 presents the main results. Section 6 examines the robustness of these results to number of specification checks and placebo tests. Policy counterfactuals are presented in Section 7. Section 8 concludes.

## 2 Background

Medicare is the primary health insurer of individuals aged 65 and older in the United States. In 2012, Medicare covered 50 million beneficiaries at an annual cost of \$472 billion, accounting for about 13% of government expenditure and 3% of GDP.<sup>10</sup> By 2025, Medicare is projected to cover 71 million beneficiaries, consuming 20% of government expenditure and 5.5% of GDP.<sup>11</sup>

Medicare beneficiaries choose between publicly administered traditional fee-for-service (FFS) Medicare coverage and private Medicare Advantage policies.<sup>12</sup> Most Medicare beneficiaries (85%) select FFS Medicare coverage. The remaining 15% hold private Medicare Advantage policies which have premiums subsidized by Medicare and which generally offer more generous financial coverage of health care services in exchange for beneficiaries accepting a more restricted network of providers. In contrast, traditional FFS Medicare coverage allows beneficiaries their choice of doctor and the ability to see a specialist without a referral. To control costs, FFS Medicare uses cost-sharing, exposing beneficiaries to a large share of the cost of care received on the margin.

The details of FFS Medicare coverage for 2005 (the last year of our sample) are shown in Table 1. Beneficiaries face a deductible of nearly \$1,000 for each hospital admission and additional cost-sharing for long hospital stays. Medicare requires beneficiaries to pay 20% of all physician expenditures.<sup>13</sup> A key feature of FFS Medicare is that there is no annual or lifetime out-of-pocket maximum, so individuals are exposed to significant financial risk. Figure 1 shows the distribution of uncovered Medicare spending, which is defined as Medicare-eligible spending for which the patient is responsible.<sup>14</sup> The mean uncovered spending is \$1,186, and 3.8% of individuals in each

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<sup>10</sup>Aggregate information on Medicare net outlays and beneficiaries in 2012 are from CBO (2013).

<sup>11</sup>The percent of GDP numbers are gross of premiums; the percent of budget numbers are net (Trustees, 2009; OMB, 2010).

<sup>12</sup>We will refer to Medicare Part C as Medicare Advantage in this paper, although this option was called “Medicare + Choice” during the beginning of the period we analyze.

<sup>13</sup>As noted in Table 1, there is also a small annual deductible specific to Part B spending.

<sup>14</sup>Figure 1 is constructed using data from the CMS Beneficiary Summary File (1999-2005), and the sample is restricted to FFS Medicare beneficiaries who are not Medicaid recipients (non-dual eligible).

year have uncovered expenditures in excess of \$5,000.

To protect against the financial risk of FFS Medicare, the majority of FFS beneficiaries (86%) carry supplemental insurance.<sup>15</sup> Approximately 13% of FFS beneficiaries qualify for supplemental insurance at no cost through the government Medicaid program. Other beneficiaries may choose to purchase supplemental insurance offered by a former employer, and everyone has the option to purchase private Medigap coverage. Among FFS beneficiaries, 42% purchase Medigap coverage, and approximately 40% purchase supplemental insurance through a former employer.<sup>16</sup>

The federal government regulates both the form of Medigap insurance and the purchase of Medigap policies. Individuals are restricted to choose from a standardized set of plans, all of which cover the same basic benefits.<sup>17</sup> These basic benefits include coverage of the Part A deductible, Part A copays, and Part B coinsurance. Beyond the basic benefits, there is some variation across plans in the remaining coverage, though most of this variation is for less common expenses such as travel emergencies and home health care. Appendix A shows enrollment by plan and discusses Medigap plan characteristics in detail.<sup>18</sup>

In this paper, we focus on the extensive margin of whether an individual has Medigap, rather than the effect of one plan compared to another, for two reasons. First, the basic benefits that are likely to have the greatest effect on the marginal price of care are common to all plans. Thus, the extensive margin substitution into Medigap is likely to be the primary driver of the marginal cost of care.<sup>19</sup> Second, our aim is to investigate the effect of a tax on Medigap policies. Because the Medigap tax proposals under consideration do not discriminate across plans, the extensive

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<sup>15</sup>Although Medicare Advantage beneficiaries can technically sign up for supplemental insurance policies, this is discouraged by Medicare, and individuals do not seem to do this in practice. Medicare Advantage plans tend to give more financial coverage in exchange for a more restricted network of providers relative to FFS Medicare. Since supplemental insurance policies are tailored to fill gaps of FFS Medicare, the supplemental coverage that these policies would provide an individual on Medicare Advantage would be largely redundant.

<sup>16</sup>According to the MCBS estimates, approximately 10% of FFS beneficiaries carried both Medigap and Retiree Supplemental Insurance coverage during our sample period.

<sup>17</sup>There are three states in which the Medigap market is different. Massachusetts, Wisconsin, and Minnesota standardized their plans prior to federal regulation and have continued their own offerings. We exclude these three states from our analysis. The Medicare Prescription Drug, Improvement, and Modernization Act of 2003 introduced plans K and L and eliminated the sale of Medigap plans with drug benefits (H, I, and J). These changes took effect after our sample period.

<sup>18</sup>Plans C and F, by far the most popular plans, are chosen by more than 60% of the Medigap beneficiaries. Both plans cover the hospital and physician deductibles, in addition to the basic benefits common to all standardized plans.

<sup>19</sup>Because elderly individuals almost all have annual physician expenditures exceeding the \$110 Part B deductible, the relevant marginal price for physician visits is the 20% coinsurance. Although Part B deductible coverage is available only for selected plans, this variation in coverage is unlikely to drive the marginal price of care since so many people are inframarginal with respect to the Part B deductible.



margin substitution is more policy-relevant than substitution across Medigap plans.

In addition to regulating the form of Medigap policies, the federal government regulates the purchase of policies. Medigap beneficiaries typically purchase Medigap insurance within six months of turning 65 years old and signing up for FFS Medicare, during what is called the “open enrollment period.” Medigap policies purchased during this open enrollment period are guaranteed renewable as long as Medigap enrollees pay plan premiums each year.<sup>20</sup> Individuals in this market typically sign up for a Medigap plan during their open enrollment period, and renew their policy each year.<sup>21</sup> During this open enrollment period, individuals cannot be legally denied coverage for any reason, and pricing is limited to a small set of characteristics: the gender, location, and smoking status of the enrollee. In practice, premium variation is much more limited than what is legally allowed. De facto, companies rarely vary premiums for a given plan within a state.<sup>22</sup> The beneficiary-weighted average annual premium of Medigap policies is \$1,779, though the premium varies substantially across states.<sup>23</sup> In Section 4, we discuss the Medigap premium variation in more detail.

Some individuals obtain supplemental coverage through a former employer.<sup>24</sup> Unlike Medigap coverage during our sample period, Retiree Supplemental Insurance (RSI) policies typically covered prescription drugs and provided less generous coverage (or sometimes no coverage) of medical services.<sup>25</sup> According to the Kaiser Family Foundation (KFF, 2004), the average annual premium for an individual RSI policy in 2004 was \$3,144, and retirees on average contributed approximately 39% or \$1,212 of this premium. Unlike individual Medigap policies, RSI coverage is

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<sup>20</sup>The federal government regulates how Medigap policy prices can evolve. In particular, when an individual enrolls in a Medigap plan, he is choosing an age-price profile that may be adjusted with medical inflation but may not be contingent on his current or future health status. Thus, along with the contemporaneous benefits, Medigap coverage provides insurance against reclassification risk in future periods. Since the evolution of premiums over time is set by federal standards, throughout the paper we focus on the premium charged to a 65-year-old during the open-enrollment period.

<sup>21</sup>Medicare’s website ([www.Medicare.gov](http://www.Medicare.gov)) provides beneficiaries with information on selecting a Medigap policy and encourages beneficiaries to select a policy as if they will annually renew the policy because dropping coverage would mean they would face risk-rating were they to wish to re-enroll. Patterns in the available data suggest that seniors tend to renew policies year-to-year. According to the authors’ calculations, the short panel available in the Medicare Current Beneficiary Survey reveals that 87% of seniors who were on Medigap in the prior year continue to be enrolled in Medigap during the current year.

<sup>22</sup>In practice, smoking status and gender are rarely priced. Although plans are legally allowed to vary prices at the ZIP code level, in practice there tends to be very limited variation in company-plan level premiums within a state.

<sup>23</sup>The beneficiary-weighted premium is calculated using the baseline sample, as described in Table 2.

<sup>24</sup>In 2005, 32% of employers with more than 200 employees offered some form of Medicare supplemental insurance (KFF, 2004).

<sup>25</sup>According to the Kaiser Family Foundation, 98% of RSI policies offered in 2004 had some prescription drug coverage. This may have changed after the introduction of Part D.



often available to both the retiree and his or her spouse.<sup>26</sup>

### 3 Empirical Strategy

#### 3.1 Overview

Our empirical approach is to use exogenous variation in Medigap premiums to identify (i) the price sensitivity of the demand for Medigap and (ii) the fiscal externality of Medigap on Medicare costs. The variation in premiums directly allows us to identify the effect of premiums on Medigap demand. Because Medigap take-up is price sensitive, this premium variation also provides us with an instrument for Medigap coverage, allowing us to identify the impact of Medigap on Medicare costs. The local average treatment effect we identify is the effect of Medigap for individuals marginal to variation premiums (Imbens and Angrist, 1994), the same individuals who would be influenced by a counterfactual tax on Medigap premiums.

The exogenous variation we use arises from geographic discontinuities in premiums that occur at state boundaries. Medical costs exhibit considerable within-state variation due to factors ranging from household incomes to local physician practice styles to the supply of medical resources (Cutler and Sheiner, 1999; Wennberg, 1999; Wennberg, Fisher and Skinner, 2002; MedPAC, 2003). Yet despite this local variation, within-state premium variation is highly limited (Maestas, Schroeder and Goldman, 2009).<sup>27</sup> This means that on opposite sides of state boundaries, otherwise identical individuals who belong to the same local medical markets can face very different premiums for Medigap.

We focus our analysis on premium discontinuities within HSAs that span state borders. HSAs are defined by the Dartmouth Atlas as sets of adjacent ZIP codes in which residents receive most of their routine hospital care at the same facility.<sup>28</sup> HSAs are approximately the size of a county: there are 3,436 HSAs and 3,140 counties in the United States. However, unlike counties, HSAs

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<sup>26</sup>When spousal coverage is available, the average premium for a policy that covers the retiree and spouse is roughly double that for individual coverage (KFF, 2004).

<sup>27</sup>Although firms are allowed to vary premiums at the ZIP code level, Maestas, Schroeder and Goldman (2009) find there is very little within-state variation in the Medigap premiums for a given plan offered by a given insurance company. Maestas, Schroeder and Goldman (2009) cite state-level reporting requirements and regulations as a potential explanation. Because plans must, for example, meet loss ratio requirements at the state level, varying premiums more locally may be administratively burdensome.

<sup>28</sup>The Dartmouth Atlas data are constructed from the Continuous Medicare History Sample (CMHS), a 5% sample of the billing records of Medicare beneficiaries collected by the Centers for Medicare and Medicaid Services (CMS).

often span state boundaries, reflecting the fact that local medical markets are not aligned with political boundaries. There are several advantages to using HSAs to define local medical markets as opposed to a market definition based solely on distance to a state boundary. First, we identify the effect of Medigap among those individuals who receive care from the same medical providers. Second, focusing on within-HSA premium variation allows us to ignore those state border areas where geographic barriers, or sharp differences in socioeconomic factors, lead to natural breaks in the providers from which medical care is received.

Figure 2 provides a concrete example of our empirical strategy. Panel A shows a map of per capita uncovered Medicare spending in New York and Vermont by HSA; Panel B shows Medigap premiums in the same area. We define “uncovered Medicare spending” as the Medicare-eligible spending that is the responsibility of the beneficiary and is paid for either out-of-pocket or by a supplemental insurance plan. Two HSAs, centered on Bennington, VT, and Cambridge, NY, straddle the New York-Vermont border. Each of these HSAs had average per capita uncovered Medicare spending around \$900, typical of the other HSAs in the upstate NY and VT area.<sup>29</sup> However, within these cross-border HSAs, there are sharp differences in Medigap premiums. Premiums on the New York side of the border are \$1,504 per year versus \$1,058 on the Vermont side.<sup>30</sup> The reason for this premium difference is that New York state has New York City in the south, a region with substantially higher Medicare costs than the northern part of the state.<sup>31</sup> It is the high-spending metropolitan south, combined with the limited within-state variation in premiums, which inflates Medigap premiums in upstate New York, creating the source of premium variation.

Figure 3 shows these same data for the continental United States. Panel A shows a map of HSA-level per capita uncovered Medicare spending; Panel B shows average Medigap premiums. It is well known that there are vast regional differences in medical spending. Across states, average per capita spending not covered by Medicare ranges from \$763 per capita in Oregon to \$1,123 per capita in New York. A fact that is less well known, but vital to our estimation strategy, is that much of the variation in medical spending is across local medical markets within states. Per

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<sup>29</sup> Authors’ calculations use data from the CMS Beneficiary Summary File for 2000. Per-capita uncovered Medicare spending in 2000 was \$902 and \$927 in the Bennington HSA and Cambridge HSA, respectively.

<sup>30</sup> The average premium cited here is the average premium of all plans offered by all Medigap insurers in the year 2000 (adjusted to be in 2005 dollars).

<sup>31</sup> The maximum HSA-level uncovered Medicare spending is \$1,585 in the south versus \$1,087 in upstate NY.

capita spending not covered by Medicare varies by more than \$740 within New York, Florida, and Texas, and by more than \$250 within the 25 most populous states that are collectively home to the majority (83%) of the individuals in our sample.<sup>32</sup> As a result, there are many cross-border HSAs with large differences in Medigap premiums. Like the case of New York and Vermont, many of these differences are driven by within-state variation in costs outside the relevant HSA. This is precisely the variation that we isolate with the “leave-out-cost” instrument that we define below.

### 3.2 Estimating Equations

Given this cross-border variation, one way to estimate the demand for Medigap is to run regressions of Medigap coverage on premiums in the sample of cross-border HSAs. However, this approach would not identify the causal effect of premiums on demand because premiums in border-spanning HSAs are partially determined by the behavior of individuals within these cross-border medical markets.<sup>33</sup> For example, high utilization among those on the Vermont side of the Cambridge, New York, HSA would impact average spending in Vermont and therefore the Medigap premiums faced by those on the Vermont side of the Cambridge, New York, HSA. To address this concern, we instrument for premiums in an HSA with the average uncovered Medicare spending of individuals who reside outside the local medical market (HSA) but within the same state. This leave-out approach is similar to that used in other empirical studies (e.g., [Chetty et al., 2011](#)) to purge instruments of mechanical sources of correlation with the outcome of interest.

Let  $i$  denote individuals,  $j$  denote states, and  $k$  denote HSAs. Assume, to a first approximation, that Medigap premiums in a given state,  $p_j$ , are proportional to the uncovered Medicare spending of individuals within that state,  $p_j = \alpha \mathbb{E}_{i \in I_j}[c_i^u]$ , where  $c_i^u$  is the uncovered Medicare spending of individual  $i$  and the expectation is taken over  $I_j$ , the set of individuals in state  $j$ . For a HSA-state pair, we can decompose the determinants of premiums into the uncovered spending of individuals within and outside the given HSA:  $p_{jk} = \alpha \Pr[i \in I_{j,k} | i \in I_j] \times \mathbb{E}_{i \in I_{j,k}}[c_i^u] + \alpha \Pr[i \in I_{j,-k} | i \in I_j] \times \mathbb{E}_{i \in I_{j,-k}}[c_i^u]$ , where  $I_{j,k}$  denotes the set of individuals in state  $j$  and HSA  $k$ . Since the uncovered spending of individuals within the HSA is potentially endogenous, we define our leave-out cost

<sup>32</sup> Authors’ calculations use data from the CMS Beneficiary Summary File for 2000.

<sup>33</sup> Although this endogeneity may be problematic in theory, in practice the endogeneity problem shrinks substantially as we narrow the focus of the analysis to those in very close proximity to the boundary who make up a very small fraction of any state.

instrument as the average uncovered Medicare spending of those outside the HSA but within the state of interest:

$$\text{Leave-out costs}_{jk} = \Pr[i \in I_{j,-k} | i \in I_j] \times \mathbb{E}_{i \in I_{j,-k}}[c_i^u]. \quad (1)$$

Given this instrument, the first stage regression of premiums on the leave-out costs is given by:

$$p_{jk} = \alpha_c \text{Leave-out costs}_{jk} + \alpha_k + X'_{jk} \alpha_X + \epsilon_{jk}, \quad (2)$$

where  $\alpha_k$  is a vector of HSA fixed effects,  $X_{jk}$  are covariates, and  $\epsilon_{jk}$  is the error term. Including HSA fixed effects means that the coefficient on leave-out costs  $\alpha_c$  is identified by variation in the instrument within HSAs that span state boundaries.

In addition to varying by state, Medigap premiums differ by plan letter and insurance provider, which are attributes that are not observed in our data. We therefore estimate the premium sensitivity of demand using a two-sample IV approach. Let  $q_{ijk}$  be an indicator that takes a value of one if the individual reports having Medigap and zero otherwise. The reduced form regression takes the form:

$$q_{ijk} = \beta_c \text{Leave-out costs}_{jk} + \beta_k + X'_{ijk} \beta_X + v_{ijk}, \quad (3)$$

where  $\beta_k$  is a vector of HSA fixed effects,  $X_{ijk}$  are covariates, and  $v_{ijk}$  is the error term. The implied instrumental variable impact on Medigap enrollment of an increase in premiums is given by the ratio of the reduced form and first stage coefficients:  $\beta_c / \alpha_c$ . We can explore the sensitivity of our results by using  $\alpha_c$ 's from regressions of different premium measures on the instrumental variable.

We estimate the effect on utilization of Medigap using the two-sample IV approach explained above. The reduced form regression of a measure of utilization  $y_{ijk}$  on the instrument is:

$$y_{ijk} = \gamma_c \text{Leave-out costs}_{jk} + \gamma_k + X'_{ijk} \gamma_X + \mu_{ijk}, \quad (4)$$

where  $\gamma_k$  are HSA fixed effects,  $X_{ijk}$  are covariates, and  $\mu_{ijk}$  is the error term. The implied instrumental variable effect on utilization of Medigap is given by  $\gamma_c / \beta_c$ . The implied instrumental variable impact on utilization of an increase in Medigap premiums is given by  $\gamma_c / \alpha_c$ . To account

for the fact that determinants of medical care may be related within local medical markets, we calculate robust standard errors clustered at the HSA level in each stage of the estimation.

## 4 Data and Identifying Variation

### 4.1 Data

We use data from several sources. The primary medical spending and utilization information comes from Medicare administrative data obtained from the Centers for Medicare and Medicaid Services (CMS) and covers the years 1999 through 2005. The CMS Denominator file contains administrative data on the universe of Medicare enrollees, and includes information on sex, age, Medicaid status, Medicare Advantage enrollment, and ZIP code of residence. To investigate beneficiary-level spending and utilization, we combine the CMS Denominator file with the CMS Beneficiary Summary File which covers the universe of Medicare FFS beneficiaries. The Beneficiary Summary File data contains information on health care spending (Medicare spending and beneficiary spending), utilization by category of care (e.g., hospitalizations, Part B claims), and chronic conditions.<sup>34</sup>

To further investigate which types of utilization are elastic to Medigap enrollment, we also examine Medicare claims data. Outpatient claims data are available in the CMS Carrier data file that contains outpatient claims for a 20% random sample of FFS Medicare beneficiaries. Inpatient claims data are available in the CMS MedPAR data file which contains inpatient claims for 100% of FFS Medicare beneficiaries.

The CMS administrative data do not contain information on Medigap enrollment.<sup>35</sup> Thus, we must rely on survey data to estimate the demand for Medigap.<sup>36</sup> To maximize statistical power, we combine estimates from two comparable surveys: the Medicare Current Beneficiary Survey (MCBS) from 1992 to 2005 and the National Health Interview Survey (NHIS) from 1992 to 2005. Both surveys ask questions regarding supplemental insurance coverage among Medicare benefi-

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<sup>34</sup>Data on spending, utilization, and chronic conditions are available only for FFS Medicare beneficiaries (no data are available for those on Medicare Advantage). Thus, it is key that we show that individuals do not substitute to Medicare Advantage to be able to interpret our results.

<sup>35</sup>The lack of CMS data on Medigap is perhaps not surprising since Medigap enrollment does not affect Medicare's reimbursement formulas so claims can be processed without this information.

<sup>36</sup>Because the Medigap variable and spending outcomes are not available in the same dataset, we use a two-sample IV approach and estimate reduced form specifications as detailed in the prior section.

ciaries and contain similar demographic and health information. Appendix B describes how we construct the key variables from each survey.

Premium data come from Weiss Ratings.<sup>37</sup> The premium data contain Medigap premiums for policies purchased during the open-enrollment period for 2000. Prior work reveals that within-state premium variation in plan-level Medigap premiums is very limited (Robst, 2006; Maestas, Schroeder and Goldman, 2009). In practice, firms do not tend to vary premiums across localities within a state, and firms rarely price gender or smoking status. For the analysis in this paper, we use premium data aggregated to the state-plan-firm level. In Section 5, we demonstrate that our instrument is a powerful predictor of premiums.

Geographic crosswalks from the Dartmouth Atlas are used to match localities with their associated local medical markets (HSAs).<sup>38</sup> We also merge supplemental data from several other sources to implement the analysis. ZIP code-level demographic covariates are obtained from the Census of Population and Housing 2000, Special Tabulation on Aging (available through ICPSR). Another key geographic control included in the estimation are the geographic adjustment factors contained in Medicare’s provider reimbursement formulas. Although our analysis looks at individuals within local medical markets who (by definition) tend to use the same providers, we also control for any mechanical reasons that provider reimbursement may vary geographically as people may disproportionately use nearby providers for basic medical care. Details on Medicare provider reimbursement formulas are obtained from CMS.<sup>39</sup>

## 4.2 Summary Statistics

Our baseline sample consists of the universe of continuously enrolled FFS Medicare beneficiaries excluding those who are simultaneously enrolled in Medicaid and excluding those who qualify for Medicare before age 65 due to disability. In Section 5, we demonstrate that individuals do not substitute into Medicaid or Medicare Advantage based on the premium variation within cross-

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<sup>37</sup>We thank John Robst for sharing these data.

<sup>38</sup>We exclude the District of Columbia from our analysis because more than 99% of the individuals in this region belong to a single HSA. We also exclude beneficiaries from the three states that do not have standardized Medigap products: Wisconsin, Massachusetts, and Minnesota. Lastly, we exclude a small number of HSAs where the remainder of the state accounts for less than 80% of the sample population since the leave-out costs instrument for these HSAs are extreme outliers.

<sup>39</sup>Medicare provider reimbursement varies across geography in part because of formulaic geographic adjustments. Our baseline analysis controls for these mechanical sources of variation using Medicare reimbursement formulas obtained from CMS. Specifically, our analysis controls for the Part B GAF and Part A OWI adjustment factors.

border HSAs. This lack of substitution into Medicaid or Medicare Advantage means that our baseline sample is valid for estimating of the effect of Medigap.

Table 2 presents summary statistics for our data. The first column displays summary statistics for the full sample. Three-quarters of Medicare beneficiaries have traditional FFS coverage without Medicaid coverage, 14.5% have coverage from a Medicare Advantage plan, and 11.0% are dual-eligibles with coverage from both Medicare and Medicaid. Within the baseline sample of FFS non-Medicaid beneficiaries, 47.9% hold a Medigap policy, 46.3% hold an RSI policy, and 15.8% have no supplemental coverage. These numbers sum to greater than 100% as some individuals report having both Medigap and RSI coverage. Medigap premiums have a mean value of \$1,779 per year. Within the baseline sample, total Medicare payments average \$6,290, and approximately 56% of payments are for inpatient care. On average, Medicare beneficiaries spend two days in a hospital annually and have 26 Part B events, where an event is defined as a line-item claim.

The second column of Table 2 presents the same summary statistics for the 11% of beneficiaries who reside in HSAs that span state boundaries. This sample is of particular interest as variation in our instrument among these individuals identifies the demand and utilization elasticities. The baseline FFS non-Medicaid sample is about 8 percentage points larger and Medicare Advantage enrollment is 8 percentage points lower in the cross-border sample. This is likely because these regions are more rural and Medicare Advantage penetration was lower in rural areas during our time period. The percent of dual-eligibles is virtually identical in the full and cross-border samples. The cross-border sample is very similar in terms of enrollment in supplemental insurance and demographic information (age, sex, and race). The average Medigap premium among those in cross-border HSAs is 3% lower than in the full sample. In the cross-border sample, Part A days are 2% less and Part B events are 7% less than in the full data. Taken together, these statistics indicate that the border-spanning sample is broadly similar to the full sample of Medicare beneficiaries.

### 4.3 Identifying Variation

Figure 4 illustrates the identifying variation, plotting a histogram of the leave-out costs instrument in cross-border HSAs net of the mean of the instrument within each HSA. The instrument



is constructed using data on the baseline sample from the 2000 CMS Beneficiary Summary File.<sup>40</sup> Leave-out costs exhibit substantial dispersion, with an interquartile range of \$64 and a 90-10 percentile range of \$166. This implies a jump of at least \$64 in 50% of the cross-border regions, or 7.2% of the mean leave-out cost value in cross-border HSAs of \$886. In 20% of the regions, there is a jump of at least \$166 or 18.7% of the mean.

The identification assumption is that the within-HSA variation in leave-out costs affects the dependent variable of interest (e.g., Medigap enrollment, medical utilization) only through Medigap premiums. Although we cannot test this assumption directly, we provide several pieces of empirical evidence that help to make the case that the identifying assumption holds. Below, we show that the instrument does not covary with individual and local characteristics (potential omitted variables) within cross-border HSAs. In Section 6, we further examine the robustness of our results by (i) examining the stability of the estimates when we control for potential confounding factors and (ii) conducting falsification tests on outcomes and individuals that should not be affected by our source of variation.

To examine the correlation between local characteristics and the instrument, we estimate regressions of the form:

$$w_{zjk} = \delta_c \text{Leave-out costs}_{jk} + \delta_k + v_{zjk}, \quad (5)$$

where  $w_{zjk}$  are local characteristics,  $\delta_k$  are HSA fixed effects, and  $v_{zjk}$  is the error term. Table 3 shows the results of these regressions. The dependent variables are ZIP code-level demographics from the Census 2000 Special Tabulation on Aging. The results reveal that within cross-border HSAs, the leave-out cost instrument is largely unrelated to characteristics of the elderly population such as education, veteran status, labor force participation, income, and relocation. Nearly all of the ZIP code-level Census demographics have a statistically insignificant relationship with the leave-out cost instrument.<sup>41</sup> Using data on the universe of Medicare beneficiaries from the CMS Denominator file, we similarly investigate whether our identifying variation is related to Part B

<sup>40</sup>We use Beneficiary Summary File data from 2000 because our premium data are also from this year.

<sup>41</sup>Although there is one exception (the coefficient on Veteran Status among Males 65+), the reported standard errors are not corrected for multiple hypothesis testing, and if we were to do so, many corrections would lead us to conclude that we could not reject the hypothesis that all the coefficients are statistically indistinguishable from zero. For example, a simple Bonferroni correction for multiple hypothesis testing would mean that the effective p-values should be multiplied by the number of hypotheses we are testing (in this case, 13). Thus, the corrected p-value on the coefficient on Veteran Status among Males 65+ would be 0.39.

coverage rates or to the fraction of individuals originally qualifying for Medicare through SSDI. The results in Table 3 reveal that neither is related to our identifying variation.

We also aggregate across these Census demographic variables by examining the correlation between an individual’s predicted level of medical spending and the instrumental variable, where the predicted level of spending is the fitted value from an OLS regression of individual-level Medicare spending on these Census demographic variables and the controls in our baseline specification. The results are reported in the final line of Table 3. The estimated coefficient from this exercise is statistically indistinguishable from zero with a p-value of 0.537. Overall, this evidence suggests that observables plausibly related to medical spending are unrelated to the identifying variation.

## 5 Results

This section presents the baseline estimates. We start by showing that the leave-out cost instrument is a powerful predictor of premiums. We then use variation in leave-out costs to estimate the demand for Medigap and the effect of Medigap on Medicare utilization and spending.

### 5.1 Premiums

Table 4 presents estimates of the first stage regression of premiums on the leave-out costs instrument, HSA fixed effects, and controls (see Section 3, Equation 2). The first column displays results for a plan-level specification that includes all plans offered by United Healthcare and Mutual of Omaha, the two largest insurance companies with a combined market share of 69%.<sup>42</sup> The second and third columns restrict attention to the most popular plans sold by these insurance companies, Plan C and Plan F. The coefficient on the instrument ranges from 1.12 to 0.93 across specifications, indicating that the instrument shifts premiums on an approximately one-for-one basis. The coefficient on the instrument is precisely estimated with p-values of less than 0.01 across the specifications. The specifications explain much of the premium variation within cross-border HSAs, with the specifications having an R-squared ranging from 0.84 to 0.93.

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<sup>42</sup>This number is taken from [Starc \(2010\)](#), which summarizes data from the National Association of Insurance Commissioners.

## 5.2 Demand

The demand estimation proceeds in two stages. First, CMS administrative data are used to show that variation in leave-out costs does not induce substitution into Medicaid or Medicare Advantage plans.<sup>43</sup> Second, survey data are used to estimate Medigap demand. We are required to use survey data for these estimates since Medigap is not recorded in the CMS administrative data. See Section 4 for more information about the data.

### 5.2.1 Alternative Coverage: Medicare Advantage and Medicaid

We examine the potential for substitution into Medicare Advantage and Medicaid coverage with regressions of coverage indicators on the leave-out costs instrument, HSA fixed effects, and controls. That is, we estimate the reduced form specification for Medigap demand replacing the dependent variable with indicators for Medicare Advantage or Medicaid (see Section 3, Equation 3). The data are the pooled 1999 to 2005 CMS Denominator File, which contains information on the universe of Medicare beneficiaries over this time period.

Table 5 presents the results of these regressions. The leave-out costs instrument does not have a perceptible effect on either Medicare Advantage or Medicaid coverage. The point estimates indicate that a \$100 increase in leave-out costs reduces Medicare Advantage by 0.8 percentage points on a base of 12.3%, and raises Medicaid coverage by 0.7 percentage points on a base of 11.3%. Both estimates are statistically indistinguishable from zero with p-values of 0.23 and 0.19, respectively.

These results are not surprising given the institutional setting. Medicaid provides supplemental insurance, of similar generosity as Medigap, to poor beneficiaries for no premium. So it would be strange if variation in Medigap premiums had an impact on Medicaid coverage. During the time period we analyze, Medicare Advantage plans were typically organized as Health Maintenance Organizations (HMOs), which place significant restrictions on provider choice.<sup>44</sup> The lack of substitution into Medicare Advantage is consistent with other evidence on limited substitution between HMO and FFS insurance plans (e.g., [Bundorf, Levin and Mahoney, 2012](#)). Given this

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<sup>43</sup>The CMS Denominator file records beneficiary Medicare Advantage and Medicaid status for all Medicare beneficiaries.

<sup>44</sup>The FFS Medicare-Medigap pairing places virtually no restrictions on provider choice.

evidence, the utilization and cost analysis focuses on Medicare beneficiaries who are not enrolled in Medicare Advantage or Medicaid—the baseline sample that we refer to as FFS, non-Medicaid Medicare beneficiaries.

### 5.2.2 Supplemental Coverage: Medigap

We estimate the demand for Medigap with regressions of coverage indicators on the leave-out costs instrument, HSA fixed effects, and controls (see Section 3, Equation 3). We use data from two surveys, the 1992 to 2005 Medicare Current Beneficiary Survey (MCBS) and the 1992 to 2005 National Health Interview Survey (NHIS). The baseline sample in the MCBS has 114,561 observations and the baseline sample in the NHIS has 121,009 observations over our time period.

Our ability to precisely measure Medigap coverage varies across the datasets. In the MCBS, we have a relatively accurate measure of Medigap coverage, and we use this measure as an outcome variable. In contrast, the NHIS survey questions make it more difficult to distinguish Medigap from other forms of supplemental insurance.<sup>45</sup> We therefore estimate the effect in the NHIS using a broader measure of supplemental insurance that captures whether the individual has any supplemental insurance, including Medigap but also Medicare Advantage and Medicaid. Because our results using the administrative data indicate that the identifying variation does not cause substitution into Medicare Advantage or Medicaid, we interpret the effect on the broad measure as reflecting the response of Medigap coverage to leave-out costs.<sup>46</sup>

In both surveys, our estimates are identified by cross-border HSAs in which we observe individuals on both sides of the state border. Of the 259 total cross-border HSAs, we observe individuals on both sides of state borders in 27 HSAs in the MCBS and 37 HSAs in the NHIS. This means that the HSA-level estimates are identified using 2,903 of the 114,561 observations in the MCBS and 5,690 of the 121,009 observations in the NHIS. To increase the precision of our estimates, we also estimate the same specifications using a more aggregate definition of local medical

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<sup>45</sup>The MCBS survey asks several questions regarding the source of coverage that we can use to cross-validate responses. In addition, the MCBS makes some effort to check Medicare Advantage and Medicaid enrollment against administrative records. In contrast, the NHIS contains very few questions regarding sources of coverage, and responses are not checked against administrative records.

<sup>46</sup>Our broader measure of Medigap includes any supplemental insurance including Medigap, Medicare Advantage, Medicaid, or RSI. Our analysis with the administrative data reveals that there is no substitution into Medicaid or Medicare Advantage based on our premium variation. The prior literature has traditionally assumed there is no substitution between Medigap and RSI, and the results presented in Table 5 are consistent with no substitution into RSI based on our variation.

markets called a Hospital Referral Region (HRR). The Dartmouth Atlas defines an HRR as the set of adjacent ZIP codes in which individuals use the same hospitals for major medical care (such as cardiovascular surgery). While there are 3,436 HSAs across the nation, there are only 306 HRRs. Of the 140 total cross-border HRRs, we have observations on opposite sides of state borders in 66 HRRs in the MCBS and 70 HRRs in the NHIS. In these HRR-level specifications, the estimates are identified by 32,915 of the 114,561 observations in the MCBS and 39,060 of the 121,009 observations in the NHIS.<sup>47</sup>

Table 5 presents the results of these regressions.<sup>48</sup> The estimates in the MCBS indicate that a \$100 increase in leave-out costs reduces Medigap demand by 6.6 to 9.0 percentage points. The estimates are similar whether we use variation at the HSA or HRR level. The results are robust whether we measure Medigap coverage using the narrow Medigap coverage variable or the broader measure of supplemental insurance coverage. In the NHIS, where we have only the broader measure of Medigap coverage, we find that a \$100 increase in leave-out costs lowers Medigap coverage by 0.9 to 3.0 percentage points depending on whether we use variation at the HSA or HRR level.

Our preferred estimates combine the point estimates from the MCBS and the NHIS using the Delta Method to construct the appropriate standard errors.<sup>49</sup> These estimates indicate that a \$100 increase in leave-out costs reduces our broad measure of Medigap by 3.9 to 4.8 percentage points. The HSA level estimate is statistically distinct from zero with a p-value of 0.03, and the HRR level estimate is statistically distinguishable from zero with a p-value of 0.01. Given that within the baseline sample the Medigap market-share is 47.9% in the MCBS and the mean inflation-adjusted

<sup>47</sup>The reported demand coefficients in Table 5 for the HRR-level specifications are scaled by the premium first stage at the HRR level to be comparable with the HSA-level coefficients. See Table 5 for details.

<sup>48</sup>Appendix C illustrates that the demand estimates are robust to inclusion of fewer or more controls than in these baseline specifications. The baseline specifications in Table 5 include year fixed effects, local medical market fixed effects, basic demographic controls, and controls for geographic price indexes (GAF and OWI).

<sup>49</sup>Let  $\beta_i$ ,  $se_i$  and  $n_i$  denote the point estimate, standard error, and sample size in dataset  $i$ . The combined point estimate is constructed as the sample-size weighted average of the point estimates in the two samples:

$$\beta_{Combined} = \frac{n_{MCBS}\beta_{MCBS} + n_{NHIS}\beta_{NHIS}}{n_{MCBS} + n_{NHIS}}$$

Using the Delta Method and assuming that the point estimates are uncorrelated, the standard error of the combined estimate is given by:

$$se_{Combined} = \frac{\sqrt{n_{MCBS}^2 se_{MCBS}^2 + n_{NHIS}^2 se_{NHIS}^2}}{n_{MCBS} + n_{NHIS}}$$

premium is \$1,779, these estimates translate into a demand elasticity of -1.5 to -1.8.

### 5.3 Utilization and Spending

We examine the effect on utilization and spending with regressions of these measures on the leave-out costs instrument, HSA fixed effects, and controls (see Section 3, Equation 4). We restrict the sample to FFS Medicare, non-Medicaid beneficiaries. The main source of data is the pooled 1999 to 2005 Beneficiary Summary Files, which provide us with annual beneficiary-level cost and utilization data on the universe of Medicare beneficiaries over this time period. We also use the 1999 to 2005 Carrier File for analysis that requires claim-level data. For these data, we have information on a randomly selected 20% sample of the universe of Medicare beneficiaries.

#### 5.3.1 Utilization

Table 6 presents estimates of the effect on utilization. The first column displays the dependent variable, and each row shows results from a separate regression. The coefficient on leave-out costs can be interpreted as the effect on the dependent variable of a \$100 increase in leave-out costs. Given the one-for-one relationship between the instrument and premiums (Table 4), we can interpret this coefficient as the effect of a \$100 increase in Medigap premiums.<sup>50</sup> We also translate the estimates into an implied effect of Medigap by dividing these coefficients by the coefficient on leave-out costs from the preferred HSA-level demand specification of -0.048. When we evaluate the effect of taxing Medigap in Section 7, we consider robustness to the alternative demand estimates in Table 5. The baseline specifications include controls for demographics (age, sex, and race), geographic price indexes (GAF and OWI), and chronic conditions.<sup>51</sup> Standard errors are clustered at the HSA level.

Table 6 shows that most categories of utilization are decreasing in leave-out costs—implying that Medigap coverage increases Medicare utilization. The first row shows that a \$100 increase in leave-out costs reduces Part B events (line-item claims) by 0.42, and this estimate is statistically sig-

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<sup>50</sup> Another way to interpret the coefficient is that the coefficient gives us the cost savings to the Medicare program from Medigap dis-enrollment from  $\$ \frac{100}{\rho}$  tax where the pass-through rate of the tax is  $\rho$ . This interpretation makes sense as long as the demand estimate implies that a tax of size  $\$ \frac{100}{\rho}$  is not so large as to make the Medigap market disappear. This condition is satisfied for our demand estimate.

<sup>51</sup> Appendix D displays the full list of chronic health condition controls. Appendix Table E1 shows that the exclusion of chronic conditions controls has a statistically indistinguishable effect on the utilization estimates.

nificant with a p-value of 0.02. Dividing by the demand coefficient on leave-out costs implies that Medigap increases Part B events by 8.7 or 33.7% of the average number of events. The second and third rows examine subcategories of Part B events that are often considered more discretionary and may be more elastic to variation in cost-sharing. We find that a \$100 increase in leave-out costs reduces imaging events (e.g., X-rays, CT scans, MRIs) by 0.08, implying a Medigap effect of 1.7 or 42.4% of the average. We find that a \$100 increase reduces testing events (e.g., glucose tests, bacterial cultures, EKG monitoring) by 0.41, implying a Medigap effect of 8.5 or 74.6% of the average.<sup>52</sup>

We also use the 20% sample of claims data from the CMS Carrier file to examine effects on other measures of Part B utilization. For each line-item Part B claim, these data provide us the relative value units (RVUs) of the care provided. A RVU is a measure constructed by CMS that is intended to reflect relative input intensity, and CMS scales this measure to determine Medicare payments. The estimates indicate that a \$100 increase in leave-out costs reduces RVUs by 1.3, implying a Medigap effect of 26.9 or 38.0% of the average. The effect is statistically significant with a p-value less than 0.01.

The next two rows show the effects of the instrument on Part A hospital utilization. We find evidence that Part A hospital stays and Part A hospital days decrease with the instrument (increase with Medigap enrollment). The estimates suggest that a \$100 increase in the instrument reduces the number of Part A hospital stays by 0.004 with an implied Medigap effect of 23.9%. A \$100 increase in leave-out costs reduces the number of Part A hospital days by 0.06, for an implied Medigap effect of 1.3 or 61.6%. The associated p-values of these estimates are 0.065 and 0.001, respectively.

There is suggestive evidence that the reduction in Part A hospital utilization may be due in part to substitution from Part A hospital care to Skilled Nursing Facility (SNF) care. SNFs provide care to recently discharged patients who need skilled medical and rehabilitative care. Although receiving Part A care requires significant cost-sharing, Medicare provides complete coverage for SNF care with no deductible for the first 20 days per benefit period.<sup>53</sup> Thus, patients without

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<sup>52</sup>As indicated in the Beneficiary Summary File data documentation, imaging events are defined as claims with a line BETOS code that starts with the letter "I." Testing events are claims with a line BETOS code that starts with the letter "T."

<sup>53</sup>To qualify for SNF coverage during a benefit period, beneficiaries must have a qualifying hospital stay of 3 days or longer and enter the SNF within 30 days of hospital discharge for services related to the hospital stay.



Medigap have an incentive to obtain this care at an SNF. We find suggestive evidence that an increase in leave-out costs *raises* SNF Days and SNF Stays. While the estimates are not statistically distinguishable from zero, the point estimate for SNF Days suggests that substitution to SNF may explain 19.3% ( $=0.012/0.062$ ) of the decline in Part A Days caused by Medigap.

### 5.3.2 Medicare Payments

We begin by presenting graphical evidence of the effect of Medigap on Medicare payments. Figure 5 shows the effect of a \$1,000 increase in leave-out costs on the distribution of Part A, Part B, and total Medicare payments. Solid lines show the CDF of payments in each category.<sup>54</sup> Dashed lines depict the effect of a \$1,000 increase in leave-out costs.<sup>55</sup> The lines are calculated using the coefficient on leave-out costs from regressions of the form  $\Pr(\text{Payments}_{ijk} < X) = \gamma_c \text{Leave-out costs}_{jk} + \gamma_k + X'_{ijk} \gamma_X + \mu_{ijk}$  where  $X = 500, 1,000, \dots, 32,000$ . Dotted lines show the 95% confidence intervals of these estimates, calculated using standard errors clustered at the HSA level.

Medicare payments, like utilization, are decreasing in leave-out costs—implying that Medigap coverage increases Medicare payments. The effects are largest for the lowest levels of spending and decrease monotonically across the spending distribution. These results are consistent with the view that the cost-sharing elasticity of medical spending is decreasing over the spending distribution.<sup>56,57</sup>

Table 7 presents estimates of the effect on Medicare payments. The table layout is identical to Table 6 on utilization. The first column displays the dependent variable, and each row shows results from a separate regression. We show the coefficient on leave-out costs (measured in hundreds of dollars) and the implied effect of Medigap. These baseline specifications include controls for demographics (age, sex, and race), geographic price indexes (GAF and OWI), and chronic conditions.<sup>58</sup> Standard errors are clustered at the HSA level.

<sup>54</sup>The distribution is censored at \$32,000 per year.

<sup>55</sup>Our discussion of Figure 5 focuses on the overall patterns in the figure rather than the level shift in the cdf, and in this sense, the exact shift in leave-out costs used to construct the figure is not important. However, to put the estimates into context, a \$1,000 increase in leave-out costs is just large enough to eliminate the Medigap market based on our preferred Medigap demand estimates.

<sup>56</sup>These estimates contrast with Kowalski (2010), who estimates constant elasticities across the distribution of spending in the sample of non-elderly individuals that she studies.

<sup>57</sup>We cannot rule out the possibility that the larger effects for lower levels of spending are caused by a larger response of Medigap demand for low spending individuals.

<sup>58</sup>Appendix D displays the full list of chronic health condition controls. Appendix Table E2 shows that the exclusion of chronic health condition controls has a statistically indistinguishable effect on the payment estimates.

Table 7 shows that a \$100 increase in leave-out costs reduces Part A payments by \$47.59 and Part B spending by \$21.80. These estimates imply that Medigap raises Part A spending by \$992 or 32.8% and Part B spending by \$454 or 17.1%. Similar to the utilization results, we find that SNF Payments are decreasing in leave-out costs, although the estimate lacks statistical precision. The point estimate for SNF payments suggests that a \$100 increase in leave-out costs raises SNF spending by \$3.44. The implied Medigap effect is -\$72 or a reduction of 17.9%.

The top row of Table 7 shows the effect on total Medicare payments. A \$100 increase in leave-out costs reduces total Medicare payments by \$67.02, and this estimate is statistically significant with a p-value of 0.043. This estimate implies that Medigap increases Medicare payments by \$1,396 on a mean of \$6,290 or 22.2%.<sup>59</sup>

Our preferred estimate—that Medigap increases Medicare payments by 22.2%—is comparable to non-quasi-experimental estimates of the effect of Medigap and implies a price elasticity similar to standard estimates in the literature. For example, the CBO (2008) summarizes the non-quasi-experimental literature as showing that Medigap increases Medicare utilization by about 25% relative to the counterfactual of no supplemental insurance.<sup>60</sup> As emphasized by Aron-Dine, Einav and Finkelstein (2013), summarizing the effect of health insurance with a single elasticity parameter is difficult because non-linear health insurance contracts do not exhibit a well-defined out-of-pocket “price” for medical care. This is particularly true for Medicare since cost-sharing is nonlinear in the level of utilization (e.g., Part A deductible, copays) and cost-sharing varies across categories of medical care (e.g., Part A, Part B, SNF). However, if we put those concerns aside and assume that Medigap reduces cost-sharing from 20% to 0%, then our preferred estimate that Medigap increases utilization by 22.2% implies an arc-elasticity of -0.11, which is in the same range as the classic RAND estimate of -0.2 (Keeler and Rolph, 1988).<sup>61</sup>

<sup>59</sup>Given the sizable effects on utilization and Medicare payments, one might be interested in testing whether Medigap reduces mortality. Appendix F shows results consistent with Medigap having no effect on mortality. Specifically, Appendix F demonstrates that the age distribution (conditional on reaching age 65) is unrelated to the identifying variation.

<sup>60</sup>The non-quasi-experimental literature tends to use regression analysis with several covariates for demographics, individual SES, and health conditions to control for observable differences between those that do and do not have Medigap. According to a recent GAO report (GAO, 2013), the raw mean cost differences suggest that those with Medigap spend nearly 100% more than those with no supplemental coverage. Comparing these estimates to our estimates, one can see some prior estimates that control for rich set observables are similar to our quasi-experimental estimate and comparing raw means provides a very inaccurate estimate of the effect of Medigap.

<sup>61</sup>Let  $q_1$  and  $p_1$  be the quantity and price without supplemental insurance and let  $q_2$  and  $p_2$  be the price and quantity with Medigap. The arc elasticity is given by  $\epsilon_{arc} = \frac{q_2 - q_1}{(q_2 + q_1)/2} / \frac{p_2 - p_1}{(p_2 + p_1)/2}$ .

Section 6 presents robustness checks of the effects on utilization and spending. Specifically, we present more evidence in support of our identifying assumption, and we demonstrate that the baseline results are robust to alternative specifications.

## 6 Robustness

The basic threat to our identification is that there may be some omitted factor that is correlated with both our instrument and Medicare utilization. In Section 4, we showed that ZIP code-level demographic characteristics such as income, labor force participation, and education are not correlated with our instrument. Below, we present three additional pieces of evidence in support of our identification strategy. First, we demonstrate that our baseline results are robust to the inclusion of additional control variables. Second, we conduct two sets of falsification tests to demonstrate that omitted factors that change sharply at state boundaries are unlikely to be driving our results. Third, we estimate several specifications that allow us to assess whether our results are driven by unrelated spatial trends in medical spending.

### 6.1 Alternative Specifications

Table 8 shows the results of alternative specifications for the cost regressions. The first row displays the baseline Medicare spending results for reference. The second row displays the results when ZIP code-level Census demographic variables are added to the baseline specification. The third row displays the results when HSA-year fixed effects are added to the baseline specification.<sup>62</sup> The point estimates are stable across all the specifications, with an implied Medigap effect ranging from \$1,396 to \$1,089. These results demonstrate that the baseline estimates are robust to the inclusion of additional controls.

### 6.2 Falsification Tests

It would be a problem for the identification strategy if there are omitted factors related to Medicare spending that are also correlated with the instrument. For example, if the underlying health of the

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<sup>62</sup>In the baseline specification, HSA and year fixed effects enter separately.

population changed sharply at state boundaries in a way that was correlated with our instrument, our results may simply reflect this health differential and not the effect of Medigap.

We present two pieces of evidence below that help to alleviate this concern. First, we show that procedures that are very urgent (and thus should not be affected by our instrument) are indeed not correlated with the instrument. Second, we demonstrate that health outcomes do not covary with our instrument for individuals younger than 65 who are not eligible for Medigap. Together, these tests reveal that factors affecting utilization in general (for example, the underlying health of the population) are not driving the results.

### 6.2.1 Unaffected Procedures: Urgent Procedures

We investigate the relationship between our instrument and urgent procedures using definitions of urgent procedures from the literature. First, we investigate urgent Part B claimed RVUs using the characterization from [Clemens and Gottlieb \(2013\)](#) based on a BETOS code classification. Second, we investigate urgent hospital admissions based on the methodology of [Card, Dobkin and Maestas \(2009\)](#) that characterizes urgent hospitalizations as those with similar daily frequencies on weekdays and weekends.<sup>63</sup> We consider two variants of this definition of urgent hospitalizations. First, we investigate the ten most common non-deferrable conditions identified by [Card, Dobkin and Maestas \(2009\)](#) in their data. Second, we use the same methodology as [Card, Dobkin and Maestas \(2009\)](#) with our data (the CMS MedPAR data) to characterize the set of urgent hospitalizations. Appendix G describes all three characterizations of urgent procedures in detail.

Table 8 presents the results of these regressions, which repeat the baseline specification replacing the dependent variable with the number of urgent procedures based on the characterizations described above.<sup>64</sup> Across the different classifications, there is no evidence of an effect of leave-out costs on urgent procedures. The point estimates are statistically insignificant (p-values ranging from 0.20 to 0.51), and the point estimates vary greatly in terms of magnitude and sign. These results suggest that it is unlikely that discontinuities in other health-related factors are driving the

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<sup>63</sup>This analysis is done using the CMS MedPAR files that contain hospital claims data for 100% of FFS Medicare beneficiaries.

<sup>64</sup>As in the baseline specification, these regressions are run at the individual-year level, so the measure of urgent procedures is also at the individual-year level. The [Clemens and Gottlieb \(2013\)](#) measure is based on the 20% of individuals for which Part B claims data are available (in the CMS Carrier file). The two [Card, Dobkin and Maestas \(2009\)](#) measures are created using the CMS MedPAR data available for 100% of beneficiaries.

main results.

### 6.2.2 Unaffected Individuals: Non-Elderly Individuals

Next, we show that the instrument is unrelated to outcomes for individuals who should be unaffected by the instrument. To do this, we use data from the NHIS on non-elderly adults (adults aged 18-64). We examine effects on utilization measures including hospital stays and hospital days. In addition, we examine the effect on self-report health, measured with a Likert Scale that runs from 1 to 5, with 1 indicating "Excellent" and 5 indicating "Poor."

Table 8 presents the results of these regressions, which as before repeat the baseline specification replacing the dependent variable with these measures of utilization and health status among the non-elderly. Across the three measures, the coefficient on leave-out costs is statistically indistinguishable from zero. Although the limited sample size of the NHIS prevents us from ruling out effects, these falsification tests show no evidence of any covariance between health outcomes and our instrument for individuals younger than 65 who are ineligible for Medigap.

## 6.3 Robustness to Spatial Trends in Utilization

Many determinants of health care utilization vary continuously over geography, including provider choice, environmental factors, and behavioral factors. If these determinants of health care utilization are correlated with the instrument, our identification assumption will not hold. We address this concern in two ways. First, we re-estimate the baseline specification, restricting the sample of individuals within cross-border HSAs to be those within a very short distance of the state boundary. The idea behind this sample restriction is that if there are spatial trends in health care utilization (driven by characteristics such as provider choice and demographics), then those individuals who live closest to one another are the best controls for one another. Table 8 reports the results. The point estimates remain statistically significant and similar in magnitude when we concentrate on the samples with 20 kilometers and 10 kilometers of state boundaries.<sup>65,66</sup> This is

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<sup>65</sup>The samples used for these specifications drop individuals in cross-border HSAs that reside more than 20 km and 10 km from the border, respectively. These specifications still include all individuals who do not reside in cross-border HSAs, as these individuals continue to assist in identifying the coefficients of the control variables.

<sup>66</sup>Within cross-border HSAs, the mean distance from a ZIP code centroid (our most disaggregated measure of location) to the state boundary is 25 km and the median distance is 16 km.

reassuring as this restricted sample contains individuals who are most similar to one another in terms of continuously trending unobservables.

Second, we verify that our estimates are robust to spatially trending omitted variables by estimating a specification with carefully defined placebo borders. Specifically, we partition each HSA-state segment in cross-border HSAs into two areas: the *border area* within 20 km of the state boundary and the *near border area* consisting of the remainder of the HSA-state. The placebo border is then the division between these two areas, meaning that placebo border is entirely internal to the state in question. We then assign the *border area* a counterfactual instrument equal to the instrument of the neighboring state, while the *near border area* has the true value of the instrument as in our baseline estimation. With this newly defined instrument determined by the placebo border, we then run the same regressions as in the baseline specification replacing the HSA fixed effects with HSA-state fixed effects. The results are reported in Table 8. If the baseline results are not picking up the causal effect of Medigap but instead reflecting unrelated spatial trends in medical spending, then one would expect the coefficient from this specification to be the same as in our baseline specification. In contrast to the significant results in our baseline estimation, we see that the coefficient in this specification is statistically indistinguishable from zero (with a p-value of 0.54). This test reveals that our estimated effect of Medigap is not simply reflecting unrelated, continuous spatial trends in medical utilization.

## 7 Policy Counterfactuals

A natural policy to address the externality from Medigap is a tax on Medigap premiums. The idea of taxing Medigap premiums is not new. In fact, many policymakers have suggested taxing Medigap beneficiaries as a means to reduce Medicare costs. For example, the Obama administration's 2013 budget proposal calls for a 15% tax on Medigap policies. In *Budget Options Volume I: Health Care*, the Congressional Budget Office (CBO) considered a 5% excise tax on Medigap premiums as a mechanism for reducing the federal deficit. Below, we investigate the effect of corrective taxation on Medicare's budget and welfare.

## 7.1 Medicare's Budget

A tax presents two sources of savings for the Medicare program. First, a tax discourages some individuals from enrolling in Medigap, which reduces their Medicare spending by removing the externality estimated above. Second, tax revenues are raised from those remaining Medigap purchasers. Thus, to calculate the impact of a tax requires both our estimate of the price sensitivity of demand for Medigap (to estimate Medigap enrollment under the tax) and our estimate of the moral hazard externality associated with Medigap (to estimate the cost savings from individuals who drop coverage).<sup>67</sup>

We use the results presented in Section 5 to produce estimates of the effect of a tax on Medigap premiums in the following manner. First, the counterfactual Medigap market share is calculated using the estimated demand elasticity, assuming the tax is fully passed through to consumers.<sup>68</sup> The demand curve used for these calculations has a slope equal to  $\partial q_{ijk} / \partial \text{Leave-Out costs}_{jk} = -0.048$  (as the coefficient in the premium regressions was approximately one) and an intercept pinned down by the national Medigap market share of 48% and the national average premium of \$1,779. The tax revenue raised is then calculated by multiplying the tax by the resulting Medigap market share. Medicare cost savings are calculated by applying the Medigap externality calculated above to all those who drop their Medigap coverage due to the tax (the change in the Medigap market share). The total budgetary impact is simply the sum of the tax revenue raised and Medicare cost savings from Medigap dis-enrollment. The parameters used in this calculation are estimated using local variation in premiums, and the projected effects of larger taxes should be viewed with appropriate caution.

Table 9 shows the results of this exercise. Each row displays the results for a different tax rate; the columns display the implied tax revenue raised, the Medicare cost savings obtained through Medigap dis-enrollment, and the total budgetary impact on the Medicare program. The per-capita numbers presented in this table refer to the non-dual eligible, FFS Medicare population (the esti-

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<sup>67</sup>The one exception occurs in the calculation of budgetary effects of the optimal Pigouvian tax, in which the tax is set equal to the moral hazard externality. In this special case, it is necessary only to estimate the moral hazard externality. The elasticity of demand is irrelevant since an individual has the same budgetary impact whether he maintains Medigap coverage (and pays the excise tax) or drops coverage (and generates cost savings of the same magnitude).

<sup>68</sup>The calculations in Table 9 assume that the tax is fully passed through to consumers. If the pass-through rate is  $\rho$ , it would take a tax of size  $\frac{x}{\rho}\%$  to achieve the Medicare budgetary impact we calculate for an  $x\%$  tax.



mation sample). A 15% tax on Medigap premiums would raise \$94 per beneficiary in tax revenue and reduce Medicare costs per beneficiary by \$179 for a total savings of \$273 per beneficiary or 4.3% of per-capita costs. Appendix Table [H1](#) shows that this estimate varies from 3.9% to 4.8% using alternative demand estimates. Combining the standard errors associated with our demand and cost estimates, we calculate the standard error of our baseline estimate of 4.3% total savings is 1.7 percentage points.<sup>69</sup>

We can translate this calculated per-capita savings into the aggregate savings for the current Medicare program. In 2012 dollars, the per-capita savings from a 15% tax for non-Medicaid eligible, FFS Medicare enrollees is \$321. By law, Medicare Advantage payments are set to be a function of the local FFS Medicare spending. Thus, if we assume that Medicare Advantage capitation payments are reduced by the same amount as the FFS Medicare spending as a result of the tax, then the per-capita savings for Medicare Advantage beneficiaries is also \$321. There are roughly 27.4 million FFS, non-Medicaid eligible Medicare beneficiaries and 12.7 million Medicare Advantage enrollees ([KFF, 2012](#)). Summing across these beneficiaries, the total savings for the Medicare program from a 15% tax is estimated to be \$12.9 billion.

Table [9](#) shows that a Pigouvian tax that fully accounts for the estimated externality would completely eliminate the Medigap market, saving the Medicare program \$670 per capita or 10.7% of total Medicare costs. When we adjust for inflation and assume that the savings are internalized by the Medicare Advantage program, this translates into total savings for the Medicare program of roughly \$31.6 billion in 2012 dollars.

## 7.2 Welfare

We can also use our parameter estimates to assess the welfare effects of a tax on Medigap. For this welfare discussion, we consider the constrained world where we take Medigap and Medicare coverage as given, and we discuss distortions relative to a world with only Medicare coverage. Because a more general evaluation of welfare and the optimal cost-sharing structure of Medicare

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<sup>69</sup>Appendix Table [H1](#) displays the projected total savings and standard errors associated with a 15% tax using the various demand estimates. To calculate the standard error on the total savings, we first separately calculate the standard error on the tax revenue raised (from the corresponding demand estimate) and the standard error from the Medicare cost savings from Medigap dis-enrollment (from the reduced form). We then obtain the standard error on the total savings by aggregating these standard errors using the Delta Method assuming no covariance between the demand and cost estimates.

would require many more modeling assumptions, we limit our discussion to the setting we can evaluate with our estimates.

Figure 6 displays supply and demand in the Medigap market under the assumption of perfect competition and constant marginal costs. Under these assumptions, we have the standard “price equals marginal costs,” and the private marginal cost curve can be approximated by a horizontal line at the average Medigap premium of \$1,779. The social marginal cost curve is the sum of private costs and the externality and is depicted in the figure by the horizontal line at \$3,175 ( $=\$1,396 + \$1,779$ ). The equilibrium with no tax is represented by point A, the intersection of the private marginal cost curve and the demand curve. The social optimum results in the elimination of the Medigap market. The deadweight loss from the fiscal externality of Medigap is given by the trapezoid AIHG. In this figure, the net efficiency gain from a Pigouvian tax is 64% of the total impact on Medicare’s budget; the remaining 36% is a transfer of surplus from individuals who otherwise would have purchased Medigap to taxpayers.

Figure 6 also illustrates the private marginal cost curve in the case of a smaller tax  $\tau$  that does not cause the Medigap market to disappear. The effect of a tax  $\tau$  on Medicare’s budget is depicted by the sum of two rectangles: CEFB (the tax revenue raised) and ACJG (the cost savings from Medigap dis-enrollment). The net efficiency gain is represented by the deadweight loss trapezoid ABJG. Comparing this welfare gain to the overall impact on Medicare’s budget shows that only a fraction of the impact on Medicare’s budget is a net welfare gain. The remainder of Medicare’s total savings comes from transfers from Medigap purchasers and individuals deterred from purchasing Medigap because of the tax. These transfers are represented in the figure by the rectangle CEFB (tax revenue raised from Medigap purchasers under the tax) and the triangle ACB (consumer surplus forgone by individuals discouraged from purchasing Medigap because of the tax).

There are at least two important caveats to note about the counterfactual calculations. First, our analysis focuses on evaluating the effect of a tax on Medigap premiums taking the form of Medigap and Medicare as given. Although the first-best policy to address the Medigap externality may involve broader changes to Medigap or Medicare coverage, we limit our analysis to policy counterfactuals for which we can obtain credible estimates.<sup>70</sup> Still, it should be noted that taxing

<sup>70</sup>See Pauly (2000) for a theoretical discussion of the efficiency trade-offs involved in the simultaneous public and

Medigap premiums is one of the most discussed policies related to Medigap, and our identifying premium variation gives us a unique opportunity to evaluate the effect of a tax on Medigap premiums.

Second, the welfare discussion above abstracts from market power. To the extent that there is market power in the Medigap market, the distortions from market power already act as an implicit tax, raising the price relative to the social marginal cost.<sup>71</sup> In general, when calculating an optimal tax, one needs to trade off the forgone consumer surplus from higher premiums with the reduction in moral hazard. It turns out that our estimate of the Medigap externality is large enough that an optimal tax would substantially reduce the size of the Medigap market regardless of the form of competition.<sup>72</sup> Of course, the exact welfare effect of such a tax would need to be measured relative to the correct equilibrium and cost curves (which, in the case of market power, would differ from those depicted in Figure 6). It is also important to note that regardless of whether there is market power, our estimates of the budgetary impact of Medigap are still equally applicable.<sup>73</sup>

## 8 Conclusion

Medicare includes cost-sharing to reduce unnecessary utilization. Since beneficiaries can purchase supplemental insurance from Medigap, they are able to reduce their exposure to this cost-sharing, potentially increasing utilization and imposing a negative externality on the Medicare system.

The goal of this paper is (i) to provide a causal estimate of the externality Medigap imposes on the Medicare system and (ii) to estimate how a corrective tax on Medigap would impact Medicare costs and welfare. Using Medigap premium discontinuities that occur at state boundaries and an estimated demand elasticity of -1.8, we find that Medigap increases overall Medicare costs by \$1,396 per year on a base of \$6,290 or by 22.2%. These results are robust to alternative specifications of private provision of insurance within the Medicare context.

<sup>71</sup>Starc (2010) estimates that markups are substantial in this market (on the order of 30%). From the point of view of Medicare's budget, such markups can help to reduce costs much in the way a tax would operate.

<sup>72</sup>Let us assume firms face constant marginal costs equal to the observed average uncovered Medicare spending \$911 in our data. (Note that this number is likely conservatively low relative to insurer average costs if there is either adverse selection in the Medigap market or administrative costs associated with Medigap policies.) In this case, regardless of the structure of competition, a Pigouvian tax would bring the after-tax premium to a minimum of \$2,307 ( $=\$911 + \$1,396$ ), as insurers will avoid making losses; the implied Medigap market share at a premium of \$2,307 is approximately 23%. In other words, our estimated Medigap externality is high enough that an optimal Pigouvian tax would cause the Medigap market to shrink by at least 50% of its current size regardless of the form of competition.

<sup>73</sup>The only potential caveat is that one may want to adjust the pass-through rate to reflect market power, though interpreting our findings for a different pass-through rate is simple (as discussed in Table 9).

tions and to falsification tests using individuals and procedures that should be unaffected by the variation in premiums.

Our estimates indicate that a 15% tax on Medigap premiums, with full pass-through, would decrease Medigap coverage by 13 percentage points on a base of 48% and reduce net government costs by 4.3% per Medicare beneficiary or \$12.9 billion in 2012 dollars. About 35% of these savings would come from tax revenue while the remainder would come from lower Medigap enrollment. A Pigouvian tax requires us to extrapolate outside the premium variation in the data. To a first approximation, such a tax would generate combined savings of 10.7% per beneficiary or \$31.6 billion in 2012 dollars.

In closing, we want to emphasize that taxing Medigap is not the only way to address the externality from Medigap. Although taxing Medigap has received substantial attention, such a tax is a fairly blunt instrument for increasing Medicare's efficiency, and other policies may lead to even larger efficiency gains. For example, the idea of banning Medigap and re-designing Medicare coverage to include an annual out-of-pocket maximum may offer beneficiaries almost as much risk protection as Medigap while leading to less excess Medicare utilization.<sup>74</sup> In general, the optimal policy to address this externality—and the optimal design of health insurance coverage more broadly—should trade off the risk-protection benefits against the moral hazard costs.

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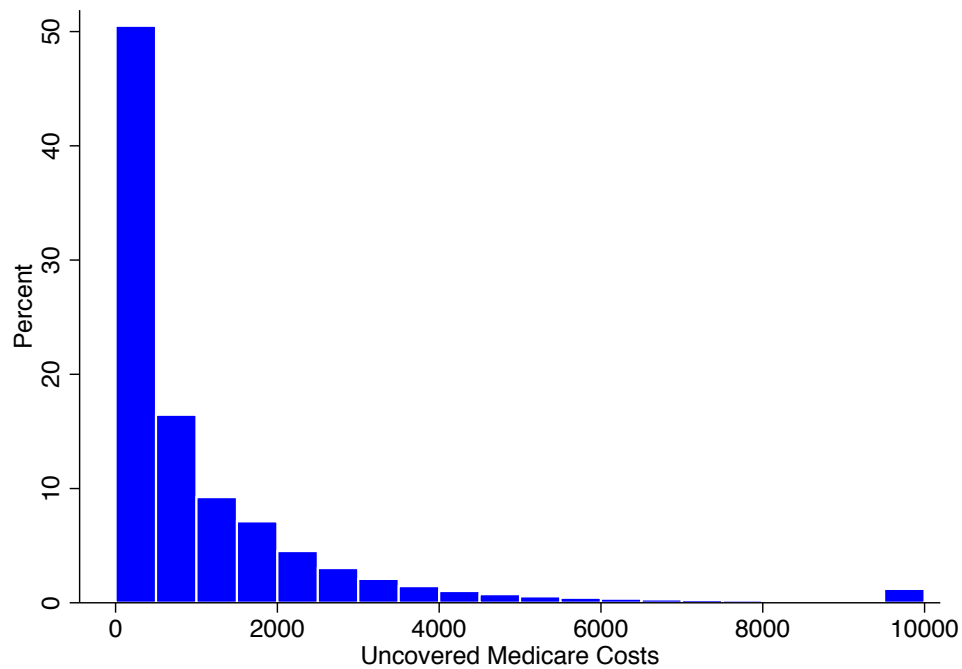
<sup>74</sup> Although our estimates cannot speak to welfare under this alternative directly, our analysis of the effect of Medigap on the distribution of Medicare payments suggests that there may be very little moral hazard at high expenditure levels.

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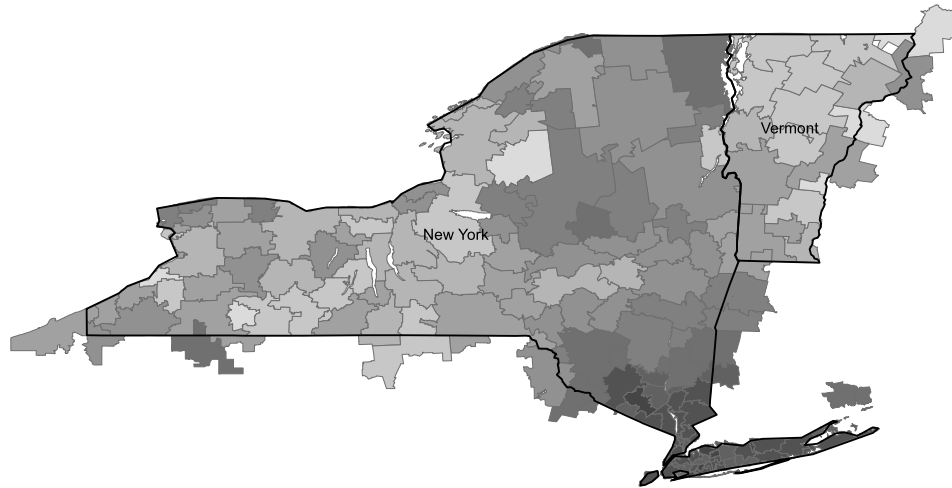
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**Figure 1: Uncovered Medicare Costs**



**Notes:** This figure shows a histogram of uncovered Medicare spending, defined as Medicare-eligible spending that is the responsibility of the beneficiary, and thus is paid out-of-pocket by the beneficiary, paid by supplemental insurance, or absorbed by the medical provider as bad debt. The data are from the 2005 CMS Beneficiary Summary File and cover the universe of FFS Medicare, non-Medicaid beneficiaries. Costs are top-coded at \$10,000. Approximately 3.8% of beneficiaries have uncovered Medicare spending exceeding \$5,000, and approximately 1% of beneficiaries have out-of-pocket annual uncovered Medicare costs greater than \$10,000. All dollar values are stated in terms of 2005 dollars (inflation-adjusted using the CPI-U).

**Figure 2:** Uncovered Medicare Expenditures and Medigap Premiums for NY and VT



(a) Uncovered Medical Expenditures by HSA, NY and VT

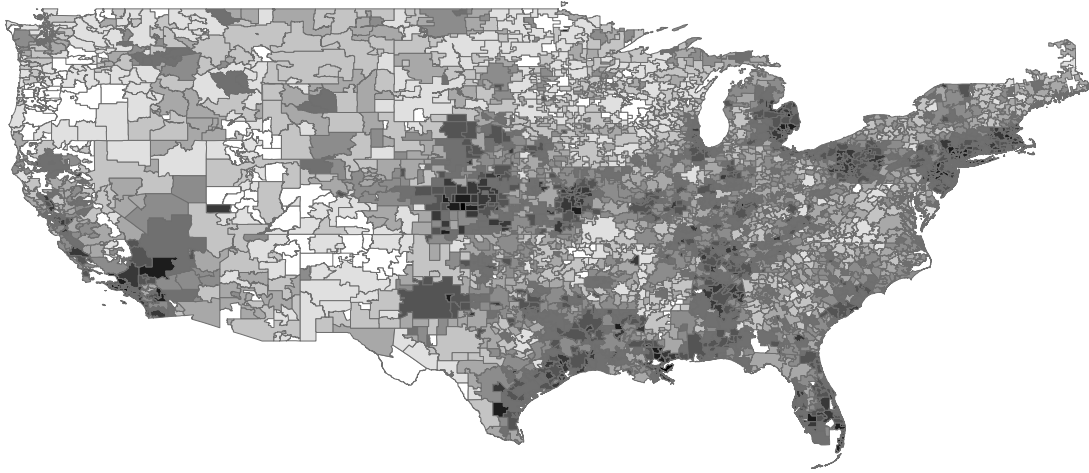


(b) Medigap Premiums for NY and VT

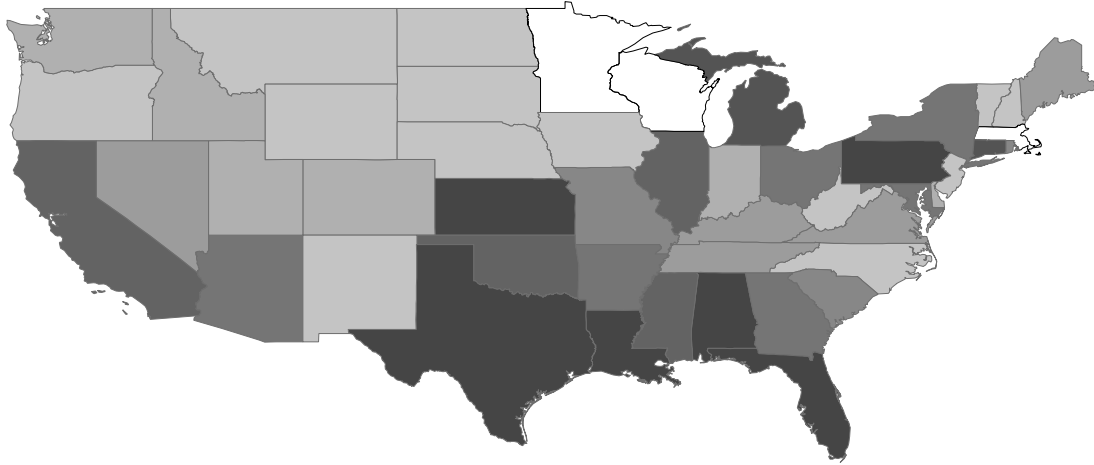
**Notes:** The above two panels focus on the states of New York and Vermont. The map in Panel (a) displays by HSA the average annual “Uncovered Medicare Expenditures,” where we define this as the Medicare-eligible medical spending that is the beneficiary’s responsibility (expenditures either paid out-of-pocket by the beneficiary or by the supplemental insurer). The CMS Beneficiary Summary File for 2000 was used to make this map. Uncovered Medicare Expenditures by HSA range from \$766 per capita in the HSA centered on Lowville, NY (a village in upstate NY) to \$1,585 in the HSA centered on Far Rockway, NY (a neighborhood of NYC). Among HSAs within these two states, the 5th percentile of Uncovered Medicare Expenditures is \$843, the 10th percentile is \$847, the median is \$956, the 90th percentile is \$1,293, and the 99th percentile is \$1,404. Panel (b) displays the average state-level Medigap premiums among all Medigap insurers. Premium data from Weiss ratings for 2000 were used to make this map. The average annual Medigap premium is \$1,504 in New York and \$1,058 in Vermont. All dollar values are stated in terms of 2005 dollars (inflation-adjusted using the CPI-U).



**Figure 3: Uncovered Medicare Expenditures and Medigap Premiums**



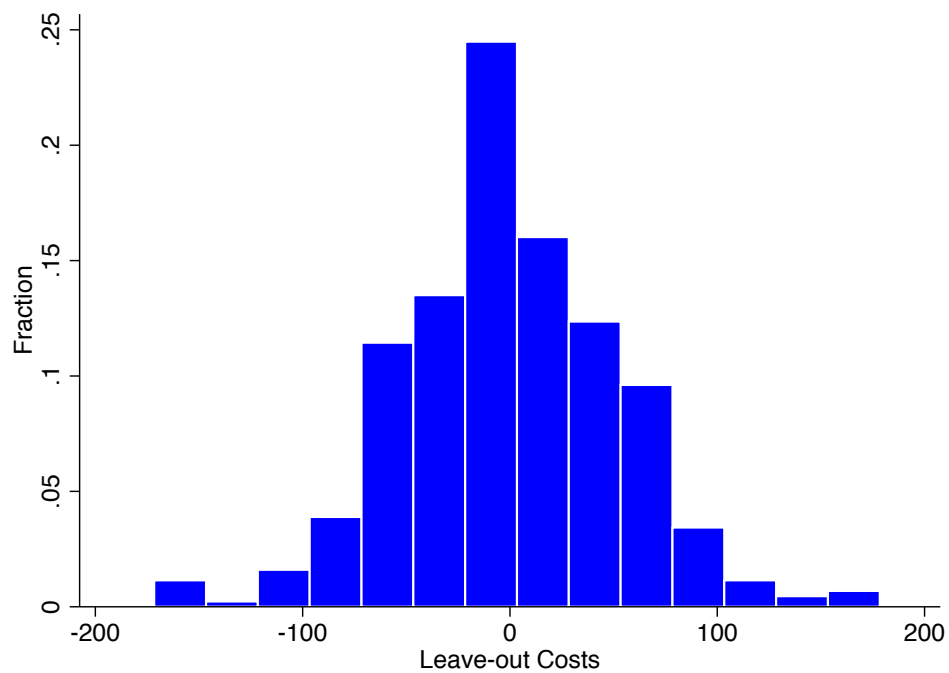
(a) Uncovered Medicare Expenditures by HSA



(b) Medigap Premiums by State

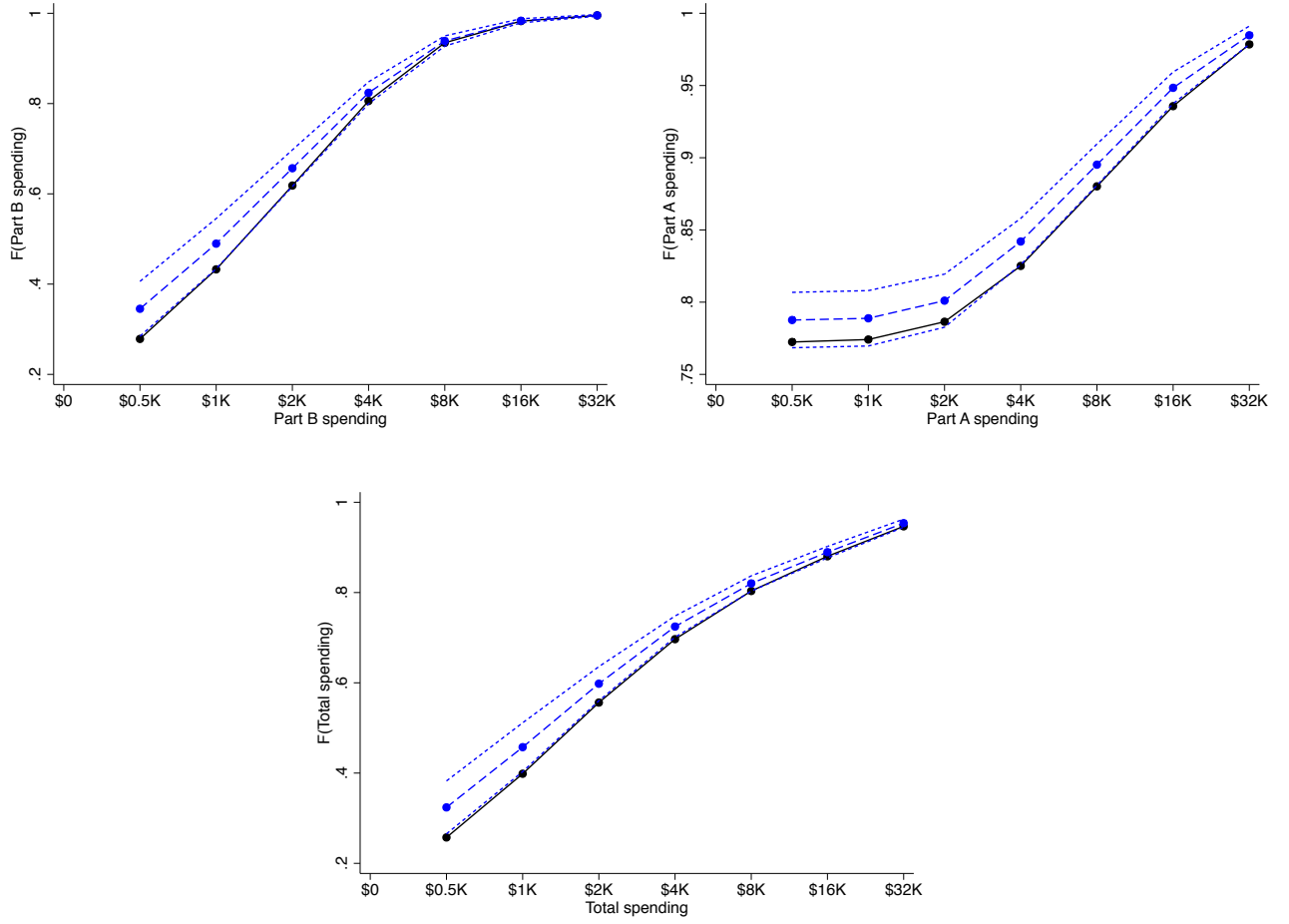
**Notes:** The map in Panel (a) displays by HSA the average annual “Uncovered Medicare Expenditures,” where we define this as the Medicare-eligible medical spending that is the beneficiary’s responsibility (expenditures either paid out-of-pocket by the beneficiary or by the supplemental insurer). The CMS Beneficiary Summary File for 2000 was used to make this map. The 5th percentile of HSA-level Uncovered Medicare Expenditures is \$705, the 10th percentile is \$801, the median is \$944, the 90th percentile is \$1,131, and the 99th percentile is \$1,368. Panel (b) displays the average, annual average state-level Medigap premiums for the two largest insurers, United Healthcare and Mutual of Omaha. Premium data from Weiss ratings for 2000 was used to make this map. Premium data do not exist for Wisconsin, Massachusetts, and Minnesota. These three states do not have standardized Medigap policies. The average Medigap premium is \$1,456, and the median is \$1,448. The 90th percentile is \$1,772 and the 10th percentile is \$1,232. All dollar values are stated in terms of 2005 dollars (inflation-adjusted using the CPI-U).

**Figure 4: Leave-Out Costs**



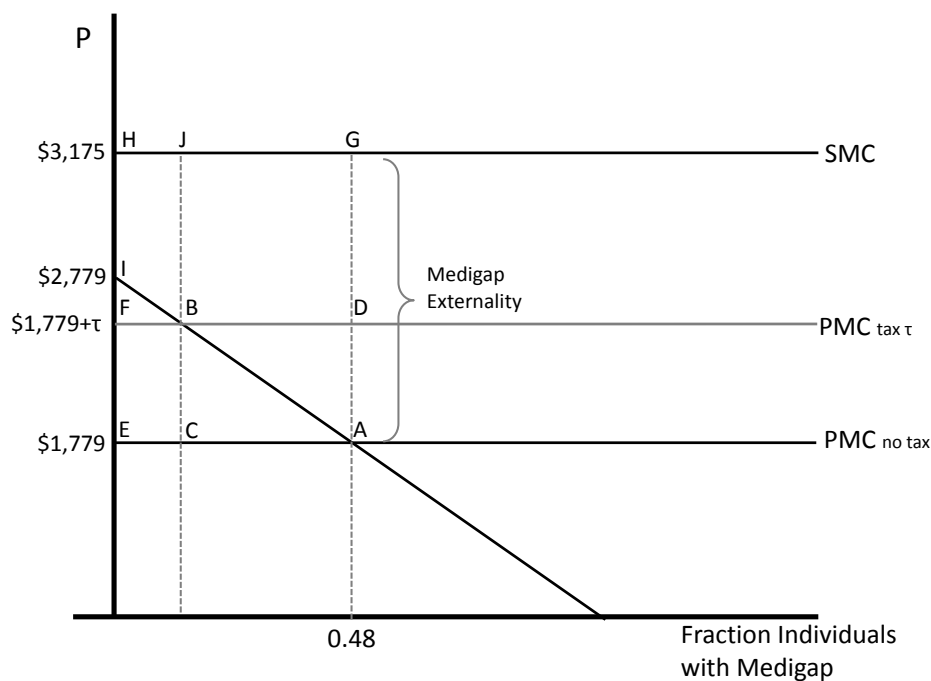
**Notes:** This figure shows a histogram of the leave-out costs instrument net of the mean of the instrument within border-spanning HSAs. The leave-out cost instrument is defined using data from the CMS Beneficiary Summary File for 2000 (the same year for which premium data are available). All dollar values are stated in terms of 2005 dollars (inflation-adjusted using the CPI-U).

**Figure 5: Effect on CDF of Payments**



**Notes:** This figure shows the impact of a \$1,000 increase in leave-out costs on the CDF of Part B payments, Part A payments, and total Medicare payments. The solid line show the empirical CDF of payments. The dashed lines show the estimated CDF under an \$1,000 increase in leave-out costs. These CDFs are constructed using the coefficient on leave-out costs from regressions of the form  $\Pr(\text{Payments} < X) = \gamma_c \text{Leave-out costs}_{jk} + \gamma_k + X'_{ijk} \gamma_X + \mu_{ijk}$  for  $X = 500, 1,000, \dots, 32,000$ . Dotted lines show the 95% confidence intervals of these estimates. All dollar values are stated in terms of 2005 dollars (inflation-adjusted using the CPI-U).

**Figure 6: Welfare Under Taxation**



**Notes:** The linear demand curve displayed above has a slope equal to  $\partial q_{ijk} / \partial \text{Leave-Out costs}_{jk} = -0.048$  (as the coefficient in the premium regressions was approximately one) and an intercept pinned down by the national market equilibrium ( $p=17.79$  hundreds of dollars and  $q=0.48$ ). Supposing that  $p=MC$  and marginal cost is constant, the figure displays the private marginal cost curve under no tax as a horizontal line at the observed average premium (\$1,779). The social marginal cost curve is the private marginal cost curve shifted upward by the Medigap externality. The deadweight loss of Medigap is the trapezoid AIHG. The figure also displays the private marginal cost curve under a smaller specific tax of magnitude  $\tau$ . All dollar values are stated in terms of 2005 dollars (inflation-adjusted using the CPI-U).

**Table 1: Medicare Cost-Sharing**

Part A: Hospital Expenditures			Part B: Physician Expenditures		SNF	
Deductible	Per-Day Copay		Deductible	Coinsurance	Deductible	Per-Day Copay
	Days 61-90	Days 91-150				Days 21-100
\$912	\$228	\$456	\$110	20%	\$0	\$114

**Notes:** The table shows FFS Medicare cost-sharing in 2005. (The cost-sharing is similar for the other years in the sample.) Part A cost-sharing is applied separately to each benefit period, where a benefit period begins upon a hospital or SNF admission and ends when the patient has been out of the hospital/SNF facility for 60 days. The patient faces the deductible and an increasing daily copay schedule for each benefit period where the number of benefit periods is unlimited. Medicare only pays for Part A hospitalizations in excess of 90 days through the draw down of 60 lifetime reserve days. Part B coverage above is on an annual basis with an annual deductible of \$110. Skilled Nursing Facility (SNF) coverage has no deductible for each benefit period, provides coverage with a daily copay of \$114 for days 21-100, and provides no coverage for SNF stays longer than 100 days.

**Table 2: Summary Statistics**

	Full Sample	Cross-Border HSA (11.0%)
<b>All Beneficiaries</b>		
Medicare Type (Denominator File, 1999-2005)		
Traditional Medicare (FFS)	74.5%	81.9%
Medicare Advantage	14.5%	6.6%
Medicaid (Dual-Eligible)	11.0%	11.5%
<b>Baseline Sample: FFS Medicare, Non-Medicaid Beneficiaries</b>		
Supplemental Insurance* (MCBS, 1992-2005)		
Medigap	47.9%	50.0%
Retiree Supplemental Insurance	46.3%	45.9%
None	15.8%	14.1%
Medigap Premiums	\$1,779	\$1,727
Costs (Beneficiary Summary File, 1999-2005)		
Part A Payments	\$3,020	\$2,776
Part B Payments	\$2,649	\$2,396
SNF Payments	\$399	\$337
Total Medicare Payments	\$6,290	\$5,760
Utilization (Beneficiary Summary File, 1999-2005)		
Part A Days	2.10	2.06
Part A Stays	0.34	0.34
Part B Events	25.82	24.01
Part B RVUs	70.79	64.87
SNF Days	1.37	1.25
SNF Stays	0.06	0.06
Demographics (Denominator File, 1999-2005)		
Sex		
Male	41.5%	41.5%
Race		
White	92.2%	92.5%
Black	5.6%	5.8%
Other	2.1%	1.7%
Age		
65-74	50.1%	51.7%
75-84	37.4%	36.7%
85+	12.5%	11.7%

\* The numbers for Medigap, RSI, and None add up to more than 100% because roughly 10% of individuals report both RSI and Medigap coverage.

**Notes:** The table above displays summary statistics from the Medicare administrative data obtained from CMS. The top panel summarizes the breakdown of insurance among all Medicare beneficiaries. The rest of the table presents summary statistics for the baseline sample (FFS Medicare, non-Medicaid beneficiaries). The cost and utilization data come from the CMS Beneficiary Summary File. The demographic data and insurance information come from the CMS Medicare Denominator File. All dollar values are stated in terms of 2005 dollars (inflation-adjusted using the CPI-U). The total number of beneficiary-year observations in the baseline sample in the CMS data is 133,467,621. The Medigap premium information comes from the Weiss Ratings premium data, and the premium measure displayed here is the average premium for plans offered by the top two insurers in 2000 to beneficiaries during the open-enrollment period.

**Table 3: Balance in Demographic Characteristics**

Dependent Variable	Coefficient on Leave-out Cost (in Hundreds)			
	Est.	Std. Err.	P-Value	Mean of Dep Var.
High School Degree, 65+	-0.013	(0.015)	0.39	0.65
Bachelors, 65+	-0.013	(0.012)	0.27	0.16
Veteran, Male 65+	-0.018	(0.008)	0.03	0.65
Veteran, Female 65+	-0.001	(0.001)	0.32	0.02
Labor Force Participation, Female 65-69	-0.008	(0.006)	0.26	0.20
Labor Force Participation, Male 65-69	-0.009	(0.011)	0.36	0.31
Income <100% FPL, age 65+	0.003	(0.009)	0.71	0.11
Log Median Income, Age 65-74	-0.021	(0.042)	0.62	10.33
Log Median Income, Age 75+	-0.036	(0.033)	0.28	10.00
Log Mean House Value	-0.027	(0.057)	0.65	11.75
Renters, Age 65+	-0.009	(0.020)	0.66	0.23
Move Homes, 65-74	-0.002	(0.014)	0.88	0.29
Move Homes, 55-64	-0.002	(0.012)	0.88	0.38
Part B Coverage	0.004	(0.004)	0.24	0.90
Original Medicare Eligibility Through SSDI	-0.001	(0.003)	0.71	0.08
Predicted Medicare Spending	21.4	(34.7)	0.537	6,310

**Notes:** The table above displays regression results for Equation 5. The first column lists the dependent variable, while the second and third columns list the coefficient and standard error, respectively, on the leave-out costs variable (in hundreds). The fourth column displays the associated p-value. The fifth column lists the mean of the dependent variable. ZIP code-level demographic dependent variables are from the Census 2000 Special Tabulation on Aging (available from ICPSR). In the specification labeled, "Part B coverage" the dependent variable is the ZIP code-level fraction of Medicare beneficiaries with Part B coverage. In the specification labeled "Original Medicare Eligibility Through SSDI," the dependent variable is the ZIP code-level fraction of Medicare beneficiaries that originally obtained Medicare coverage through SSDI. These two dependent variables come from the CMS Denominator File, 1999-2005. Zip code-level observations are weighted by the population residing in those ZIP codes. All specifications include HSA fixed effects and robust standard errors clustered at the HSA level. The dependent variable in the final row is the fitted value of an OLS regression of individual-level Medicare Spending on the Census demographics listed in the table and controls used in our baseline specifications (age, sex, race, year, and health risk controls). This fitted value is then regressed on the leave-out instrument and HSA fixed effects, with the resulting coefficient on leave-out costs in the table above. All dollar values are stated in terms of 2005 dollars (inflation-adjusted using the CPI-U).

**Table 4:** Premiums: Regressions of Medigap Premiums on Leave-Out Costs

	Dep. Var.: Medigap Premiums		
	Plans A-J	Plan C	Plan F
	(1)	(2)	(3)
Leave-Out Costs	1.118 (0.111)	0.937 (0.153)	0.974 (0.155)
HSA FE	X	X	X
Insurer FE	X	X	X
Plan FE	X		
R-Squared	0.926	0.841	0.877
N	45,129	6,298	6,449

**Notes:** This table shows estimates from regressions of Medigap premiums on the leave-out costs instrument, HSA fixed effects, and controls (see Section 3, Equation 2). In addition to HSA fixed effects, these specifications contain controls for log GAF/OWI adjustment factors. Observations are at the HSA-state-plan-company level. The first column displays results for plan-level specification that includes all plans offered by United Healthcare and Mutual of Omaha, the two largest insurers. The second and third columns restrict attention to the most popular plans offered by these companies, Plan C and Plan F, respectively. Robust standard errors are clustered at the HSA level. All dollar values are stated in terms of 2005 dollars (inflation-adjusted using the CPI-U).



**Table 5: Demand: Regressions of Insurance Coverage on Leave-Out Costs**

	Leave-Out Costs (Hundreds)			Mean of
	Est	Std. Err.	P-Value	Dep Var
All Beneficiaries				
Denominator File, 1999-2005				
Medicare Advantage	-0.008	(0.006)	0.23	0.12
Medicaid (Dual-Eligible)	0.007	(0.005)	0.19	0.11
MCBS alone				
Supplemental Coverage (HSA level)	-0.066	(0.038)	0.08	0.90
Supplemental Coverage (HRR level)	-0.068	(0.026)	0.01	0.90
Medigap (HSA level)	-0.083	(0.060)	0.17	0.36
Medigap (HRR level)	-0.090	(0.049)	0.07	0.36
NHIS Alone				
Supplemental Coverage (HSA level)	-0.030	(0.026)	0.24	0.79
Supplemental Coverage (HRR level)	-0.009	(0.015)	0.56	0.79
Combined MCBS+NHIS				
Supplemental Coverage (HSA level)	-0.048	(0.023)	0.03	0.85
Supplemental Coverage (HRR level)	-0.039	(0.016)	0.01	0.85

**Notes:** This table shows estimates from regressions of insurance coverage indicators on leave-out costs, HSA/HRR fixed effects, and controls (see Section 3, Equation 3). For the Medicare Advantage and Medicaid specifications, the regressions rely on data from the CMS Denominator file from 1999 to 2005 aggregated to the ZIP code-level where observations are weighted by the population residing within each ZIP code. The analysis for Medigap and any supplemental insurance uses the MCBS and NHIS data from 1992 to 2005. The “Supplemental Insurance” specifications use a dependent variable that indicates if the individual has any coverage beyond the basic FFS Medicare (including Medigap, Medicare Advantage, Medicaid, and RSI). The “Medigap” specifications use a dependent variable indicating Medigap status based on the MCBS data. The HRR level results displayed above are scaled by the premium first-stage coefficient on the instrument at the HRR level to make the reported coefficient comparable to the HSA level results. Our estimates from first-stage at the HRR-level (in Appendix Table C2) show that the coefficient on the instrument ranges from 0.24 to 0.25 depending on the specification; thus, we scale the HRR demand estimates up by a factor of four to make them comparable to the HSA estimates (for which the first-stage coefficient is between 1.1 and 0.94). All specifications include HSA or HRR fixed effects and year fixed effects, in addition to controls for age, sex, and log GAF/OWI adjustment factors. Robust standard errors are clustered at the HSA or HRR level depending on the specification. All dollar values are stated in terms of 2005 dollars (inflation-adjusted using the CPI-U).

**Table 6: Utilization: Regressions of Medicare Utilization on Leave-Out Costs**

Dependent Variable	Leave-Out Costs (Hundreds)			Mean of Dep. Var.	Implied Medigap Effect	
	Est	Std. Err.	P-Value		Level	%
Part B Events	-0.4180	(0.1810)	0.021	25.82	8.71	33.7%
Imaging Events	-0.0812	(0.0323)	0.012	3.99	1.69	42.4%
Testing Events	-0.4090	(0.1470)	0.005	11.42	8.52	74.6%
Total RVUs	-1.2900	(0.4960)	0.009	70.79	26.88	38.0%
Part A Days	-0.0621	(0.0188)	0.001	2.10	1.29	61.6%
Part A Stays	-0.0040	(0.0021)	0.065	0.34	0.08	23.9%
SNF Days	0.0120	(0.0201)	0.552	1.37	-0.25	-18.2%
SNF Stays	0.0003	(0.0008)	0.761	0.06	-0.01	-8.8%

**Notes:** This table displays estimates from regressions of Medicare utilization measures on leave-out costs, HSA fixed effects, and controls (see Section 3 Equation 4). Each row displays the results from a separate regression. The first column lists the dependent variable; the second and third columns list the coefficient and standard error on the leave-out costs variable (in hundreds); the fourth column lists the associated p-value; the fifth column lists the mean of the dependent variable; and the final columns interpret the coefficients in terms of the implied Medigap effect (by scaling the relevant coefficient by the coefficient from the Medigap demand specification). Pooled 1999-2005 data from the CMS Beneficiary Summary File, CMS Denominator File, and CMS Carrier File (for RVU analysis) are used to create these estimates. All specifications include HSA fixed effects, and controls for age, sex, race, health risk, and log GAF/OWI adjustment factors. Robust standard errors are clustered at HSA level. All dollar values are stated in terms of 2005 dollars (inflation-adjusted using the CPI-U).

**Table 7: Payments: Regressions of Medicare Payments on Leave-Out Costs**

Dependent Variable	Leave-Out Costs (Hundreds)			Mean of Dep. Var.	Implied Medigap Effect	
	Est	Std. Err.	P-Value		Level	%
Medicare Payments	-67.02	(33.11)	0.043	6,290	1396.25	22.2%
Part A Payments	-47.59	(22.76)	0.037	3,020	991.54	32.8%
Part B Payments	-21.80	(15.90)	0.159	2,649	454.23	17.1%
SNF Payments	3.44	(5.25)	0.513	399	-71.61	-17.9%

**Notes:** This table displays estimates from regressions of Medicare utilization measures on leave-out costs, HSA fixed effects, and controls (see Section 3, Equation 4). Each row displays the results from a separate regression. The first column lists the dependent variable; the second and third columns list the coefficient and standard error on the leave-out costs variable (in hundreds); the fourth column lists the associated p-value; the fifth column lists the mean of the dependent variable; and the final columns interpret the coefficients in terms of the implied Medigap effect (by scaling the relevant coefficient by the coefficient from the Medigap demand specification). Pooled 1999-2005 data from the CMS Beneficiary Summary File and CMS Denominator File are used to create these estimates. All dependent variables are winsorized at \$64,000. All specifications include HSA fixed effects, and controls for age, sex, race, health risk, and log GAF/OWI adjustment factors. Robust standard errors are clustered at HSA level. All dollar values are stated in terms of 2005 dollars (inflation-adjusted using the CPI-U).

**Table 8: Robustness Checks**

	Leave-Out Costs (Hundreds)			Mean of Dep. Var.	Implied Medigap Effect	
	Est	Std. Err.	P-Value		Level	%
Baseline Specification						
Medicare spending	-67.02	(33.11)	0.043	6,290	1396.25	22.2%
Alternative Specifications (dep var=Medicare spending )						
Census ZIP code-level controls included	-52.30	(30.40)	0.085	6,290	1089.49	17.3%
Region-year fixed effects included	-59.96	(30.16)	0.047	6,290	1249.09	19.9%
Unaffected Procedures						
Urgent RVUs (Clemens & Gottlieb Defn)	0.05	(0.07)	0.421	4.3	-1.13	-26.5%
Urgent admissions (Card, Dobkin, & Maestas Defn 1)	-1.31E-03	(1.03E-03)	0.201	0.077	0.03	35.4%
Urgent admissions (Card, Dobkin, & Maestas Defn 2)	-6.89E-04	(1.03E-03)	0.505	0.125	0.01	11.5%
Unaffected Individuals						
Non-elderly adults in NHIS						
Hospital days	0.03	(0.08)	0.65	0.364	-0.71	-196.3%
Hospital stays	0.01	(0.01)	0.23	0.091	-0.20	-225.3%
Self-reported health	0.02	(0.06)	0.75	1.968	-0.40	-20.5%
Robustness to Spatial Trends						
Medicare spending, restricted to ZIP codes within...						
20 km of border	-65.88	(32.06)	0.040	6,290	1372.50	21.8%
10 km of border	-58.57	(32.53)	0.072	6,290	1220.14	19.4%
Placebo Borders						
Medicare spending	-14.27	(22.97)	0.535	6,290	297.23	4.7%

**Notes:** Each row represents a separate reduced form regression. The first column lists the dependent variable and the relevant alternative sample restrictions (if applicable), while the second and third columns list the coefficient and standard error, respectively, on the leave-out costs variable (in hundreds). The fourth column displays the associated p-value. The fifth column lists the mean of the dependent variable. The final columns interpret the coefficients in terms of the implied Medigap effect (by scaling the relevant coefficient by the coefficient from the Medigap demand specification). The data used for these regressions come from pooling 1999-2005 data from the CMS Beneficiary Summary File, the CMS Denominator File, the CMS Carrier File (for "Urgent RVU" analysis), the NHIS data (the "Unaffected Individuals analysis"), and the CMS MedPAR data (the "Urgent Admissions" analysis). All specifications include HSA fixed effects as well as controls for age, sex, race, health risk, and log GAF/OWI adjustment factors. Robust standard errors are clustered at HSA level except for the "Placebo Borders" specification in which robust standard errors are clustered at the HSA-state level (see text for a full description). All dollar values are stated in terms of 2005 dollars (inflation-adjusted using the CPI-U).

**Table 9: Counterfactuals: Taxing Medigap**

Tax	Medigap	Tax Revenue	Medicare Savings	Total Budgetary Impact	
	Market Share	(per Beneficiary)	(per Beneficiary)	(per Beneficiary)	%
0%	48%	\$0	\$0	\$0	0%
5%	44%	\$39	\$60	\$99	1.6%
10%	39%	\$70	\$119	\$189	3.0%
15%	35%	\$94	\$179	\$273	4.3%
20%	31%	\$110	\$238	\$348	5.5%
30%	22%	\$119	\$358	\$477	7.6%
40%	14%	\$99	\$477	\$575	9.1%
Pigouvian Tax	0%	\$0	\$670	\$670	10.7%

**Notes:** The first column lists the tax as a percent of Medigap premiums, using the average premium of \$1,779 as a baseline. The second column lists the implied Medigap market share assuming full pass-through of the tax. The linear demand curve used in these calculations has a slope equal to  $\partial q_{ijk} / \partial \text{Leave-Out costs}_{jk} = -0.048$  (as the coefficient in the premium regressions was approximately one) and an intercept pinned down by the national market equilibrium ( $p=17.79$  hundreds of dollars and  $q=0.48$ ). The remaining columns list the tax revenue raised, Medicare cost savings from Medigap dis-enrollment, and total budgetary impact, respectively. These results are based on the Medigap externality calculated to be \$1,396. These calculations assume that the ad valorem tax on Medigap is fully passed through to consumers. If instead the pass-through rate is  $\rho < 1$ , then the results displayed above for x% tax on Medigap can be interpreted as the effect of a  $\frac{x}{\rho}\%$  tax on Medigap. All dollar values are stated in terms of 2005 dollars (inflation-adjusted using the CPI-U).

## APPENDIX

### A Medigap Plans: Plan Features and Enrollees by Plan Letter

The form and pricing of Medigap policies are regulated by the federal government. During our sample period, firms were permitted to sell standardized policies labeled A-J. Table A1 describes the features of these different Medigap policies. As one can see, all the policies contain the “basic benefits,” which include coverage of Part A copays and deductibles, Part B coinsurance, blood, and additional lifetime hospital days. Much of the differentiation among the plans is for niche services such as home health care and foreign travel emergencies.

**Table A1: Medigap Benefits by Plan Letter**

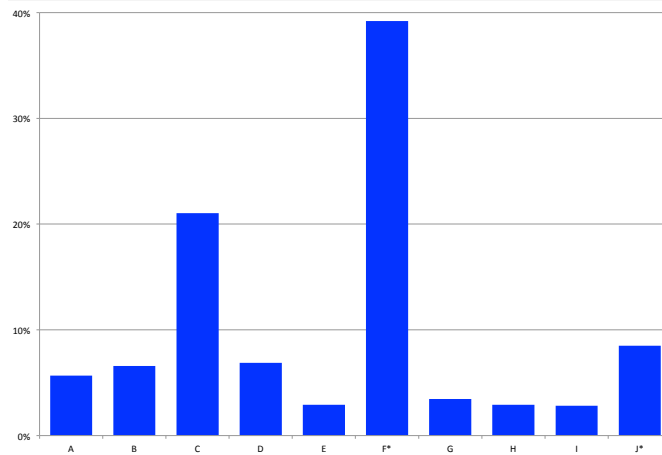
	Medigap Plan Letter									
	A	B	C	D	E	F*	G	H	I	J*
Basic Benefits	X	X	X	X	X	X	X	X	X	X
Part A Copays and Deductible										
Part B Coinsurance										
Blood										
Additional Lifetime Hospital Days										
SNF Coinsurance		X	X	X	X	X	X	X	X	X
Part B Deductible		X				X				X
Part B Excess Charges						X	80%			
Foreign Travel Emergency		X	X	X	X	X	X	X	X	X
Home Health Care				X			X		X	X
Prescription Drugs								X	X	X
Preventive Medical Care				X						X

**Notes:** This table describes the plan features of the different Medigap plans that firms may sell. As one can see from the table, the “basic benefits” are common to all plans. According to federal regulations, firms that offer any Medigap plan must offer Plan A and either Plan C or Plan F. Plans C and F are the most popular Medigap plans sold (see Figure A1 for more information on the relative frequency of Medigap plans purchased).

\*Plans F and J have high-deductible options that require beneficiaries to pay \$1,580 before receiving Medigap benefits in any given calendar year. These plans are rarely offered and have very few enrollees.

Figure A1 illustrates the distribution of Medigap enrollees by plan letter. This distribution is calculated from self-reported Medigap plan letter information from the MCBS (which is reported by roughly half of the respondents who report having Medigap coverage). As one can see, Plan C and Plan F are the most popular plans. Federal government regulations required firms that offered any Medigap policy to offer two options as a subset of the available plans: Plan A and either Plan C or Plan F.

**Figure A1: Medigap Enrollment by Plan Letter**



Notes: The above histogram displays the enrollment of Medigap by plan letter. This histogram is constructed using data from the Medicare Current Beneficiary Survey (MCBS), 1992-2005. A Medigap plan letter is reported by approximately half of the MCBS respondents who report having a Medigap policy.

## B Supplemental Insurance in MCBS and NHIS datasets

To investigate the elasticity of Medigap enrollment, we use data from two surveys: the Medicare Current Beneficiary Survey (MCBS) and the National Health Interview Survey (NHIS). Below, we describe how we translate the variables in these surveys into the insurance dependent variables we use in our estimation. These surveys are used to investigate Medigap and our broader definition of “Supplemental Insurance,” so the coding of these variables is explained below. In the MCBS and the NHIS data, supplemental insurance is inferred based on survey questions.

**MCBS.** MCBS insurance variables are available in the “ric 4” data file for each year. We code individuals as having Medigap if they report having private coverage and report the plan is “self-purchased” and either purchased directly or through AARP. We code individuals as having supplemental coverage if we can infer that they have any source of supplemental coverage, including Medicaid, Medicare Advantage, Medigap, or RSI. Specifically, the following MCBS variables are used in coding individual insurance status: d\_phi, d\_hmo, d\_caaid, d\_obtnp1-5. <sup>75</sup>

**NHIS.** NHIS insurance variables are available in the “personx” data file for each year. Relative to the MCBS, the NHIS has fewer survey questions regarding sources of coverage, and the NHIS survey responses are not verified against administrative data. We code individuals as having supplemental insurance if we can infer that they have any source of supplemental coverage, including Medicaid, Medicare Advantage, Medigap, or RSI. Specifically, the following NHIS variables are used in coding individual insurance status: mchmo, medicare, plnpay21, plnpay22, private, plnwrkn1, plnwrk2, medicaid.

## C Robustness of Demand Results

**Robustness to Alternative Control Variables** In the following table, we display our demand results with alternative sets of controls. The table shows the estimates from the baseline specifica-

<sup>75</sup>The characterization leads to roughly the same market shares as displayed in [GAO \(2001\)](#).

tion for reference (as in Table 5). All specifications include year fixed effects, local medical market fixed effects, and controls for Medicare geographic payment adjustments. The “Fewer Controls” specification includes no additional controls, and the “Baseline Controls” specification includes demographic controls for sex, race, and age. The “More Controls” specification includes demographic controls as well as controls for the incidence of chronic conditions including arthritis, heart disease, diabetes, non-skin cancer, and previous heart attack. The table displays the results for the dependent variables indicating Medigap (in the MCBS) and any supplemental coverage (in the NHIS and the MCBS). Regardless of which set of controls are used, the results are qualitatively and quantitatively very similar. The results indicate that Medigap enrollment is price-sensitive, and the implied elasticity from the combined specification is in the range of -1.5 to -1.8. Within the MCBS, the effects on Medigap and any supplemental insurance are very similar, consistent with the evidence from the administrative data on the lack of substitution into alternative coverage based on our variation.

**Table C1: Demand: Robustness to Alternative Controls**

	Baseline Controls		Fewer Controls		More Controls		Mean of
	Est	Std. Err.	Est	Std. Err.	Est	Std. Err.	Dep Var
All Beneficiaries							
Combined MCBS+NHIS							
Supplemental Coverage (HSA level)	-0.048	(0.023)	-0.046	(0.023)	-0.047	(0.023)	0.85
Supplemental Coverage (HRR level)	-0.039	(0.016)	-0.038	(0.016)	-0.041	(0.015)	0.85
MCBS alone							
Supplemental Coverage (HSA level)	-0.066	(0.038)	-0.068	(0.038)	-0.064	(0.040)	0.90
Supplemental Coverage (HRR level)	-0.068	(0.026)	-0.071	(0.028)	-0.073	(0.028)	0.90
Medigap (HSA level)	-0.0825	(0.060)	-0.0797	(0.064)	-0.0788	(0.060)	0.36
Medigap (HRR level)	-0.090	(0.049)	-0.088	(0.047)	-0.092	(0.048)	0.36
NHIS Alone							
Supplemental Coverage (HSA level)	-0.0302	(0.026)	-0.0253	(0.025)	-0.0321	(0.025)	0.79
Supplemental Coverage (HRR level)	-0.00884	(0.015)	-0.00604	(0.016)	-0.00996	(0.015)	0.79
Controls							
Year and Local Medical Market Fixed Effects	x		x		x		
Demographic	x				x		
Chronic Conditions					x		

**Notes:** This table shows estimates from regressions of insurance coverage indicators on leave-out costs, HSA/HRR fixed effects, and controls (see Section 3, Equation 3). The analysis uses the MCBS and NHIS data from 1992-2005. The “Supplemental Insurance” specifications use a dependent variable that indicates if the individual has any coverage beyond the basic FFS Medicare (including Medigap, Medicare Advantage, Medicaid, and RSI). The “Medigap” specifications use a dependent variable indicating Medigap status based on the MCBS data. The HRR-level results displayed above are scaled by the premium first-stage coefficient on the instrument at the HRR level to make the reported coefficient comparable to the HSA level results. Our estimates from first-stage at the HRR level (in Appendix Table C2) show that the coefficient on the instrument ranges from 0.24 to 0.25 depending on the specification; thus, we scale the HRR demand estimates up by a factor of four to make them comparable to the HSA estimates (for which the first-stage coefficient is between 1.1 and 0.94). All specifications include HSA or HRR fixed effects and year fixed effects, in addition to controls for year, age, sex, and log GAF/OWI adjustment factors. Robust standard errors are clustered at the HSA or HRR level. All dollar values are stated in terms of 2005 dollars (inflation-adjusted using the CPI-U).

**First-stage at HRR level.** The following table presents the first stage at the HRR level. The estimates illustrate that leave-out costs defined at the HRR level are predictive of premiums, with a coefficient around 0.25. Recall that the analogous coefficient at the HSA level was approximately 1. This indicates that the demand results at the HRR level would need to be scaled up by a factor



of four to be comparable with the HSA level estimates. This is what is done in reporting the results for the demand coefficients.

**Table C2:** Premiums: Regressions of Medigap Premiums on Leave-Out Costs at HRR level

	Dep. Var.: Medigap Premiums		
	Plans A-J	Plan C	Plan F
	(1)	(2)	(3)
Leave-Out Costs	0.241 (0.057)	0.244 (0.059)	0.249 (0.068)
HRR FE	X	X	X
Insurer FE	X	X	X
Plan FE	X		
R-Squared	0.917	0.805	0.838
N	44,765	6,246	6,397

**Notes:** This table shows estimates from regressions of Medigap premiums on the leave-out costs instrument, HRR fixed effects, and controls (see Section 3, Equation 2). In addition to HRR fixed effects, these specifications contain controls for log GAF/OWI adjustment factors. Observations are at the HSA-state-plan-company level. The first column displays results for plan-level specification that includes all plans offered by United Healthcare and Mutual of Omaha, the two largest insurers. The second and third columns restrict attention to the most popular plans offered by these companies, Plan C and Plan F, respectively. Robust standard errors are clustered at the HRR level. All dollar values are stated in terms of 2005 dollars (inflation-adjusted using the CPI-U).

## D Control Variables in CMS Data

We include controls for individual chronic conditions in several specifications to improve the precision of our estimates. The chronic condition information comes from the CMS Beneficiary Summary File. The chronic condition controls we include are dummy variables that indicate when the following conditions are present:

- Acute Myocardial Infarction (AMI)
- Alzheimer’s Disease (ALZH)
- Alzheimer’s Disease and Rltd Disorders or Senile Dementia (ALZHDMTA)
- Atrial Fibrillation (ATRIALFB)
- Cataract (CATARACT)
- Chronic Kidney Disease (CHRNKIDN)
- Chronic Obstructive Pulmonary Disease (COPD)
- Heart Failure (CHF)
- Diabetes (DIABETES)
- Glaucoma (GLAUCOMA)
- Hip/Pelvic Fracture (HIPFRAC)
- Ischemic Heart Disease (ISCHMCHT)
- Depression (DEPRESSN)
- Osteoporosis (OSTEOPRS)
- Rheumatoid Arthritis or Osteoarthritis (RA\_OA)

- Stroke or Transient Ischemic Attack (STRKETIA)
- Breast Cancer (CNCRBST)
- Colorectal Cancer (CNCRLRC)
- Prostate Cancer (CNCRPRST)
- Lung Cancer (CNCRLUNG)
- Endometrial Cancer (CNCRENDM)
- Anemia (ANEMIA)
- Asthma (ASTHMA)
- Hyperlipidemia (HYPERL)
- Benign Prostatic Hyperplasia (HYPERP)
- Hypertension (HYPERT)
- Acquired Hypothyroidism (HYPOTH)

The CMS corresponding variable used to derive each of these indicator variables is included in the list above in parentheses after each chronic condition. For more information on the CMS algorithm for determining whether these conditions are present, see the documentation at: <http://www.resdac.org/cms-data/files/mbsf/data-documentation>.

## E Alternative Specifications

The baseline specifications reported in the text include controls for age, sex, race, chronic conditions, and log GAF/OWI adjustment factors. In Table E1, we report the results for the utilization dependent variables when chronic health condition controls are omitted. Overall, the results are qualitatively similar as when the chronic health condition controls are included.

**Table E1:** Utilization: Regressions of Medicare Utilization on Leave-Out Costs, Without Health Controls

Dependent Variable	Leave-Out Costs (Hundreds)			Mean of Dep. Var.	Implied Medigap Effect	
	Est	Std. Err.	P-Value		Level	%
Part B Events	-0.3210	(0.1990)	0.106	25.82	6.69	25.9%
Imaging Events	-0.0561	(0.0337)	0.096	3.99	1.17	29.3%
Testing Events	-0.3400	(0.1710)	0.047	11.42	7.08	62.0%
Total RVUs	-0.9550	(0.4970)	0.055	70.79	19.90	28.1%
Part A Days	-0.0354	(0.0218)	0.105	2.10	0.74	35.1%
Part A Stays	0.0002	(0.0025)	0.931	0.34	0.00	-1.3%
SNF Days	0.0246	(0.0251)	0.327	1.37	-0.51	-37.3%
SNF Stays	0.0009	(0.0011)	0.414	0.06	-0.02	-29.9%

**Notes:** This table displays estimates from regressions of Medicare utilization measures on leave-out costs, HSA fixed effects, and controls (see Section 3, Equation 4). Each row displays the results from a separate regression. The first column lists the dependent variable; the second and third columns list the coefficient and standard error on the leave-out costs variable (in hundreds); the fourth column lists the associated p-value; the fifth column lists the mean of the dependent variable; and the final columns interpret the coefficients in terms of the implied Medigap effect (by scaling the relevant coefficient by the coefficient from the Medigap demand specification). Pooled 1999-2005 data from the CMS Beneficiary Summary File, CMS Denominator File, and CMS Carrier File (for RVU analysis) are used to create these estimates. All specifications include HSA fixed effects, and controls for age, sex, race, and log GAF/OWI adjustment factors. The difference between these results and those presented in Table 6 is that this specification excludes health controls. Robust standard errors clustered at the HSA level. All dollar values are stated in terms of 2005 dollars (inflation-adjusted using the CPI-U).

Table E2 reports the results when the baseline specification for the payment dependent variables is estimated omitting the chronic health condition controls. The results are less statistically

**Table E2:** Payments: Regressions of Medicare Payments on Leave-Out Costs, Without Health Controls

Dependent Variable	Leave-Out Costs (Hundreds)			Mean of Dep. Var.	Implied Medigap Effect	
	Est	Std. Err.	P-Value		Level	%
Medicare Payments	-11.36	(16.76)	0.498	6,290	236.73	3.8%
Part A Payments	-4.70	(22.18)	0.832	3,020	97.87	3.2%
Part B Payments	-9.98	(33.45)	0.766	2,649	207.83	7.8%
SNF Payments	6.90	(6.42)	0.283	399	-143.84	-36.0%

**Notes:** This table displays estimates from regressions of Medicare utilization measures on leave-out costs, HSA fixed effects, and controls (see Section 3, Equation 4). Each row displays the results from a separate regression. The first column lists the dependent variable; the second and third columns list the coefficient and standard error on the leave-out costs variable (in hundreds); the fourth column lists the associated p-value; the fifth column lists the mean of the dependent variable; and the final columns interpret the coefficients in terms of the implied Medigap effect (by scaling the relevant coefficient by the coefficient from the Medigap demand specification). Pooled 1999-2005 data from the CMS Beneficiary Summary File and CMS Denominator File are used to create these estimates. All dependent variables are winsorized at \$64,000. All specifications include HSA fixed effects, and controls for age, sex, race, and log GAF/OWI adjustment factors. The difference between these results and those presented in Table 7 is that this specification excludes health controls. Robust standard errors clustered at the HSA level. All dollar values are stated in terms of 2005 dollars (inflation-adjusted using the CPI-U).

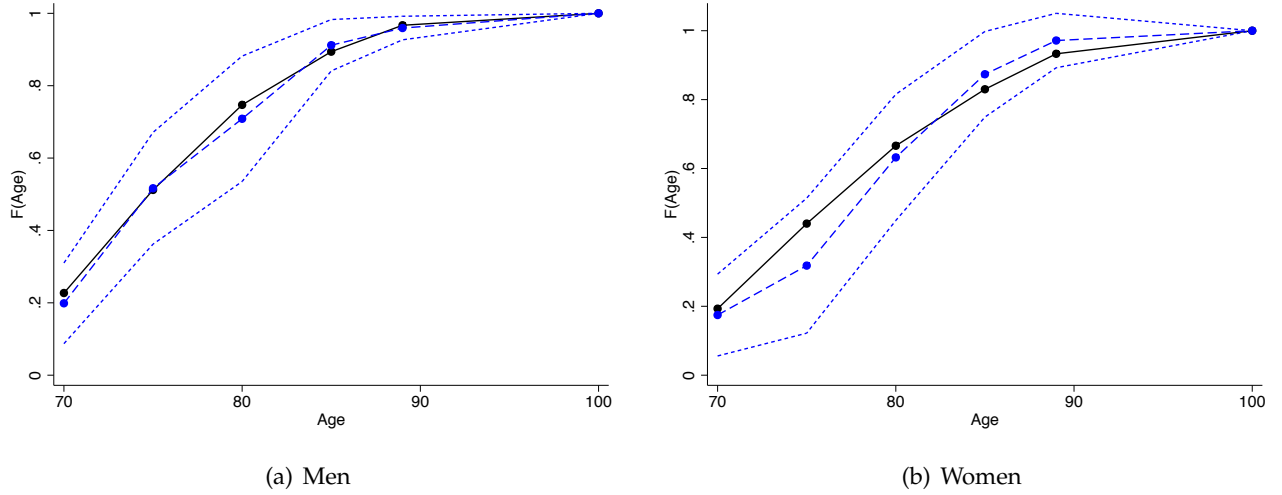
precise when these controls are omitted. However, these results are statistically indistinguishable from the point estimates in the baseline specification. It is perhaps not surprising that the health controls are important for precision as the R-squared increases from 0.03 without health controls to 0.43 with health controls for the “Medicare Payments” specification.

## F Premium Variation and Mortality

We can analyze the effect of Medigap on mortality by examining the effect of the instrument on the cross-sectional age distribution of Medicare beneficiaries. If Medigap reduces mortality, then higher leave-out costs, and the corresponding lower Medigap take-up, should lead to earlier death, shifting the age distribution in an inward direction.

Figure F1 displays the impact of a \$10 increase in leave-out costs on the cross-sectional age distribution. Solid lines show the empirical CDF of payments. Dashed lines show the estimated CDF under a \$10 increase in leave-out costs. These CDFs are constructed using the coefficient on leave-out costs from regressions of the form  $\Pr(\text{Age}_{ijk} < X) = \gamma_c \text{Leave-out costs}_{jk} + \gamma_k + X'_{ijk} \gamma_X + \mu_{ijk}$  for  $X = 70, 75, \dots, 100$ . Dotted lines show the 95% confidence intervals of these estimates. Overall, the plots show that the instrument has no detectable effect on the age distribution of Medicare beneficiaries. Although this evidence is consistent with Medigap having no mortality effect, our research design does not have the power to detect small to moderate effects on mortality

**Figure F1: Effect on Cross-Sectional Age Distribution**



**Notes:** This figure shows the impact of a \$10 increase in leave-out costs on the cross-sectional age distribution. Solid line show the empirical CDF of payments. Dashed lines show the estimated CDF under a \$10 increase in leave-out costs. These CDFs are constructed using the coefficient on leave-out costs from regressions of the form  $\Pr(\text{Age} < X) = \gamma_c \text{Leave-out costs}_{jk} + \gamma_k + X'_{ijk} \gamma_X + \mu_{ijk}$  for  $X = 70, 75, \dots, 100$ . Dotted lines show the 95% confidence intervals of these estimates. All dollar values are stated in terms of 2005 dollars (inflation-adjusted using the CPI-U).

## G Definition of Urgent Procedures

### G.1 Betos Code Characterization from Clemens and Gottlieb (2013)

We follow the categorization used by [Clemens and Gottlieb \(2013\)](#) to group Part B RVUs by BETOS code to determine which procedures are for less discretionary care. Specifically, we define urgent procedures as Part B claims associated with the following BETOS codes:

- P1A: Major procedure—breast
- P1B: Major procedure—colectomy
- P1C: Major procedure—cholecystectomy
- P1D: Major procedure—turp
- P1F: Major procedure—hysterectomy
- P1G: Major procedure—other
- P2B: Major procedure, cardiovascular—aneurysm repair
- P3A: Major procedure, orthopedic—hip fracture repair
- P4A: Eye procedure—corneal transplant
- P4C: Eye procedure—retinal detachment
- P5C: Ambulatory procedure—groin hernia repair
- P7A: Oncology—radiation therapy
- P7B: Oncology—other
- P9A: Dialysis services

## G.2 Weekend Versus Weekday Daily Frequency Characterization from [Card, Dobkin and Maestas \(2009\)](#)

[Card, Dobkin and Maestas \(2009\)](#) characterize urgent hospitalizations by inspecting the weekend versus weekday daily frequency of ICD-9 codes for hospital admissions originating in the ER. We consider two definitions urgent hospitalizations based on this characterization. For the first definition, we define a procedure as urgent if it is listed in Table I of [Card, Dobkin and Maestas \(2009\)](#) as one of the ten highest frequency urgent ICD-9 diagnoses based on their data and characterization. Below is the list of procedures that this first definition encompasses.

- Obstructive chronic bronchitis with acute exacerbation
- Respiratory failure
- AMI of other inferior wall (1st episode)
- AMI of other anterior wall (1st episode)
- Intracerebral hemorrhage
- Chronic airway obstruction, n.e.c.
- Fracture of neck of femur intertrochanteric section
- Cerebral artery occlusion, unspecified
- Convulsions unknown cause
- Asthma, unspecified with status asthmaticus

For the second definition, we apply the same procedure as [Card, Dobkin and Maestas \(2009\)](#) to the 2002 CMS MedPAR data to identify urgent procedures. Specifically, we construct the fraction of hospitalizations originating from the ER during the weekend for each ICD-9 code. We then define a hospitalization as urgent if the T-stat on this fraction being equal to  $\frac{2}{7}$  is less than or equal to 0.3713 (the 10th percentile of the distribution of T-stats).<sup>76</sup> Below are the descriptions of the ten highest frequency ICD-9 codes that are characterized as urgent through this second methodology:

- Escherichia coli infections
- Paralytic ileus
- Home accidents (Accident in home)
- Acute pancreatitis
- Other abnormal blood chemistry (Abn blood chemistry NEC)
- Diverticulitis of colon (without mention of hemorrhage) (Dvrtcli colon w/o hmrhg)
- Infection with microorganisms resistant to penicillins (Inf mcrg rstn pncllins)
- Benign neoplasm of colon (Benign neoplasm lg bowel)
- Other closed transcervical fracture of neck of femur (Fx femur intrcaps NEC-cl)
- Acute myocardial infarction of other inferior wall

## H Robustness of Policy Counterfactuals

Table [H1](#) examines the sensitivity of our estimates of the budgetary effect of a tax on Medigap premiums. To examine robustness to heterogeneity in the price-elasticity of demand, rows of Table [H1](#) recalculate the effect of a 15% tax using the different demand estimates from Table [5](#). We find that across these different estimates, the total budgetary savings to Medicare range from 3.9% to 4.8%. We also show standard errors for each specification. For our baseline estimate of 4.3% total savings, the standard error is 1.7 percentage points.

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<sup>76</sup>Note that this definition of urgent procedures is more conservative than that in [Card, Dobkin and Maestas \(2009\)](#). [Card, Dobkin and Maestas \(2009\)](#) define a procedure as urgent if the T-stat is less than 0.965.

**Table H1: Tax Counterfactuals: Robustness to Alternative Demand Estimates**

Tax	Demand Parameter Used	Medigap Market Share	Tax Revenue (per Beneficiary)		Medicare Savings (per Beneficiary)		Total Budgetary Impact (per Beneficiary)			
			Estimate	SE	Estimate	SE	Estimate (\$)	Estimate (%)	SE (\$)	SE (%)
15%	Supp Cov, Combined Hsa (baseline)	35%	\$94	\$16	\$179	\$88	\$273	4.3%	\$104	1.7%
15%	Supp Cov, MCBS Hsa	30%	\$81	\$27	\$179	\$88	\$260	4.1%	\$115	1.8%
15%	Supp Cov, NHIS Hsa	40%	\$107	\$18	\$179	\$88	\$286	4.5%	\$107	1.7%
15%	Medigap, MCBS Hsa	26%	\$69	\$43	\$179	\$88	\$248	3.9%	\$131	2.1%
15%	Supp Cov, Combined Hrr	38%	\$100	\$11	\$179	\$88	\$279	4.4%	\$100	1.6%
15%	Supp Cov, MCBS Hrr	30%	\$79	\$19	\$179	\$88	\$258	4.1%	\$107	1.7%
15%	Supp Cov, NHIS Hrr	46%	\$122	\$11	\$179	\$88	\$301	4.8%	\$99	1.6%
15%	Medigap, MCBS Hrr	24%	\$64	\$35	\$179	\$88	\$243	3.9%	\$123	2.0%

**Notes:** The first column lists the tax as a percent of Medigap premiums, using the average premium of \$1,779 as a baseline. The second column describes which demand estimate from Table 5 is used in the calculation. The third column lists the implied Medigap market share assuming full pass-through of the tax. The linear demand curve used in these calculations has a slope equal to  $\partial q_{ijk} / \partial \text{Leave-Out costs}_{jk}$  (as the coefficient in the premium regressions was approximately one) and an intercept pinned down by the national market equilibrium ( $p=17.79$  hundreds of dollars and  $q=0.48$ ). The remaining columns list the tax revenue raised, Medicare cost savings from Medigap dis-enrollment, and total budgetary impact, respectively. The associated standard errors are also displayed above. To calculate the standard error on the total budgetary savings, we first separately calculate the standard error on the tax revenue raised (from the corresponding demand estimate) and the standard error from the Medicare cost savings from Medigap dis-enrollment (from the reduced form cost estimates). We then obtain the standard error on the total savings by aggregating these standard errors using the Delta Method assuming no covariance between the demand and cost estimates. These calculations assume that the ad valorem tax on Medigap is fully passed through to consumers. If instead the pass-through rate is  $\rho < 1$ , then the results displayed above for an  $X\%$  tax on Medigap can be interpreted as the effect of an  $\frac{X}{\rho}\%$  tax on Medigap. All dollar values are stated in terms of 2005 dollars (inflation-adjusted using the CPI-U).