Does Privatized Medicare Benefit Patients or Producers?
Evidence from the Benefits Improvement and Protection Act

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Abstract

The debate over privatizing Medicare stems from a fundamental disagreement about whether privatized Medicare would primarily generate consumer surplus for individuals or producer surplus for insurance companies and health care providers. This paper investigates this question by studying an existing form of privatized Medicare called Medicare Advantage. Using differences-in-differences variation in Medicare Advantage payments brought about by payment floors established by the 2000 Benefits Improvement and Protection Act, we find that for each dollar in increased payments, Medicare Advantage plans reduced premiums by 29 cents and increased the actuarial value of benefits by 9 cents, for a combined pass-through rate of less than 40%. Low pass-through could potentially be explained by advantageous selection or market power. Analysis of Medicare costs shows that very little of the low pass-through can be accounted for by selection, suggesting that market power among insurers is an important factor determining the incidence of payments.

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1 Introduction

Medicare is the second largest social insurance program in the United States and the primary source of health insurance for the elderly. In 2011, Medicare spent $554.3 billion on health care, a 6.2% increase over the previous year.\(^1\) Given the large scale of the program and rapid growth in spending, reforming Medicare is a perpetual policy issue.

One commonly discussed reform proposal is the privatization of Medicare. Proponents of privatization argue that it would reduce costs by encouraging competition among private insurers, and it would allow individuals to select coverage that better matches their preferences. Opponents of privatizing Medicare argue that such a move would lead to large profits for private insurers and health care providers and the eventual erosion of insurance benefits. At its core, the debate is about economic incidence: Does privatized Medicare primarily generate consumer surplus for individuals or producer surplus for insurance companies and health care providers?

This paper investigates this question by studying an existing form of privatized Medicare called Medicare Advantage.\(^2\) In most regions of the country, Medicare beneficiaries can choose to be covered by public fee-for-service Traditional Medicare or to obtain subsidized coverage through their choice of a private Medicare Advantage (MA) insurance plan. MA plans are differentiated from Traditional Medicare in having restricted provider networks and alternative cost-sharing arrangements, and have traditionally been offered by health maintenance organizations (HMOs). These plans receive a capitation payment from the Medicare system for each enrolled beneficiary and typically charge beneficiaries a supplemental premium. The current system is therefore quite similar to recent proposals to expand the private provision of Medicare through a system of “premium supports.” Like the current system, such privatization proposals typically include requirements for a minimum level of basic benefits and a traditional fee-for-service coverage option.\(^3\)

We examine the incidence of privatized Medicare on consumer and producer surplus by studying a sharp change in capitation payments to MA insurers brought about by the 2000 Benefits Improvement and Protection Act (BIPA). Baseline MA capitation payments are determined at the county level

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\(^2\) During our sample period, this private option was called Medicare Part C or Medicare+Choice. Since the passage of the Medicare Modernization Act in 2003, these plans have been called Medicare Advantage so throughout this paper we use the current naming convention.

\(^3\) For recent examples, see the 2012 Burr-Coburn plan or the 2014 Ryan proposal.
using a moving average of Traditional Medicare expenditures in the county over a five-year window. BIPA reformed this payment system by instituting a system of payment floors that raised payments in 72% of the counties in our data. We show that MA capitation payments in the counties below these payment floors were on parallel trends before the payment reform, but payments in the affected counties increased by an average of about $600 per beneficiary per year or 12% when BIPA was implemented, providing us with a source of difference-in-differences variation.

Using this difference-in-differences variation, we find that that MA plans passed-through a low share of their capitation payment increases. For each dollar in higher payments, we find that consumer premiums were reduced by 29 cents and the actuarial value of plan benefits were increased by 9 cents in the 3 years following the reform. These effects are precisely estimated, with 95% confidence intervals that rule out a combined pass-through effect greater than 59 cents on the dollar under a variety of empirical specifications. Difference-in-differences plots that flexibly allow the effect of BIPA on premiums to vary by year show no impacts on premiums in pre-reform years, providing evidence in support of the parallel trends identifying assumption. We show the effects are robust to specifications that isolate different sources of variation in payments brought about by BIPA.

We then investigate why consumers receive so little surplus on the margin from privatized Medicare. Drawing on prior work by Weyl and Fabinger (2013) and Mahoney and Weyl (2013), we build a model that illustrates that the observed low pass-through could potentially be explained by two factors: the degree of advantageous selection in the market and the market power of private MA insurance plans. If there is substantial advantageous selection into MA, then private plans will not pass-through the payments in reduced premiums because they will attract enrollees that are differentially high cost on the margin. If firms have market power, then they may not face pressure to pass-through increased payments through lower premiums or reduced cost-sharing.

We estimate the degree of advantageous selection into MA by estimating the slope of the Traditional Medicare cost curve using spending data for Traditional Medicare beneficiaries. Using the same difference-in-differences research design, our estimates indicate little advantageous selection into MA on the margin from the BIPA payment reform. Within our theoretical framework, the selection estimates imply that the low pass-through is primarily due to market power among MA insurers. We provide several pieces of additional evidence consistent with the finding that market power is an important determinant of pass-through, including showing that counties with a greater number of
MA contracts in the pre-reform period passed-through a larger share of the payment increases.

Our paper makes three contributions. First, we provide estimates of the pass-through of MA payments into premiums and benefits. The estimates complement work by Duggan, Starc and Vabson (2014) on MA conducted in parallel to our study.\(^4\) The estimates also contribute to a broader empirical literature on pass-through such as Atkin and Donaldson (2014).

Second, we provide estimates of selection between MA and Traditional Medicare. A major obstacle to studying selection into MA is that cost data is not reported for MA beneficiaries. As explained in detail in Section 6, our identification strategy allows us to use cost data on Traditional Medicare beneficiaries to estimate the degree of selection into MA among those marginal to the BIPA capitation payment increase.\(^5\) In contrast to the well-established fact that selection into MA is advantageous overall, we find little advantageous selection into MA along the margin of our variation in enrollment.

Third, our results contribute to the central policy debate on privatizing Medicare. Our finding that consumers receive only a small share of the surplus from increased Medicare payments raises questions about the efficacy of privatizing Medicare. The fact that selection makes only a very small contribution to this result suggests that improvements in risk-adjustment since our period of analysis are unlikely to have improved the situation. Although evaluating the merits of specific policy proposals are outside the scope of our analysis, our estimates indicate that efforts to make the market more competitive may be key to increasing consumer surplus.

The remainder of the paper proceeds as follows. Section 2 provides some background on MA payments and describes our data. Section 3 presents our empirical strategy. Section 4 reports estimates of pass-through. In Section 5, we present the model that allows us to investigate the determinants of pass-through. In Section 6 we empirically evaluate the roles of selection and market power in explaining the low pass-through [COMING SOON]. Section 7 concludes.

\(^4\)Our work is also related to Town and Liu (2003) who estimate consumer and producer surplus generated by MA using a logit discrete choice framework assuming no adverse selection.
\(^5\)The results of these studies suggest that MA plans are “overpaid” relative to what beneficiaries would utilize in Traditional Medicare (Brown et al., 2011; McWilliams, Hsu and Newhouse, 2012a). Prior research on MA has been limited to estimating correlations between spending and coverage among those that choose to switch between MA and Traditional Medicare.
2 Background and Data

2.1 Medicare Advantage Payments

Private Medicare Advantage insurance plans are given monthly capitated payments for each enrolled Medicare beneficiary. Insurers can supplement these payments by charging premiums directly to enrollees. Baseline capitated payments are determined by historical average monthly costs for the Traditional Medicare program in the enrollee’s county of residence, and a statutory minimum payment.\(^6\)

Let \( j \) denote counties and \( t \) denote years. Baseline MA payments \( b_{jt} \) are given by

\[
 b_{jt} = \max \left\{ \frac{1}{5} \sum_{\tau=t-7}^{t-3} \bar{c}_{jt\tau}, \ b_{jt} \right\} 
\]

where the maximand is taken over the five-year moving average of county fee-for-service costs \( \bar{c}_{jt\tau} \), lagged by three years and the minimum payment rate \( b_{jt} \), which depends on whether county \( j \) is urban or rural and the time period \( t \).

The final capitation payment received by MA insurers is determined by multiplying the county baseline rate by a risk adjustment factor to account for the relative riskiness of those individuals enrolling in MA versus Traditional Medicare. Prior to 2000, this adjustment was done using age and sex. From 2000 through 2003, the risk adjustment formula began to additionally place a small weight on inpatient diagnoses. Overall, the risk adjustment done prior to 2004 explained no more than 1.5% of the variation in medical spending.\(^7\) Extensive risk adjustment of MA capitation payments was introduced in 2004 (see Brown et al., 2011; McWilliams, Hsu and Newhouse, 2012b).

\(^6\) The Balanced Budget Act of 1997 (BBA) set the first minimum payment floors, which applied to rural counties only beginning in 1998. The Benefits Improvement and Protection of 2000 (BIPA) raised the rural floor and added an urban floor. The Medicare Modernization Act of 2003 (MMA) increased payments to all counties, reducing the extent to which BIPA floors were binding.

\(^7\) Between 2000 and 2003, ninety percent of the payment adjustment was based on sex and age, while 10% was based on inpatient diagnoses, if any. This mixture explained approximately 1.5% of the variation in medical spending (Brown et al., 2011), and its purpose was not to correct for geographic variation in illness or utilization, which is fully captured in the local county average, but to address sorting between TM and MA. Following the prior literature, we abstract for the risk adjustment based on inpatient diagnoses during this time period and focus solely on the demographic risk adjustment in our analysis.
2.2 Benefits Improvement and Protection Act (BIPA)

Our source of identifying variation arises from the Benefits Improvement and Protection Act of 2000 (BIPA), which implemented two new baseline payment floors in March 2001: one floor applicable to rural counties and one floor for urban counties. Counties already receiving payments in excess of the floors received a uniform 1% increase in their payment rates.

The historical context for BIPA was a contraction in the MA program in the late 1990s. In 1998, the Balanced Budget Act of 1997 (BBA) put in place growth caps for payments to MA plans with the aim of limiting cost growth. As a result, enrollment growth in the MA program slowed, and between 1999 and 2000 the number of MA enrollees shrunk for the first time since the program’s inception in 1985. Under pressure from insurers to reverse the payment cuts (Achman and Gold, 2002), Congress passed BIPA in December of 2000.

2.3 Data

Most of our analysis relies on publicly available administrative data on the Medicare Advantage program. We combine data from several sources: MA rate books, which list the administered payment rates for each county in each year; the census of MA insurer contracts offered by county-year; county-level MA enrollment summaries; and plan premium data for every contract.\(^8\) For some years we can supplement our data with detailed information on the benefits and other financial characteristics of each plan.\(^9\)

We focus on the 8 year period from 1996 to 2003, which provides us with 5 years of data from before the passage of BIPA and 3 years of data after the bill was signed into law. We end our sample at 2003 to avoid confounding factors introduced by the 2004 implementation of the Medicare Modernization Act of 2003 (MMA), which reformed the capitation payment system extensively.\(^10\)

Table 1 displays summary statistics for the pooled 1996 to 2003 sample. Panel A shows values for the full set of 3,143 counties. Panel B restricts the sample to the 2,257 counties that were affected

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\(^8\)Plan premium sources vary by year and include the Medicare Compare database, the Medicare Options Compare database, and an Out of Pocket Cost database provided by CMS.

\(^9\)The detailed descriptions of plan benefits are sometimes referred to as landscape files or plan services files. These data are only currently available beginning in 2000.

\(^10\)The 2003 Medicare Modernization Act changed the formula by which the baseline payment is calculated substantially. In addition, the act introduced meaningful risk-adjustment applied on top of the baseline rate to calculate the overall capitation payment. Several prior papers examine the effects of various aspects of the 2003 Medicare Modernization Act reform including Brown \textit{et al.} (2011), McWilliams, Hsu and Newhouse (2012\textit{b}), and Woolston (2012).
by the BIPA payment floors. The table shows statistics calculated on county level data, so that for example the minimum base rate in the table is the lowest county level base rate in our sample. The observations are weighted by the number of Medicare beneficiaries per county so that statistics reflect the experience of the average Medicare beneficiary.

Panel A shows that base rates average $541 per month for all counties, but range from $380 to $873 per month across the sample. More than 70% of Medicare beneficiaries live in a county with at least one plan, and more than two plans are available to the average beneficiary. The statistics on plan premiums and benefits are calculated for the sample of counties with at least one plan in all years. Premiums average $32 per month and also vary substantially. Copays for physician and specialists visits average $10 and $14 per visit, respectively. Approximately two-thirds of plans offer drug and dental coverage, and one-third cover hearing products and eyewear.

Panel B shows that base rates average $486 per month for the payment floor counties, approximately 10% less than the average in the full sample. Slightly more than 50% of beneficiaries live in a county with at least one plan, and the average number of plans is slightly more than one. For counties with a plan, average premiums are higher, averaging $40 per month. Benefits are moderately less generous. For instance, only 55% of plans in these counties cover prescription drugs, compared to 69% in the full sample.

3 Research Design

In this section we present the research design we use to examine the effects of BIPA on outcomes like premiums and plan benefits. We start by showing descriptive evidence of the change in payments and then present our econometric model.

3.1 Identifying Variation

Figure 1 plots payments for each county in February and March of 2001, the month before and after the payment floors came into effect. The left panel shows the full sample; the right panel magnifies the region of interest. The figure shows that the BIPA lead to a sharp increase in payments, with urban counties having their base rates raised to a minimum of $525 per month and rural counties having their base rates raised to a minimum of $475 per month.
Figure 1 also illustrates the two key sources of variation that we use in our analysis. The first source of variation arises from the fact that counties with the same base rate in February 2001 received different payment increases depending on their urban or rural status, with urban counties receiving increases of $50 per month more than rural counties with the same February 2001 base rate level. The second source of variation arises from the fact that counties with the same urban or rural status received different payment increases depending on their February 2001 base rate level, with, for example, affected urban counties with lower base rates receiving relatively larger payment increases than affected urban counties with higher February 2001 base rate levels.

Table 2 provides some basic statistics on the increase in payments. On average, the payment floors lead to a 11.2% payment increase in affected rural counties and a 12.8% increase in affected urban counties. There was substantial variation, for example, with the bottom quantile of urban floor counties received a payment increase below 6.7% and the top quantile received an increase above 17.8%.

3.2 Econometric Model

We examine the effects of this rate change using a differences-in-differences research design that compares outcomes for counties that received rate changes of varying sizes due to the BIPA payment floors. Let \( j \) indicate counties and \( t \) indicate years. We measure exposure to BIPA with a county-level distance-to-the-floor variable defined as the difference between the pre-period monthly payment in 2000, \( b_{j,2000} \), and the relevant March 2001 payment floor, \( b_{j,2001} \):

\[
\Delta b_j = \max\left\{b_{j,2001} - b_{j,2000}, 0\right\}. \tag{2}
\]

Our baseline econometric model is a difference-in-differences specification that allows the coefficient on the distance-to-the-floor variable \( \Delta b_j \) to flexibly vary by year. Let \( y_{jt} \) be an outcome in county \( j \) in year \( t \). Our baseline regression specification takes the form:

\[
y_{jt} = \sum_{t \neq 2000} \beta_t \cdot \Delta b_j + \delta_j + \delta_t + f(X_{jt}) + \epsilon_{jt} \tag{3}
\]

where \( \delta_j \) and \( \delta_t \) are county and year fixed effects, \( f(X_{jt}) \) is a flexible set of controls, and \( \epsilon_{jt} \) is the error term. We normalize \( \beta_{2000} = 0 \) so that the \( \beta \)'s can be interpreted as the change in the outcomes relative
to the year 2000 when BIPA was passed. This normalization is not particularly restrictive since the 
econometric approach allows us to observe any anticipatory or delayed response to the legislation.

The identifying assumption for this differences-differences research design is the parallel trends 
assumption: in the absence of BIPA, outcomes for counties that were differentially affected by the 
payment floors would have evolved in parallel. We have two broad approaches to assess the validity 
of this assumption. Our first approach is to plot the β coefficients over time. This approach allows us 
to visually determine whether there is evidence of a spurious pre-existing trend, as well as to observe 
any anticipatory or delayed response to the payment increases from BIPA.

Our second approach is to estimate specifications that isolate the two key sources of variation in 
our data. For all of our outcomes, we estimate a specification that includes as a control the base rate 
in year 2000 interacted with a linear time trend. This approach controls for any differential trends 
over time across counties with different base rate levels, such as differential medical cost growth or 
regression to the mean. With these controls, the estimates are largely identified by differences in the 
payment increases across urban and rural counties with the same base rate levels.

We also estimate specifications where we take the complementary approach of including as a 
control the urban status of the county interacted with a linear time trend. This approach controls for 
differential trends over time across urban and rural counties. With these controls, the estimates are 
largely identified by differences in the size of the payment increase within the sets of urban and rural 
counties.

Figure 2 shows the effect of the change in payments on monthly payment rates, plotting the co-
efficients on distance-to-the-floor from the difference-in-differences specification of Equation 3 with 
the base rate as the dependent variable. The figure shows coefficients from our baseline specification 
with county and year fixed effects as well as the base rate in 2000 interacted with a linear time trend. 
Table 3 shows the corresponding regression estimates from the baseline specification as well as spec-
fications with alternative controls. Both the figure and table show that a dollar increase in payments 
translates one-for-one into a change in county payments. In the remainder of the paper, we interpret 
reduced form effects of distance-to-the-floor on outcomes such as premiums and benefits as resulting 
from a one-for-one change in county monthly payments.
4 Main Results

In this section, we examine the effects of the increase in payments on premiums and plan characteristics. We start by presenting the results on premiums. We then examine the effects on plan benefits, such as copayments and drug coverage, as well as a variable that monetizes these benefits by combining them with data on utilization.

4.1 Pass-through into Premiums

Figure 3 examines the effect on premiums by plotting the coefficients on distance-from-the-floor from difference-in-differences regressions (Equation 3) with different measures county-level premiums as the dependent variable. The dashed horizontal line at -1 indicates the benchmark of full pass-through, which occurs when a dollar increase in payments translates one-for-one into a dollar decline in premiums. The dashed horizontal line at zero indicates the benchmarked of no pass-through. The figures shows coefficients from our baseline specification with county and year fixed effects, as well as the year 2000 base rate interacted with a linear time trend, which captures any differential trends among counties in the pre-BIPA period. Standard errors in all specifications are clustered by county. The capped vertical bars show 95% confidence intervals.

Panel (a) shows estimates from a specification with mean county-level premiums as the dependent variable. The plot shows no evidence of a trend in the pre-BIPA period, providing support for our parallel trends identifying assumption. Following BIPA, the plot shows modest evidence of pass-through. In 2001, the first year following the introduction of payment floors, mean premiums declined by 14 cents for each dollar of increased payments. Premiums for 2001 were set prior to the resolution of uncertainty surrounding the passage of BIPA, though the pattern suggests that insurers who were lobbying for BIPA correctly anticipated its passage. In 2002, the first plan year for which the BIPA payment changes could have been fully priced into MA plans, premiums affected by the payment floors drop further, yielding a pass-through of 26 cents. In 2003, this pass-through increased again to 29 cents. The ramp-up over time may reflect the dynamics of convergence to a medium-run response, which may be larger than the immediate response since more factors are adjustable in the medium-run.

Panels (b), (c), and (d) of Figure 3 show difference-in-differences plots with the county-level median, minimum, and maximum premiums as the dependent variable. Plots follow the same general
pattern of Panel (a), showing no evidence of pre-trends in the years before BIPA and statistically significant pass-through in the post-period. The effects are smaller for minimum premiums and larger for maximum premiums but always show pass-through effects of substantially less than than one-for-one. Indeed, across all of the plots we are able to rule out pass-through of more than 50 cents on the dollar with a 95% confidence interval.

Table 4 presents parameter estimates from corresponding difference-in-differences regressions. In addition to the baseline specification, which include year 2000 base rates interacted with a linear time trend as an additional control, the table shows specifications with only county and year fixed effects and specifications that include an urban indicator interacted with a linear time trend as an alternative set of controls. With the alternative sets of controls, the point estimates are slightly smaller than the baseline specification, but in no case statistically different from these estimates. Importantly, pass-through is substantially less than full in all of the specifications.

4.2 Pass-through into Benefits

It is possible that insurers adjusted plan characteristics other than premiums in response to payment floors. If benefit packages became more generous, it would suggest insurers passed-through the BIPA base rate changes partially in the form of lower cost sharing or newly-covered services. If plans became less generous, than the already low premium pass-through we find would be an overestimate. We explore effects on the benefits package in two ways, first analyzing changes to various plan characteristics and then monetizing the changes by combining them with data on utilization.

We begin by examining a set of plan characteristics likely to be particularly salient to consumers choosing among plans. Recall that MA plans are differentiated from traditional Medicare primarily by offering lower out-of-pocket payments via smaller copays and by offering fringe benefits that are unavailable in Traditional Medicare. Therefore, we analyze the impact of BIPA payments on personal physician and specialist copays and on the likelihood that plans offered coverage for dental, vision (contacts and glasses), hearing aids, and prescription drugs. These fringe benefits were excluded from Traditional Medicare during our study period.

\footnote{In general, for a small change in production costs, an envelope condition predicts that price, but not other characteristics of plans, should respond. However, the changes in net costs due to BIPA were relatively large, suggesting that insurers may have adjusted their benefits package.}

\footnote{Traditional Medicare offered minimal coverage in some of these categories. For example, it covered prescription drugs administered during hospitalizations but not otherwise, and it covered eyeglasses following a cataract surgery only. We examine effects on coverage of these items above and beyond the standard benefits.}
Figure 4 plots the coefficients on distance-from-the-floor from difference-in-differences regressions (Equation 3) with measures of plan benefits as the dependent variable. Panel (a) and Panel (b) examine effects on mean copays for office visits with personal physician and with specialists. Data on plan benefits exists for only one pre-reform year, 2000, which is sufficient for identifying effects, but does not allow us to examine whether there are pre-existing trends. To aid interpretation, in these plots we scale the coefficient on the distance-to-the-floor variable by $50, which is approximately the average base rate increase for affected counties and also approximately 10% of the $523 mean pre-BIPA base rate in our data.

Panels (a) and (b) show that the increase in payments had a sharp effect on average copayments for physician and specialist visits. By 2003, the $50 or 10% increase in monthly payments reduced physicians copayments by 45% or $3.34 on a pre-BIPA base of $7.37 and reduced specialist copayments by 48% or $4.86 on a pre-BIPA base of $10.12. However, these large elasticities do not imply large economic effects. The average Medicare beneficiary has 8 combined physician and specialist visits per year or two-thirds of a visit per month, implying that the $50 increase in monthly payments translates into an expected reduction in copayments of approximately $3 per month.

Panels (c) to (f) of Figure 4 plot coefficients from analogous specifications where the dependent variable is the percent of plans that offer drug, dental care, hearing aids, and vision coverage. As before, the effects are scaled to a $50 increase in monthly payments. The plots show that increased payments have no effect for drug and dental coverage but relatively large effects on the percent of plans offering hearing aids and vision coverage. By 2003, the point estimate indicate that a $50 or 10% increase in payments raises the share of plans offering hearing aids by 17 percentage points of a base of 43% and the share of plans offering vision coverage by 11 percentage points on a base of 73%.

Table 5 displays parameter estimates from the corresponding difference-in-differences regressions. Means of the dependent variables and implied elasticities are included at the bottom of the table. For each dependent variable, the table shows the baseline specification which includes as a control the year 2000 base rate interacted with a linear time trend and an alternative specification which controls for an urban indicator interacted with a linear time trend. All specifications include county and year fixed effects. The effects are consistent across these different specifications.

In order to understand how these benefits changes the actuarial value of this insurance, we combine these estimates with data from the Medical Expenditure Panel Survey (MEPS). The MEPS con-
Preliminary and Incomplete information on utilization and insurer- and patient-paid portions of medical expenditure by category. We use data from the MEPS 2000, restricting the sample to individuals who are at least 65 years of age. For each of the fringe benefits (dental, vision, hearing aids, and drugs), we find total spending and the insurer-covered portion among respondents with non-zero insurer claims in order to estimate service-specific coinsurance rates. We then multiply these rates by the unconditional total spending in each category, generating actuarial values of coverage for each fringe benefit. For copays, we simply multiply the copay amount by the number the average annual number of physician visits. Finally, we sum across all categories and divide the measure by 12, since the utilization and expenditure tallies in the MEPS are annual and our payment floor variation is in monthly rates. We describe the exercise in fuller detail in the appendix. The procedure delivers a monetized measure of plan generosity that reflects the actuarial value of the benefits.

Figure 5 plots effects on this measure of the actuarial value of benefits. The vertical axis offers the same pass-through interpretation as in the premium figures, where a coefficient of 1 corresponds to a dollar increase in plan benefits for a dollar increase in plan subsidies due to BIPA. Pass-through is small and statistically significant only in 2003. The point estimates indicate that for each dollar in increase payments 9 cents is passed-through to beneficiaries in the form of more generous benefits. Tables 5 shows the the corresponding regression estimates for the baseline and alternative specifications. The effects are very stable across the different sets of identifying variation and confirm the finding that a dollar increase in monthly payments translates into a 9 cent increase in the actuarial value of plan generosity.

4.3 Plan Availability

Since floor counties saw a differential rise in capitation payments, we might expect MA insurers would provide more plans in floor counties relative to non-floor counties after BIPA was implemented. We investigate the effect of BIPA on plan availability in Figure 6. Panel (a) plots the effect of a $50 increase in the monthly payment on the presence of at least one plan in a county. The results indicate that there was no effect on the presence of at least one plan. In other words, BIPA did not improve access to MA among beneficiaries that would have otherwise had no MA option in their county.

Although BIPA had no impact on whether an MA plan was offered, it could be that BIPA had an
impact on the number of plans offered. We investigate this in Panel (b), which plots the effect of a $50 increase in the monthly payment on the number of plans offered within a county. Here we find some effect. A $50 increase in monthly payments increased the number of MA plans offered by 0.13 in the first year of implementation and 0.35 by three years after implementation. The results ramp up over time, which aligns with intuition that entry-and-exit decisions may happen over the medium-run. Table 7 displays coefficients from the baseline and alternative specifications. The results are stable across the specifications, with an implied elasticity on the number of plans ranging from 1.42 to 1.65.

5 Model of Pass-Through

In the previous section, we showed that MA plans pass-through only a small share of increased capitation payments in the form of lower premiums and more generous benefits. In this section, we show that limited pass-through can possibly be explained by (i) advantageous selection into MA and (ii) market power among MA insurers and medical providers. To build intuition, we start by presenting simplified graphs that illustrate these forces. We then present a model which, under assumptions on the nature of selection and competition, allows us to generate quantitative predictions on the relationship between these forces and pass-through. The model provides a quantitative framework for interpreting the empirical evidence which follows.

5.1 Graphical Analysis

Figure 7 presents this graphical analysis. We model demand for MA as linear, and we define the marginal cost of providing MA to an individual as the expected cost of providing medical care net of the capitation payment from Medicare. Within this framework, we can depict the increase in capitation payments under BIPA as a downward shift of the the marginal cost curve. Our graphical approach is closely related to Einav, Finkelstein and Cullen (2010) who examine selection in a perfectly competitive environment and Mahoney and Weyl (2013) who examine the interaction of imperfect competition and selection.

Panel A of Figure 7 examines the impact of selection on pass-through in a perfectly competitive market. In a perfectly competitive market, firms earn zero profits and the equilibrium is defined by the intersection of the demand and the average cost curves. When there is no selection, firms
face a horizontal average cost curve, and a downward shift in the average cost curve translates one-for-one into a reduction in premiums, depicted by the transition from the point A to the point B in the figure. When there is advantageous selection, average costs are upward slopping, as the marginal consumer is more expensive than the average. Panel A illustrates that under advantageous selection a identically sized downward shift in the average cost curve is not fully passed through, as firms offset the higher costs of the marginal consumers with higher prices to maintain zero profits in equilibrium.

Panel B examines the impact of market power on pass-through in a market with no selection. To simplify the exposition, we consider the extremes of perfect competition and monopoly. As described above, when there is perfect competition and no selection, a downward shift in the marginal cost curve is fully passed through to consumers, moving the equilibrium from point A to point B. The monopolist sets price such that marginal revenue is equal to marginal cost. With a linear demand curve, this leads to 50% pass-through, shifting the equilibrium from point C to point D in the figure. More generally, Bulow and Pfleiderer (1983) show that the pass-through of a small cost shock in this setting is determined by the ratio of the slope of the demand curve to the slope of the marginal revenue curve.

5.2 Model

We build on and generalize this graphical analysis by constructing a model of pass-through in imperfectly competitive selection markets, drawing upon previous work by Weyl and Fabinger (2013) and Mahoney and Weyl (2013). We direct the reader to these papers for technical details and micro-foundations that support the modeling choices.

Consider an industry in which symmetric firms compete to offer symmetric products. Let $b$ denote the per-beneficiary capitation payment and $p$ denote the additional premium paid by each beneficiary. In equilibrium, all firms charge the same price, with aggregate demand at this price given by $q(p)$. Total costs for the industry are summarized by an aggregate cost function $c(q)$ which is equal to the aggregate medical costs paid on behalf of the individuals enrolled in Medicare Advantage, and marginal costs for the industry as given by $mc(q) \equiv c'(q)$. Adverse selection at the industry level is indicated by decreasing aggregate marginal costs $mc'(q) < 0$ and advantageous selection is indicated by increasing aggregate marginal costs $mc'(q) > 0$.

When a single firm lowers its price, it attracts consumers that are new to the market and con-
consumers who are already purchasing the product from competing firms. Consumers who are new to the market have marginal costs equal to industry marginal costs $mc(q)$. Following Mahoney and Weyl (2013), we assume that consumers acquired from rivals are not selected and have costs equal to industry average cost $ac(q) = \frac{c(q)}{q}$. Perceived marginal costs for a single firm $mc_i(q)$, indicated by the $i$ subscript, are the weighted sum of marginal costs for consumers that are new to the market and marginal costs for consumers that are attracted from other firms:

$$mc_i(q) = \theta mc(q) + (1 - \theta) ac(q)$$  \hfill (4)

where the weight $\theta \in [0, 1]$ is the share of consumers that are new to the market: $\theta = \frac{\sum_j \frac{\partial q_j}{\partial p_i}}{\frac{\partial q_i}{\partial p_i}}$, the sum of consumers gained by the market divided by the consumers gained by firm $i$.$^{13}$

Weyl and Fabinger (2013) show that under a range of different micro-foundations for the nature of competition, the first order conditions for price are given by

$$p_i + b - mc_i(q) = \theta \mu(p)$$  \hfill (5)

where $\theta$, defined as above, can be thought of as conduct parameter (Bresnahan, 1989) that indexes the degree of market power and $\mu(p)$ is an absolute markup function that is equal to $p$ times the inverse elasticity of aggregate demand: $\mu(p) \equiv -\frac{q}{q'} = \frac{p}{\epsilon_p}$, where $\epsilon_p$ is the aggregate elasticity of demand.$^{14}$

Substituting for $mc_i$ using Equation 4 and rearranging allows us to write the first order condition in terms of industry marginal and average costs:

$$p_i + b = \theta [mc(q) + \mu(p)] + (1 - \theta) ac(q)$$  \hfill (6)

Perfect competition is given by $\theta = 0$, and simplifies the first order condition to the “price equals average costs” condition in Einav, Finkelstein and Cullen (2010). Monopoly is given by $\theta = 1$, and simplifies the equation to the standard Lerner Index for optimal pricing $\frac{p + b - mc(q)}{p} = \frac{1}{\epsilon_p}$.

The specification also captures standard models of imperfect competition. Cournot competition is given by $\theta = 1/n$, where $n$ is the number of firms. Differentiated product Nash-in-prices compet-

$^{13}$Equivalently, the weight can be defined as $\theta = 1 - A$, where $A$ is the aggregate division ratio: $A = -\frac{\sum_{j \neq i} \frac{\partial q_j}{\partial p_i}}{\frac{\partial q_i}{\partial p_i}}$, the sum of consumers lost by firms $j \neq i$ divided by the consumers gained by firm $i$.

$^{14}$The second order condition for $p$ is $\theta \mu' < 1$. We assume that at the optimal price this condition is satisfied.
Preliminary and Incomplete

tition can be modeled by setting \( \theta \equiv \sum_j \frac{\partial q_j}{\partial p_i} \frac{\partial q_j}{\partial p_i} \), the share of consumers that are new to the market. See Weyl and Fabinger (2013) and Mahoney and Weyl (2013) for more on the micro-foundations of this specification.

5.3 Pass-Through

We are interested in how much of an increase in payments is passed through into lower health insurance premiums. For a small change in payments, pass-through is defined as negative one times the total derivative of premiums with respect to the capitation payment \( b \): \( \rho \equiv -\frac{dp}{db} \). We will say there is full pass-through when \( \rho = 1 \) and no pass-through when \( \rho = 0 \).

Fully differentiating Equation 6 with respect to \( b \) and rearranging yields the pass-through equation:

\[
\rho = \frac{1 - \theta'(b)(\mu + mc(q) - ac(q))}{1 - \theta mc'(q)q'(p) - \theta \mu'(p) - (1 - \theta)ac'(q)q'(p)}
\]  

Consider the case when a change in payments does not impact the market structure (\( \theta'(b) = 0 \)). The intuition for this equation can be understood by considering the extremes. Under perfect competition (\( \theta = 0 \)) the pass-through formula simplifies to \( \rho = \frac{1}{1-ac'(q)q'(p)} \). Since demand is downward sloping (\( q'(p) < 0 \)), pass-through is less than 1 when the average cost curve is downward sloping (\( ac'(q) < 0 \)) and greater than 1 when the average cost curve slopes upward (\( ac'(q) > 0 \)).

When there is no selection (\( mc'(q) = ac'(q) = 0 \)) the pass-through formula simplifies to \( \rho = \frac{1}{1-\theta \mu'(p)} \). In this setting, the comparative statistics are slightly more subtle. Pass-through is decreasing in market power (higher \( \theta \)) for many standard parameterizations of demand. Technically, this occurs when \( \mu' < 0 \) or when log demand is concave, since \( (\log q)'' = \mu'/\mu^2 < 0 \iff \mu' < 0 \). When \( \mu'(p) > 0 \), the pass-through rate can be greater than 1 and be increasing in market power.  

6 Costs and Competition

Coming soon

\(^{15}\)Fabinger and Weyl (2013) prove that \( \mu' < 0 \) if demand is linear or if it is based on an underlying willingness-to-pay distribution that is normal, logistic, Type I Extreme Value (logit), Laplace, Type III Extreme Value, or Weibull or Gamma with shape parameter \( \alpha > 1 \). They show that \( \mu' > 0 \) if demand is based on a willingness-to-pay distribution that is Pareto (constant elasticity), Type II Extreme Value, or Weibull or Gamma with shape parameter \( \alpha < 1 \). They show that \( \mu \) switches from \( \mu' < 0 \) to \( \mu' > 0 \) for a log-normal distribution of willingness-to-pay.
7 Conclusion

Coming soon.
References


### Table 1: Summary Statistics

<table>
<thead>
<tr>
<th></th>
<th>Panel A: All Counties</th>
<th>Panel B: Payment Floor Counties</th>
</tr>
</thead>
<tbody>
<tr>
<td>Base Rate ($ per month)</td>
<td>540.68</td>
<td>86.65</td>
</tr>
<tr>
<td>At Least 1 Plan</td>
<td>70.8%</td>
<td>45.5%</td>
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<tr>
<td>Number of Plans</td>
<td>2.05</td>
<td>1.98</td>
</tr>
<tr>
<td>County Averages</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Premiums ($ per month)</td>
<td>32.26</td>
<td>32.63</td>
</tr>
<tr>
<td>Benefits</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Copay ($ per visit)</td>
<td>9.89</td>
<td>3.89</td>
</tr>
<tr>
<td>Specialist Copay ($ per visit)</td>
<td>14.22</td>
<td>6.53</td>
</tr>
<tr>
<td>Drugs</td>
<td>68.7%</td>
<td>31.3%</td>
</tr>
<tr>
<td>Dental</td>
<td>30.2%</td>
<td>30.5%</td>
</tr>
<tr>
<td>Hearing</td>
<td>37.3%</td>
<td>36.5%</td>
</tr>
<tr>
<td>Eyewear</td>
<td>66.1%</td>
<td>34.3%</td>
</tr>
</tbody>
</table>

**Note:** Table shows summary statistics for all counties and for counties that were affected by the payment floors. Values are calculated using the pooled 1996 to 2003 data, with the exception of benefits for which data is only available for 2000 to 2003. Premiums and benefits values are calculated using the sample of counties with at least one plan in all years.
Table 2: Effect of BIPA on County Base Rates

<table>
<thead>
<tr>
<th></th>
<th>Mean</th>
<th>Std. Dev.</th>
<th>25th</th>
<th>50th</th>
<th>75th</th>
</tr>
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<tbody>
<tr>
<td>Non Floor County (N = 886)</td>
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<td></td>
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<td></td>
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<tr>
<td>Δ Base Rate</td>
<td>5.24</td>
<td>0.57</td>
<td>4.80</td>
<td>5.11</td>
<td>5.50</td>
</tr>
<tr>
<td>% Change in Base Rate</td>
<td>1.0%</td>
<td>0.0%</td>
<td>1.0%</td>
<td>1.0%</td>
<td>1.0%</td>
</tr>
<tr>
<td>Rural Floor County (N = 1694)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Δ Base Rate</td>
<td>47.35</td>
<td>17.01</td>
<td>35.32</td>
<td>59.99</td>
<td>59.99</td>
</tr>
<tr>
<td>% Change in Base Rate</td>
<td>11.2%</td>
<td>4.3%</td>
<td>8.0%</td>
<td>14.5%</td>
<td>14.5%</td>
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<tr>
<td>Urban Floor County (N = 563)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Δ Base Rate</td>
<td>56.43</td>
<td>30.42</td>
<td>31.80</td>
<td>55.86</td>
<td>79.24</td>
</tr>
<tr>
<td>% Change in Base Rate</td>
<td>12.8%</td>
<td>7.4%</td>
<td>6.7%</td>
<td>12.4%</td>
<td>17.8%</td>
</tr>
</tbody>
</table>

Note: Table shows the effect of BIPA on base rates for non-floor counties, and for counties that were affected by the rural and urban floors. The Δ Base Rate rows show the difference between the March 2001 base rate and the February 2001 base rate in dollars per beneficiary per month. The % Change in Base Rate rows show this difference as a percent of the February 2001 base rate. See text for additional information on data construction.
Table 3: Base Rates: Impact of $1 Distance from the Floor

<table>
<thead>
<tr>
<th></th>
<th>(1)</th>
<th>(2)</th>
<th>(3)</th>
</tr>
</thead>
<tbody>
<tr>
<td>( \Delta b \times 2001 )</td>
<td>0.885</td>
<td>0.888</td>
<td>0.885</td>
</tr>
<tr>
<td></td>
<td>(0.006)</td>
<td>(0.007)</td>
<td>(0.006)</td>
</tr>
<tr>
<td>( \Delta b \times 2002 )</td>
<td>1.100</td>
<td>1.106</td>
<td>1.099</td>
</tr>
<tr>
<td></td>
<td>(0.007)</td>
<td>(0.010)</td>
<td>(0.006)</td>
</tr>
<tr>
<td>( \Delta b \times 2003 )</td>
<td>1.079</td>
<td>1.088</td>
<td>1.077</td>
</tr>
<tr>
<td></td>
<td>(0.008)</td>
<td>(0.014)</td>
<td>(0.008)</td>
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</table>

Main Effects
- County FE: X    X    X
- Year FE: X      X    X

Additional Controls
- Base-Rate X Year Trend: X
- Urban X Year Trend: X

Pre-BIPA Mean Base Rate: 522.62 522.62 522.62
R-Squared: 0.996 0.996 0.996

Note: Table shows coefficients on the distance-to-the-floor variables (\( \Delta b_j \times \) year interactions) from difference-in-difference regressions with the monthly base rate as the dependent variable. The unit of observation is the county × year, and observations are weighted by the number of beneficiaries in the county. Controls include year and county fixed effects and additional controls as noted in the table. Year 2000, which is the year prior to BIPA implementation, is the omitted category. Robust standard errors clustered at the county level (N = 397) are reported in parentheses.
Table 4: Premium Pass-Through: Impact of $1 Increase in Monthly Payments

<table>
<thead>
<tr>
<th></th>
<th>Δb X 2001</th>
<th></th>
<th></th>
<th></th>
<th></th>
<th></th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td></td>
<td>(1)</td>
<td>(2)</td>
<td>(3)</td>
<td>(4)</td>
<td>(5)</td>
<td>(6)</td>
<td>(7)</td>
<td>(8)</td>
</tr>
<tr>
<td>Δb X 2001</td>
<td>-0.097</td>
<td>-0.137</td>
<td>-0.097</td>
<td>0.011</td>
<td>-0.035</td>
<td>0.009</td>
<td>-0.203</td>
<td>-0.179</td>
<td>-0.203</td>
</tr>
<tr>
<td></td>
<td>(0.043)</td>
<td>(0.044)</td>
<td>(0.043)</td>
<td>(0.049)</td>
<td>(0.050)</td>
<td>(0.049)</td>
<td>(0.073)</td>
<td>(0.083)</td>
<td>(0.073)</td>
</tr>
<tr>
<td>Δb X 2002</td>
<td>-0.179</td>
<td>-0.260</td>
<td>-0.181</td>
<td>-0.044</td>
<td>-0.136</td>
<td>-0.047</td>
<td>-0.333</td>
<td>-0.285</td>
<td>-0.332</td>
</tr>
<tr>
<td></td>
<td>(0.051)</td>
<td>(0.056)</td>
<td>(0.051)</td>
<td>(0.068)</td>
<td>(0.070)</td>
<td>(0.068)</td>
<td>(0.081)</td>
<td>(0.102)</td>
<td>(0.081)</td>
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<tr>
<td>Δb X 2003</td>
<td>-0.171</td>
<td>-0.292</td>
<td>-0.173</td>
<td>0.005</td>
<td>-0.134</td>
<td>0.000</td>
<td>-0.348</td>
<td>-0.277</td>
<td>-0.348</td>
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<tr>
<td></td>
<td>(0.070)</td>
<td>(0.072)</td>
<td>(0.069)</td>
<td>(0.087)</td>
<td>(0.090)</td>
<td>(0.087)</td>
<td>(0.092)</td>
<td>(0.120)</td>
<td>(0.092)</td>
</tr>
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</table>

Main Effects
- County FE: X X X X X X X X X
- Year FE: X X X X X X X X X

Additional Controls
- Base-Rate X Year Trend: X X X X X X X X X
- Urban X Year Trend: X X X X X X X X X

Pre-BIPA Mean Dependent Variable
- Mean Premium: 11.81 11.81 11.81 2.96 2.96 2.96 31.58 31.58 31.58
- Minimum Premium: 0.69 0.70 0.69 0.56 0.57 0.56 0.60 0.60 0.60
- Maximum Premium: 0.69 0.70 0.69 0.56 0.57 0.56 0.60 0.60 0.60

Note: Table shows coefficients on the distance-to-the-floor variables (Δbj × year interactions) from difference-in-difference regressions. Note that the first stage results displayed in Table 3 indicate that a $1 change in the distance to the floor roughly translates to a $1 change in the monthly payments. The dependent variables are mean monthly premiums, median monthly premiums, minimum monthly premiums, and maximum monthly premiums. The unit of observation is the county × year, and observations are weighted by the number of beneficiaries in the county. Controls include year and county fixed effects and additional controls as noted in the table. Year 2000, which is the year prior to BIPA implementation, is the omitted category. Robust standard errors clustered at the county level (N = 397) are reported in parentheses.
### Table 5: Benefits Generosity: Impact of a $50 Increase in Monthly Payments

<table>
<thead>
<tr>
<th>Dependent Variable</th>
<th>Pre-BIPA Mean</th>
<th>Implied Elasticity†</th>
<th>R-Squared</th>
</tr>
</thead>
<tbody>
<tr>
<td>Physician Copay ($)</td>
<td>7.37</td>
<td>-4.53</td>
<td>0.74</td>
</tr>
<tr>
<td>Specialist Copay ($)</td>
<td>7.37</td>
<td>-1.90</td>
<td>0.72</td>
</tr>
<tr>
<td>Drug Coverage (%)</td>
<td>10.12</td>
<td>-4.80</td>
<td>0.75</td>
</tr>
<tr>
<td>Hearing Coverage (%)</td>
<td>10.12</td>
<td>-3.36</td>
<td>0.75</td>
</tr>
<tr>
<td>Vision Coverage (%)</td>
<td>74.47</td>
<td>0.40</td>
<td>0.75</td>
</tr>
<tr>
<td>Dental Coverage (%)</td>
<td>42.56</td>
<td>0.40</td>
<td>0.75</td>
</tr>
</tbody>
</table>

Note: Table shows scaled coefficients on the distance-to-the-floor variables ($\Delta b_j \times$ year interactions) from difference-in-difference regressions. Note that the first stage results displayed in Table 3 indicate that a $1$ change in the distance to the floor roughly translates to a $1$ change in the monthly payments. The dependent variables are physician copays in dollars, specialist copays in dollars, and indicators for drug, dental, hearing aid and vision coverage. The unit of observation is the county × year, and observations are weighted by the number of beneficiaries in the county. Controls include year and county fixed effects and additional controls as noted in the table. Year 2000, which is the year prior to BIPA implementation, is the omitted category. Robust standard errors clustered at the county level ($N = 397$) are reported in parentheses.

*Impact of $50$ increase.

†Elasticity calculated as percent change in benefit generosity / percent change in monthly payments using parameter estimate for 2003.
Table 6: Actuarial Value of Benefits: Impact of $1 Increase in Monthly Payments

<table>
<thead>
<tr>
<th></th>
<th>(1)</th>
<th>(2)</th>
<th>(3)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Δb X 2001</td>
<td>0.049</td>
<td>0.052</td>
<td>0.048</td>
</tr>
<tr>
<td></td>
<td>(0.026)</td>
<td>(0.029)</td>
<td>(0.026)</td>
</tr>
<tr>
<td>Δb X 2002</td>
<td>0.068</td>
<td>0.075</td>
<td>0.066</td>
</tr>
<tr>
<td></td>
<td>(0.031)</td>
<td>(0.038)</td>
<td>(0.031)</td>
</tr>
<tr>
<td>Δb X 2003</td>
<td>0.088</td>
<td>0.098</td>
<td>0.085</td>
</tr>
<tr>
<td></td>
<td>(0.033)</td>
<td>(0.045)</td>
<td>(0.033)</td>
</tr>
<tr>
<td>Main Effects</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>County FE</td>
<td>X</td>
<td>X</td>
<td>X</td>
</tr>
<tr>
<td>Year FE</td>
<td>X</td>
<td>X</td>
<td>X</td>
</tr>
<tr>
<td>Additional Controls</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Base-Rate X Year Trend</td>
<td></td>
<td>X</td>
<td></td>
</tr>
<tr>
<td>Urban X Year Trend</td>
<td></td>
<td></td>
<td>X</td>
</tr>
<tr>
<td>R-Squared</td>
<td>0.75</td>
<td>0.75</td>
<td>0.75</td>
</tr>
</tbody>
</table>

**Note:** Table shows coefficients on the distance-to-the-floor variables (Δ$b_j \times$ year interactions) from difference-in-difference regressions. Note that the first stage results displayed in Table 3 indicate that a $1 change in the distance to the floor roughly translates to a $1 change in the monthly payments. The dependent variable is the monthly actuarial value of benefits, which is constructed based on observed plan benefits in our main analysis dataset and utilization and cost data from the 2000 Medical Expenditure Panel Survey. See text for full details. The unit of observation is the county × year, and observations are weighted by the number of beneficiaries in the county. Controls include year and county fixed effects and additional controls as noted in the table. Year 2000, which is the year prior to BIPA implementation, is the omitted category. Robust standard errors clustered at the county level (N = 397) are reported in parentheses.
Table 7: Plan Availability: Impact of $50 Increase in Monthly Payments

<table>
<thead>
<tr>
<th></th>
<th>At Least One Plan (%)</th>
<th>Number of Plans</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(1)</td>
<td>(2)</td>
</tr>
<tr>
<td>Δb X 2001*</td>
<td>-1.738</td>
<td>-0.722</td>
</tr>
<tr>
<td></td>
<td>(1.417)</td>
<td>(1.508)</td>
</tr>
<tr>
<td>Δb X 2002*</td>
<td>-0.362</td>
<td>1.670</td>
</tr>
<tr>
<td></td>
<td>(1.528)</td>
<td>(1.722)</td>
</tr>
<tr>
<td>Δb X 2003*</td>
<td>1.599</td>
<td>4.648</td>
</tr>
<tr>
<td></td>
<td>(1.793)</td>
<td>(2.073)</td>
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</table>

Main Effects
- County FE
- Year FE

Additional Controls
- Base-Rate X Year Trend
- Base-Rate X Year FE
- Second-Order Base-Rate X Year FE

Pre-BIPA Mean Enrollment | 74.44 | 74.44 | 74.44 | 2.46   | 2.46   | 2.46   |
Implied Elasticity†      | 0.21  | 0.62  | 0.25  | 1.58   | 1.65   | 1.42   |
R-Squared                | 0.81  | 0.81  | 0.81  | 0.88   | 0.88   | 0.88   |

Note: Table shows scaled coefficients on the distance-to-the-floor variables (Δb_j × year interactions) from difference-in-difference regressions. Note that the first stage results displayed in Table 3 indicate that a $1 change in the distance to the floor roughly translates to a $1 change in the monthly payments. The dependent variables are the presence of any plan and the count of plans. The unit of observation is the county × year, and observations are weighted by the number of beneficiaries in the county. Controls include year and county fixed effects and additional controls as noted in the table. Year 2000, which is the year prior to BIPA implementation, is the omitted category. Robust standard errors clustered at the county level (N = 3,143) are reported in parentheses.

*Impact of $50 increase.
†Elasticity calculated as percent change in plan availability / percent change in monthly payments using parameter estimate for 2003.
Figure 1: Payment Floors: Pre- and Post-BIPA Monthly Base Rates

Note: Figure shows county base rates before (x-axis) and after (y-axis) the implementation of the BIPA urban and rural rate floors in March of 2001. The left plot shows the entire sample. The right plot restricts the sample to focus on counties close to the rate floors. The dashed line indicates the uniform 1% increase that was applied to all counties and traces the counterfactual payment rule in absence of the floors. All values are denominated in dollars per beneficiary per month.
**Figure 2: Base Rates: Impact of $1 Distance from the Floor**

![Graph showing base rates over years with notes explaining the coefficients and methodology]

**Note:** Figure shows coefficients on the distance-to-the-floor variables ($\Delta b_j \times$ year interactions) from difference-in-difference regressions with the monthly base rate as the dependent variable. The unit of observation is the county $\times$ year, and observations are weighted by the number of beneficiaries in the county. Controls include year and county fixed effects and an interaction of the pre-BIPA county base rate with a linear time trend. The capped vertical bars show 95% confidence intervals, calculated using standard errors clustered at the county level. Year 2000, which is the year prior to BIPA implementation, is the omitted category and denoted with a vertical dashed lined line. Horizontal dashed lines are plotted at the reference values of 0 and 1.
Figure 3: Premium Pass-Through: Impact of $1 Increase in Monthly Payments

(a) Mean

(b) Median

(c) Minimum

(d) Maximum

Note: Figure shows coefficients on the distance-to-the-floor variables (\(\Delta b_j \times \text{year interactions}\)) from difference-in-difference regressions. Note that the first stage results displayed in Figure 2 indicate that a $1 change in the distance to the floor roughly translates to a $1 change in the monthly payments. The dependent variables are mean monthly premiums (Panel a), median monthly premiums (Panel b), minimum monthly premiums (Panel c) and maximum monthly premiums (Panel d). The unit of observation is the county \(\times\) year, and observations are weighted by the number of beneficiaries in the county. Controls are as described in Figure 2. The capped vertical bars show 95% confidence intervals, calculated using standard errors clustered at the county level. Horizontal dashed lines are plotted at the reference values of 0 and -1, where -1 corresponds to 100% pass-through.
Figure 4: Benefits Generosity: Impact of a $50 Increase in Monthly Payments

(a) Physician Copay

(b) Specialist Copay

(c) Drug Coverage

(d) Dental Coverage

(e) Hearing Aid

(f) Vision

Note: Figure shows scaled coefficients on the distance-to-the-floor variables ($\Delta b_j \times$ year interactions) from difference-in-difference regressions. Note that the first stage results displayed in Figure 2 indicate that a $1$ change in the distance to the floor roughly translates to a $1$ change in the monthly payments. The dependent variables are physician copays in dollars (Panel a), specialist copays in dollars (Panel b), and indicators for drug (Panel c), dental (Panel d), hearing aid (Panel e) and vision (Panel f) coverage. The unit of observation is the county $\times$ year, and observations are weighted by the number of beneficiaries in the county. Controls are as described in Figure 2. In Panels a and b, the vertical axes measure the effect on copays in dollars of a $50$ difference in monthly reimbursement. In Panels c through f, the vertical axes measure the effect on the probability that a plan offers each fringe benefit, again for a $50$ difference in monthly reimbursement. The capped vertical bars show 95% confidence intervals, calculated using standard errors clustered at the county level. Year 2000, which is the year prior to BIPA implementation, is the omitted category. Horizontal dashed lines are plotted at 0.
Figure 5: Actuarial Value of Benefits: Impact of $1 Increase in Monthly Payments

Note: Figure shows coefficients on the distance-to-the-floor variables ($\Delta b_j \times$ year interactions) from difference-in-difference regressions. Note that the first stage results displayed in Figure 2 indicate that a $1$ change in the distance to the floor roughly translates to a $1$ change in the monthly payments. The dependent variable is the actuarial value of benefits, which is constructed based on observed plan benefits in our main analysis dataset and utilization and cost data from the 2000 Medical Expenditure Panel Survey. See text for full details. The unit of observation is the county $\times$ year, and observations are weighted by the number of beneficiaries in the county. Controls are as described in Figure 2. The capped vertical bars show 95% confidence intervals, calculated using standard errors clustered at the county level. The horizontal dashed line is plotted at 0.
**Figure 6:** Plan Availability: Impact of $50 Increase in Monthly Payments

(a) At Least One Plan

(b) Number of Plans

Note: Figure shows scaled coefficients on the distance-to-the-floor variables ($\Delta b_j \times$ year interactions) from difference-in-difference regressions. Note that the first stage results displayed in Figure 2 indicate that a $1$ change in the distance to the floor roughly translates to a $1$ change in the monthly payments. The dependent variables are the presence of any plan (Panel a) and the count of plans (Panel b). The unit of observation is the county $\times$ year, and observations are weighted by the number of beneficiaries in the county. Controls are as described in Figure 2. The capped vertical bars show 95% confidence intervals, calculated using standard errors clustered at the county level. The horizontal dashed lines are plotted at at the sample means, which are added to the coefficients.
Figure 7: Incomplete Pass-Through

(a) Advantageous Selection

(b) Market Power

Note: Figure shows the pass-through of an increase in capitation payments depicted by a decrease in marginal costs. Panel A examines pass-through when there is no selection and when there is advantageous selection in a perfectly competitive market. Panel B examines pass-through where there is perfectly competition and where there is a monopolist in a market with no selection.