Health Insurance and the Consumer Bankruptcy Decision: Evidence from Expansions of Medicaid*

Tal Gross
Columbia University
Mailman School of Public Health

Matthew J. Notowidigdo
The University of Chicago
Booth School of Business

December 2010

Abstract

Anecdotal evidence and several observational studies suggest that out-of-pocket medical costs are pivotal in a large fraction of consumer bankruptcy decisions. In this paper, we assess the contribution of medical costs to household bankruptcy risk by exploiting plausibly exogenous variation in publicly provided health insurance. Using cross-state variation in Medicaid expansions from 1992 through 2004, we find that a 10 percentage point increase in Medicaid eligibility reduces personal bankruptcies by 8 percent, with no evidence that business bankruptcies are similarly affected. We interpret our findings with a model in which health insurance imperfectly substitutes for other forms of financial protection, and we use the model to present simple calibration results which illustrate how our reduced-form parameter estimate affects the optimal level of health insurance benefits. We conclude with calculations which suggest that out-of-pocket medical costs are pivotal in roughly 26 percent of personal bankruptcies among low-income households.

Keywords: Health Insurance, Bankruptcy, Medicaid.
JEL Classification: G33, G38, K35, G22.

*We thank Daron Acemoglu, Josh Angrist, David Autor, Amy Finkelstein, Jonathan Gruber, and seminar participants at MIT, Columbia, and The Federal Reserve Bank of Chicago. We are extremely grateful to Kosali Simon for guidance with all simulated eligibility calculations; to Damon Seils and Kevin Schulman for assistance with the AHA uncompensated care data; and to Richard Hynes for information regarding state exemptions. Notowidigdo thanks the National Institute of Aging (NIA grant number T32-AG000186) for financial support.

1 Introduction

Bankruptcy protection is a legal procedure designed to forgive debtors their debt. It was once undertaken by few debtors, but has become common over the past few decades (Zywicki, 2005). In the 1990s, the number of personal bankruptcies in the United States rose by more than 78 percent (see figure 1). By the end of the decade, more than 1 percent of American households were declaring bankruptcy in any given year. Stavins (2000) estimates that 8.5 percent of American households have ever filed for bankruptcy.

This increase in bankruptcies has motivated research on factors that induce households to declare bankruptcy. One such factor is the burden of out-of-pocket medical costs. Several researchers have argued that a large fraction of consumer bankruptcies are driven by the high cost of health care. This conjecture has been widely publicized and has also motivated legislation to prevent "medical bankruptcies." For instance, a bill proposed in Congress, "The Medical Bankruptcy Fairness Act of 2008," would have lowered penalties on debtors forced to declare bankruptcy because of medical bills.

Currently, there exists little credible evidence regarding the relative importance of medical costs in the decision to declare bankruptcy. The few studies that have pursued such evidence rely primarily on interviews with individuals who have recently filed for bankruptcy. Such interviews are unlikely to isolate whether bankruptcy filers who experienced high medical costs would have still declared bankruptcy in the absence of any medical costs.

In this paper, we use plausibly exogenous variation in publicly provided health insurance to examine the effect of medical costs on bankruptcy risk. Specifically, we exploit state-level expansions in Medicaid eligibility during the 1990s.[1] In our

[1] The variation we exploit stems from both Medicaid expansions and the State Children's Health

preferred specification, we find that a 10 percentage-point increase in eligibility for Medicaid reduces personal bankruptcies by 8 percent.

We test the robustness of these main findings in several ways. First, we document the results of a simple falsification test: business bankruptcies are not similarly affected by Medicaid expansions. Second, we present the results of a variety of alternative specifications which control for other determinants of consumer bankruptcies. Finally, we construct a database of bankruptcies based on the dockets of bankruptcy courts, and we compile counts of bankruptcies by zip code. This database allows us to test whether Medicaid expansions—which primarily affected households with children—affected certain zip codes rather than others. We find that Medicaid expansions disproportionately reduced bankruptcies in zip codes in which children are a large share of the population as well as zip codes with a large share of low-income households. In general, all of these exercises confirm our main findings.

The empirical results suggest a robust interaction between Medicaid and the consumer bankruptcy system. To explore the welfare implications of this interaction, we construct a simple theoretical model in which health insurance is an imperfect substitute for other forms of financial protection. We calibrate this model and find that the optimal health insurance benefit rate is between 14 and 24 percent lower than would be suggested by a model which ignores the generosity of the bankruptcy system and the imperfect substitutability between health insurance and consumer bankruptcy. While the calibrations are extremely stylized, they qualitatively demonstrate the likely substantive importance of the interaction between bankruptcy and Medicaid.

The remainder of the paper proceeds as follows. The subsequent section discusses

Insurance Program (SCHIP). For simplicity, we refer to both Medicaid and SCHIP simply as "Medicaid," even though SCHIP provides health insurance to children through programs that are—in some states—distinct from Medicaid.

the state of research on personal bankruptcy. Section 3 describes Medicaid expansions, the data, and our empirical strategy. Section 4 presents our main results and robustness tests. Section 5 explores patterns in households' exposure to financial risk that may drive the main findings. Section 6 develops a model of the interaction between bankruptcy and Medicaid and presents calibration results that utilize our empirical estimates. Section 7 estimates the share of bankruptcies among low-income households that are driven by medical costs. Section 8 concludes.

2 Previous Research on the Determinants of Consumer Bankruptcy

Many studies have explored the determinants of consumer bankruptcy. The research generally falls into two categories. One strand of research emphasizes the strategic nature of the household bankruptcy decision. The studies find that households are forward-looking and optimally choose whether or not to file for bankruptcy based on the financial advantage of doing so. Households take the generosity of the bankruptcy system into account in making savings and investment decisions. As a result, the bankruptcy system leads to ex-ante moral hazard.

For example, several studies document that households respond to financial incentives when deciding whether to declare bankruptcy. Fay et al. (2002) study a sample of respondents to the Panel Study of Income Dynamics who have declared bankruptcy. The authors find that households are more likely to declare bankruptcy when the financial benefits of doing so outweigh the costs. Researchers have also documented that stigma and the availability of credit may be critical factors. Both Zywicki (2005) and Gross and Souleles (2001) conclude that the stigma of declaring bankruptcy has diminished over time. Similarly, Livshits et al. (2007) estimate a structural model of household financial decisions. The authors conclude that the rise in personal bankruptcy has been driven mainly by the increasing availability of consumer credit and a decline in the social cost of filing for bankruptcy, rather than by uncertainty or medical shocks.

A second strand of research quantifies the role of adverse, potentially unforeseen shocks that may lead to consumer bankruptcies. For instance, Keys (2009) studies the relationship between unemployment and bankruptcy. Additionally, a study by Himmelstein et al. (2005) estimates that medical costs are pivotal in more than half of all consumer bankruptcies. In interviews with bankruptcy filers, the authors find that 54 percent of respondents cite "any medical cause" when asked what led them to declare bankruptcy. The finding confirms other qualitative studies that point to adverse events as the primary driver of personal bankruptcy (Sullivan et al. 1989). One concern with this study, however, is that the authors define medical costs broadly. They include the birth or death of a family member, alcoholism, drug addiction, and uncontrolled gambling as "any medical cause." Dranove and Millenson (2006) re-analyze the same survey data using a narrower definition of medical causes and attribute far fewer bankruptcies to medical costs. They estimate that 17 percent of bankruptcies are due to medical causes, most of which involve low-income households. Recent follow-up studies suffer from similar drawbacks. Himmelstein et al. (2009) interview a sample of bankruptcy filers, 29 percent of whom state that medical costs were a reason for filing. The authors then add to this estimate respondents who did *not* state that medical costs were a factor in their bankruptcy, but who did describe substantial medical costs. In this way, the authors calculate that 62 percent of bankruptcies can be classified as "medical," even though more than half of the relevant respondents did not list medical costs as a primary cause of their decision to file for bankruptcy.

More broadly, a concern with both strands of research is that the studies do not employ quasi-experimental variation in the determinants of bankruptcy, which makes it difficult to credibly estimate the causal effect of interest. To our knowledge, this paper is the first to document the relative importance of medical costs in the bankruptcy decision using plausibly exogenous variation in public health insurance eligibility.

3 Empirical Strategy and Data

This section briefly describes the Medicaid expansions we study, the data we use, and our empirical framework.

3.1 Background on Medicaid Expansions

In the mid-1990s, states expanded Medicaid eligibility to cover all young children living in families with incomes below 133 percent of the federal poverty line, and in certain states, their parents. In 1997, the Medicaid program was augmented further with the introduction of the State Children's Health Insurance Program (SCHIP), which expanded Medicaid eligibility for children and pregnant women. Many states also went beyond the minimum federally required extended eligibility. New Jersey, for example, offered Medicaid to children whose families earned less than 350 percent of the federal poverty level (see Gruber 2000 and Gruber and Simon 2008 for more details on the Medicaid program). Many states also expanded eligibility for parents in conjunction with their SCHIP expansions.[2] Crucially for our estimation strategy, states expanded Medicaid eligibility at different times, and states chose to expand eligibility by different amounts during this time period. Figure 2 plots the increase in

[2] There exists anecdotal evidence that most Medicaid expansions during this time period involved retrospective eligibility. For instance, hospitals may apply for Medicaid on behalf of eligible but uninsured patients. To our knowledge, little research has documented the extent of this practice. We believe that retrospective eligibility may be an important part of how Medicaid affects household finances; households that are eligible but do not enroll may still be covered after illness or injury.

Medicaid eligibility from 1992 through 2004. Overall, roughly 15 percent of all U.S. households became eligible for Medicaid during this time.

3.2 Data

Our investigation into bankruptcy and public health insurance requires accurate measures of both variables. For the latter, we construct measures of public insurance eligibility using the 1992–2004 March Current Population Survey (CPS).[3] First, we calculate whether each surveyed household is eligible for Medicaid in its state of residence and year, given the household's income, number of children, and the gender of the head of household. We also perform a similar procedure to calculate simulated eligibility. Specifically, we take a 20 percent national sample from the 1996 CPS and calculate the share of this fixed population that would be eligible for Medicaid in each state and year.[4]

For bankruptcy data, we rely on the publicly-available census of consumer and business bankruptcies. This census is published annually by the Administrative Office of the U.S. Courts. The data include bankruptcy totals for each bankruptcy district. There are 94 bankruptcy districts, with one to four districts per state. We exclude bankruptcy districts in US territories and compile counts of bankruptcies by state and year.[5]

Table 1 presents some descriptive information on our sample. In 1992, states processed an average of 17,615 bankruptcies. Over the next decade, bankruptcy counts nearly doubled. The table presents descriptive statistics for the five states with the smallest expansions of Medicaid and the five states with the largest expansions of

[3]The consumer bankruptcy system changed substantially after the 2005 Bankruptcy Abuse Prevention and Consumer Protection Act. We limit our sample to bankruptcies in 2004 and earlier in order to avoid that change.

[4]We are extremely grateful to Kosali Simon for computer code that constructs these two variables.

[5]The excluded bankruptcy districts are those in the Virgin Islands, Puerto Rico, Northern Mariana Islands, and Guam.

Medicaid during our sample period.[6] For the "small expansion states," bankruptcy counts more than doubled, growing from an average of 14,336 in 1992 to an average of 30,872 in 2004. For the "large expansion states," however, bankruptcy counts grew by a smaller amount in both absolute and proportional terms, from 46,320 in 1992 to 51,585 in 2004.

3.3 Empirical Strategy

Figure 3 summarizes our approach and main results. The figure plots the difference in the logarithm of consumer bankruptcies between 1992 and 2004 for each state against the change in simulated Medicaid eligibility over that time period. The figure demonstrates that states with larger expansions in Medicaid eligibility experienced a smaller increase in bankruptcies over the 1990s. Our main empirical strategy is similar. We compare the change in the consumer bankruptcy rate across states with varying changes in Medicaid generosity.

We model the relationship between Medicaid eligibility and consumer bankruptcy as:

$$\log(c_{st}) = \alpha_s + \alpha_t + \beta M_{st} + \varepsilon_{st}, \tag{1}$$

where c_{st} denotes the number of consumer bankruptcies in state s and year t, α_s are state fixed effects, and α_t are year fixed effects. The variable M_{st} denotes the fraction of the population eligible for Medicaid, and ε_{st} represents unobserved state-year shocks that affect the number of consumer bankruptcies.

Estimating equation (1) with ordinary least squares (OLS) would likely lead to biased estimates of β. Unobserved, adverse economic shocks will lead to more consumer bankruptcies and more households qualifying for Medicaid. Moreover, measurement

[6]The large expansion states are California, Missouri, Florida, Minnesota, and the District of Columbia. The small expansion states are South Carolina, Texas, North Carolina, North Dakota, and West Virginia.

error in eligibility may attenuate estimates of β. To overcome both issues, we use simulated Medicaid eligibility as an instrumental variable (IV) for actual Medicaid eligibility.[7] The simulated instruments isolate variation in Medicaid eligibility that is driven by changes in the program parameters rather than by changes in the economic environment. Simulated Medicaid eligibility is highly correlated with actual Medicaid eligibility (the t-statistic for simulated eligibility from our first stage regression is 12.78, see appendix table 1), but is assumed not to be correlated with adverse economic shocks and other unobserved or omitted determinants of consumer bankruptcies.[8] The key identifying assumption is that absent changes in simulated Medicaid eligibility, state bankruptcy rates would have evolved similarly over time. We begin by estimating equation (1) using instrumental variables under this assumption. We then investigate the validity of this assumption with several robustness tests.

4 The Aggregate Effect of Medicaid Expansions on Bankruptcies

This section presents estimates of the effect of Medicaid expansions on aggregate bankruptcy counts. Table 2 presents our main results. The first column shows the OLS relationship between Medicaid eligibility and state consumer bankruptcy filings. The estimated negative relationship is not statistically significant at conventional levels.[9] Column 2 reports the IV estimates. The magnitude of the IV estimate

[7] Simulated instrumental variables for Medicaid eligibility were introduced by Currie and Gruber (1996). Simulated instruments for Medicaid have also been used by Gruber and Yelowitz (1999), Cutler and Gruber (1996), DeLeire et al. (2007), and Gruber and Simon (2008).

[8] As shown in the appendix table 2, we find no evidence that lagged consumer bankruptcies or other economic variables predict changes in the simulated eligibility variable. This indicates that Medicaid expansions were not driven by past trends in consumer bankruptcies.

[9] In all tables we report robust standard errors that allow for an arbitrary variance-covariance matrix within states over time.

implies that a 10 percentage point increase in Medicaid eligibility reduces consumer bankruptcies by 8 percent. Consistent with the existence of omitted variables and measurement error, the IV estimate is more than three times larger than the OLS estimate.

The remainder of table 2 reports the results of a robustness test involving business bankruptcies. One would expect Medicaid expansions to have little effect on business bankruptcies; few businesses are both nearly bankrupt and employ many individuals eligible for Medicaid. Columns 3 and 4 present OLS and IV results for business bankruptcies. Both point estimates are not statistically distinguishable from zero. The magnitude of the IV estimate is also much smaller in absolute value than the corresponding IV estimate for consumer bankruptcies. Overall, table 2 demonstrates a strong, negative effect of Medicaid eligibility on consumer bankruptcies, and no similar effect on business bankruptcies.

4.1 Robustness Checks

The remainder of this section reports results of additional specifications designed to explore the robustness of these findings and to test alternative explanations for our main results. Table 3 reports results of several alternative specifications designed to test whether the effect of Medicaid expansions can be distinguished from a linear time-trend. The first column reproduces the baseline IV results from table 2. The second column presents the reduced form, and the third column presents results testing whether a two-year lead or lag of simulated Medicaid eligibility is a potential confounder. Reassuringly, the lead effect is much smaller in magnitude than the lag effect, and the lead is not statistically significant at conventional levels.[10] The results therefore suggest that the contemporaneous effect of Medicaid eligibility on consumer

[10] The sum of the contemporaneous and lag effects in column 3 ($-.484$) is roughly the same as the baseline reduced form coefficient on simulated Medicaid eligibility (-0.493).

bankruptcies is not simply a proxy for future changes.

A further concern with the baseline specification is that state bankruptcies may follow unobserved, area-specific trends correlated with Medicaid expansions. The remainder of table 3 investigates this issue. The fourth column presents results that include a linear time-trend for each of nine census regions, while column 5 includes region-year fixed effects. Such controls have little effect on the point estimate of interest. Column 6 presents the results when state-specific linear time-trends are included. Relative to the baseline specification, the magnitude of the point estimate declines substantially and the standard errors increase slightly, from 0.347 to 0.424. Strictly interpreted, column 6 suggests a smaller interaction between Medicaid and bankruptcy, although the confidence interval does not rule out the previous estimates. The point estimate implies that a 10 percentage point expansion of Medicaid would lead to a 3.4 percent decrease in bankruptcies, though it is not statistically significant at conventional levels.

We do not choose column 6 as our preferred specification for two main reasons. Many states rolled out their Medicaid eligibility expansions over time, making eligibility itself well approximated by a state-specific linear trend.[11] Some states, however, had either no significant Medicaid expansions during our sample period or only one major expansion during this time period. We label these states "sharp expansion states," because their Medicaid eligibility trends are much better approximated by a step function than by a single, positively sloped line.[12] Column 7 presents the baseline specification restricted to these 23 states; the coefficient on Medicaid eligi-

[11] Also note that Medicaid expansions may affect a state gradually over time. Once an expansion has been passed by the legislature, a population may take up public insurance only slowly. For instance, Cunningham (2003) estimates that SCHIP dramatically reduced the share of children who are uninsured, but did so after a "slow start."

[12] We categorize a state as a sharp expansion state if it expanded eligibility by more than 2 percentage points two or fewer times within the sample. The sharp expansion states are AK, AL, AZ, CO, IL, KY, LA, MI, MS, MT, NC, ND, NJ, NM, NY, OK, OR, RI, SC, SD, TN, TX, UT, VA, WI, WV, WY.

bility is similar in magnitude and precision to the baseline result. Column 8 adds state-specific linear trends to this subsample. For these states, the point estimate is not substantially affected by the addition of state trends. Our interpretation of these results is that state-specific trends absorb much of the identifying variation for states that expanded Medicaid more smoothly over time.

Nevertheless, the results in column 6 raise the concern that Medicaid expansions may be correlated with unobserved trends within each state. Columns 7 and 8 suggest collinearity between state-specific trends and the Medicaid expansions themselves; however, we cannot rule out conclusively that state-specific trends are partially responsible for some of our findings. For that reason, we focus next on trends in potential confounders.

Table 4 reports estimates of equation (1) with controls for a variety of bankruptcy determinants from the March supplement to the CPS. Such variables have been shown to be proxies for bankruptcy risk by Nelson (1999). We find that higher income or employment is associated with fewer consumer bankruptcies. For instance, column 7 demonstrates that the state unemployment rate strongly predicts consumer bankruptcies. But the coefficient on Medicaid eligibility remains similar in magnitude when these additional controls are included. The final column reports results when we include all of the potential confounding variables. The magnitude of the coefficient on Medicaid eligibility declines by 22 percent and becomes marginally significant (p-value of 0.056).

In addition to the controls above, previous research has shown that changes to bankruptcy exemption laws may also affect bankruptcy rates (Gropp et al., 1997). That pattern alone would not affect our baseline estimates unless exemption laws happened to change at the same time that Medicaid was expanded. Nevertheless, to investigate this issue directly, we obtained data on state bankruptcy exemption levels constructed by Hynes et al. (2004).[13] The data contain information on the levels of homestead exemptions and property exemptions in each state over time. We follow Hynes et al. (2004) and focus on the homestead exemption and property exemption levels for married couples. Table 5 reports results from this analysis. Column 1 reports our baseline results without the District of Columbia, for which we do not have data on exemption levels. The remainder of the columns report results from specifications that include alternative measures of homestead exemptions and personal property exemptions as potential confounders.[14] In all columns, the estimated coefficient on Medicaid eligibility remains precisely estimated and similar in magnitude to our baseline specification. Overall, we find no evidence that changes in homestead exemptions or property exemptions can explain our estimates of the effect of Medicaid eligibility expansions on consumer bankruptcies.[15]

A final concern with our baseline specification is that the short-run effect of Medicaid expansions may differ from the long-run effect. This may occur, for example, due to the adjustment dynamics or due to a precautionary savings motive that responds slowly to policy changes.[16] If bankruptcy rates require several years to adjust to changes in public insurance, then the regressions above may not capture the full, long-run effect. Table 6 explores alternative specifications designed to address this concern. Column 2 presents the results of a regression on three-year averages of all variables. The results are similar to the baseline estimates; the point estimate is

[13] We are grateful to Richard Hynes for assistance with these data.

[14] In columns (2) through (5), we follow Hynes et al. (2004) and study homestead and personal property exemptions separately. In columns (6) and (7) we follow Gropp et al. (1997) and assume that exemptions are fungible. We therefore focus on the maximum of the homestead exemption and the personal property exemption as the relevant measure. It is also worth noting that we CPI-adjust the homestead and personal property exemption levels, though our results are very similar using nominal levels instead.

[15] Consistent with the work of White (1987), table 5 suggests that unlimited exemptions increase consumer bankruptcies.

[16] Specifically, households may save less or borrow more once they are eligible for Medicaid (Gruber and Yelowitz, 1999). Such a decrease in precautionary saving may make the effect of Medicaid on bankruptcies change over time, as some bankruptcies are postponed.

slightly higher than the baseline estimate. Column 3 presents estimates when only four years of data are included (1992, 1996, 2000, and 2004) to measure longer-run responses to changes in eligibility. The point estimates again remain roughly similar to our preferred specification, suggesting that our baseline results do not depend on short-term variation and that the longer-run effects of changes in Medicaid eligibility do not differ significantly from the short-run effects.

4.2 Within-State Analysis

The state-level results above suggest that Medicaid expansions reduce consumer bankruptcies. But the results above rely on aggregate counts of bankruptcies, and are thus unable to test whether the decrease in bankruptcies is concentrated amongst households most likely to qualify for Medicaid. For that reason, we have compiled a database of consumer bankruptcies from the administrative records of bankruptcy courts.

We received access to the records of 33 bankruptcy district courts through the Public Access to Court Electronic Records (PACER) system.[17] The PACER database lists the name and address of each bankruptcy filer. We thus were able to compile annual counts of bankruptcies by zip code for the districts in our database.[18] We then merged those counts to zip code-level demographic information from the 1990 US Decennial Census.[19]

[17] The following bankruptcy district courts had a complete record of consumer bankruptcy filings during our sample period and agreed to provide us with access to their electronic records: AK, AR(E), AR(W), HI, IN(N), IN(S), KY(E), KY(W), LA(E), LA(M), ME, MI(W), MS(N), MS(S), MT, NC(E), NC(M), ND, NE, NM, OH(S), PA(M), PA(W), RI, SD, UT, VA(E), VA(W), VT, WI(E), WI(W), WV(S), WY. We verified that aggregate counts of bankruptcies based on the court dockets matched the state-by-year aggregate totals in the reports from the administrative office of the US courts.

[18] For comparison, we restricted our state-by-year sample to the states for which we could obtain PACER data. The baseline IV specification (equation 1) leads to a point estimate of -0.589 with a standard error of 0.342.

[19] Approximately 13% of bankruptcies in the PACER data cannot be merged to the census data. These zip codes may be recorded incorrectly in the raw PACER data or may be dropped from the

We stratify zip codes based on two characteristics, both captured by the 1990 census. First, Medicaid expansions primarily extended coverage to children, so we expect the impact of the expansions to be concentrated in areas with the most children. We divide zip codes into terciles based on the share of the zip code's population that is younger than age 17. Second, income is a key factor in eligibility, so we expect the impact of the expansions to be concentrated in zip codes with many low- and middle-income households. We divide zip codes into terciles based on the share of households with less than $40,000 in annual income. We then construct counts of bankruptcies for each state and year that measure the number of bankruptcies in zip codes from each tercile.

Table 7 reports the results of our basic IV specification, when restricted to bankruptcies from zip codes in a given tercile. Panel A reports results when zip codes are stratified based on the share of the population younger than age 17, and panel B reports results when zip codes are stratified by household income. In column 1, both panels find a negative effect of eligibility on bankruptcies. The effect, however, is largest for zip codes in the third tercile, zip codes with either many children or many low-income households.[20] Columns 2 and 3 introduce region-specific time trends and state-specific time trends. In all cases, the effect of Medicaid on bankruptcies is largest in zip codes from the third tercile.

We interpret table 7 as broadly consistent with the state-level findings. It suggests that the reduction in bankruptcies was concentrated in families directly affected by the Medicaid expansions.

1990 census files if their population is too small.

[20] The point estimates are much larger than those in our baseline sample. The baseline sample suggests that a 10 percentage point increase in Medicaid eligibility will lead to an 8 percent reduction in bankruptcies; some of the estimates in table 7 suggest nearly a 20 percent reduction. However, the large estimates are for zip codes most likely to be affected by the Medicaid expansions. Therefore, these zip codes likely experienced a disproportionately large increase in Medicaid eligibility for a given 10 percentage point increase in state-wide eligibility.

5 Households' Exposure to Financial Risk From Medical Costs

The results of the previous section suggest a significant and negative effect of Medicaid eligibility on consumer bankruptcies. This section investigates the plausible mechanism behind that finding: Medicaid reduces households' exposure to financial risk through out-of-pocket medical costs.[21] Unfortunately, no data set exists with sufficient sample size that contains both detailed information on household medical costs and bankruptcy status. We instead investigate the financial risk faced by Medicaid-covered households and uninsured low-income households in the cross section.

The Medicaid expansions of the 1990s affected some adults directly, providing coverage regardless of age for households under certain income thresholds. But the expansions were meant primarily to expand coverage for children. Insurance coverage for children only helps households to the extent that children's medical costs are a burden. To measure that burden, we turn to the Medical Expenditure Panel Survey (MEPS). The MEPS collects detailed records on out-of-pocket medical costs for a nationally representative sample of households. We select all low-income households with children in the 1997, 1998, and 1999 MEPS samples.[22] Figure 4 plots the empirical cumulative distribution function (CDF) of the out-of-pocket medical costs for household's children and adults.[23] The figure makes clear that households spend much more money on out-of-pocket medical expenditures for parents than for children.

[21] We have also explored the effect of Medicaid expansions on hospital uncompensated care. Using data from the American Hospital Association (AHA) annual census of U.S. hospitals between 1994 and 1999, we estimate the effect of expanding Medicaid eligibility on hospital bad debt, hospital charity care, and total uncompensated care. The point estimates are uniformly negative and economically large for each outcome, but the very large standard errors prevent us from making any conclusions about the effect of Medicaid on uncompensated care.

[22] Since the MEPS is a two-year panel, we restrict the sample to only the first appearance of each household. We retain all households with income less than 200 percent of the relevant poverty line. All expenditures are in 2000 dollars.

[23] To construct figures 4 and 5, we select households in their first year of survey from the 1997, 1998, and 1999 MEPS survey years. There are 4,460 low-income families in that sample.

This is consistent with the general positive correlation between age and health care utilization. However, the figure also demonstrates that some households still spend substantial sums of money out-of-pocket on their children's health care. Roughly 6.9 percent of households spend more than $1,000 on medical care for their children in a given year, and 1.9 percent spent more than $3,000. Such expenditures represent large sums of money for the low-income households that compose this sample.

Figure 5 performs a similar analysis for the same sample, but divides households into those in which all children are uninsured and those in which all children are covered by Medicaid. For this cross-section, the figure demonstrates that Medicaid-covered households face a dramatically lower risk of large out-of-pocket medical costs for their children. The Kolmogorov-Smirnov test rejects the null hypothesis that the two CDFs are identical (p-value less than 0.001). Roughly 8.9 percent of uninsured households had out-of-pocket expenditures for children beyond $1,000, compared to only 2.3 percent for Medicaid-covered households. Such a cross-sectional pattern does not demonstrate a causal relationship between Medicaid eligibility and children's out-of-pocket medical costs. Nevertheless, figure 5 provides suggestive evidence that Medicaid substantially reduces financial risk, especially in the right tail of the distribution of costs.

Finally, we compare the distribution of out-of-pocket expenditures in the MEPS over time. The first panel of figure 6 compares households' out-of-pocket expenditures for children in the 1997 MEPS sample versus those in the 2004 sample.[24] Between those two survey years, a substantial share of children in the US gained health insurance through SCHIP (variation captured in section 4 by increases in actual and simulated eligibility). The figure demonstrates that households' total out-of-pocket expenditures declined slightly for children. The second panel of figure 6 makes clear

[24] The 1997 sample is the first year of the MEPS. We choose to compare this to the 2004 MEPS, because that is the last year of our simulated eligibility sample.

that over the same period, households experienced an *increase* in out-of-pocket expenditures for adults. The two panels present a type of difference-in-difference analysis, suggesting that over the late 1990s, households' out-of-pocket medical expenditures on children declined relative to the out-of-pocket medical expenditures on adults. This evidence is consistent with Gruber and Levy (2009), who document a similar pattern.[25]

6 Theoretical Implications of the Interaction Between Medicaid and Bankruptcy

This section explores the policy implications of the empirical estimates above. It presents a simple model in which health insurance imperfectly substitutes for bankruptcy. The model provides sufficient statistics that map the interaction between Medicaid and bankruptcy to the optimal generosity of Medicaid (Chetty, 2009). It is based on the one-period optimal insurance problem analyzed by Chetty (2006).

An agent faces two types of shocks: health shocks and productivity shocks. The agent suffers a health shock with fixed probability p_H, and then must choose m units of medical consumption at price $1 - b_H$. Here b_H is the co-insurance rate provided by the government. The value of medical consumption is captured by a concave, increasing function, $v(m)$.[26]

The agent suffers a productivity shock with probability $p_B(e, m)$, where e is effort exerted to avoid the productivity shock. This effort is costly, with convex cost $f(e)$. We assume a stylized version of the bankruptcy system that captures the nature of bankruptcy as social insurance, but for simplicity we do not model the financial decision taken by the debtor. We assume that if the agent suffers a productivity shock, the agent files for bankruptcy and must pay a fixed amount of debt, D, and that the

[25] See table 1 in Gruber and Levy (2009), which documents that out-of-pocket medical cost risks did not decrease for families likely affected by Medicaid expansions.

[26] The shape of the $v(m)$ function determines the ex-post moral hazard in health consumption.

bankruptcy system dissolves a share b_B of that debt.[27] Note that the probability of a productivity shock may depend on whether the agent has also suffered a health shock. This allows out-of-pocket medical costs to directly increase bankruptcy risk, which might be one mechanism through which health insurance benefits affect bankruptcy risk.

Suppose that the social planner imposes a lump-sum tax, τ, in each state of the world. Denote as c the agent's consumption in the case of no shocks. In that case, the agent's consumption is simply her wealth less taxes: $c = W - \tau$. If the agent suffers a health shock but no productivity shock, she chooses m units of medical care, but is partially compensated by the government: $c_H = W - \tau - (1 - b_H) \cdot m$. Similarly, when the agent suffers a productivity shock but no health shock, her consumption is: $c_B = W - \tau - (1 - b_B) \cdot D$. Finally, the agent may suffer both a productivity and a health shock, in which case her consumption is: $c_{BH} = W - \tau - (1 - b_H) \cdot m - (1 - b_B) \cdot D$.

Under these assumptions, the agent solves the following problem:

$$
\begin{aligned}
V^*(b_H, b_B, \tau) \equiv \max_{m,e} \quad & p_H p_B(e,m)(u(c_{BH}) + v(m)) + \\
& (1 - p_H) p_B(e,m) u(c_B) + \\
& p_H(1 - p_B(e,m))(u(c_H) + v(m)) + \\
& (1 - p_H)(1 - p_B(e,m)) u(c) - \\
& f(e).
\end{aligned}
$$

The social planner takes the agent's actions as given and maximizes V^* subject to the resource constraint $\tau = p_H b_H m + p_B(e,m) b_B D$.[28] Optimal health insurance benefits

[27] In reality, debt is likely affected by the ex-ante moral hazard of bankruptcy and the generosity of both insurance systems. One way to incorporate this in the model is to make the choice of D endogenous, which results in an additional elasticity in the optimal insurance formula: the elasticity of D with respect to b_H.

[28] The social planner maximizes V^* with respect to both b_H and b_B. This leads to two formulas for optimal insurance. The joint optimality of both insurance systems is simultaneously determined

must satisfy:

$$\frac{p_B u'(c_{HB}) + (1 - p_B) u'(c_H)}{\bar{u}'} = 1 + \frac{d \log m}{d \log b_H} + \frac{p_B b_B D}{p_H b_H m} \times \frac{d \log p_B}{d \log b_H}, \tag{2}$$

where $\bar{u}'$ is the agent's expected marginal utility of consumption. Equation (2) is analogous to the formula for optimal insurance derived by Baily (1978).[29] The formula demonstrates that a social planner will provide full health insurance if medical consumption does not respond to the health insurance benefit rate and the probability of bankruptcy does not respond to the health insurance benefit rate. If, on the other hand, the right-hand side of equation (2) is greater than 1, then less than full insurance will be socially optimal.

A large literature in health economics has estimated the price elasticity of health consumption (the first elasticity on the right-hand side of equation 2), most notably the RAND health insurance experiment (Manning et al., 1987). Most of the literature estimates a positive but small elasticity. To our knowledge, our study is the first to estimate the second elasticity: the response of the probability of filing for bankruptcy to shifts in the health insurance benefit level. Below, we estimate that this elasticity is negative: Medicaid expansions reduce the bankruptcy rate.

We use this model to calibrate the optimal health insurance benefit taking the bankruptcy system as given. We combine our empirical estimate above with parameters selected from the literature and solve equation (2). We begin by computing the elasticity of bankruptcy risk with respect to the health insurance benefit rate, $\frac{d \log p_B}{d \log b_H}$. Our preferred estimate is that expanding Medicaid eligibility by 10 percentage points

by these two equations. To simplify the exposition, we focus only on the first-order condition for b_H. There are two reasons for this simplification: first, we calibrate the optimal health insurance benefit rate taking the bankruptcy system as given. Second, the first-order condition for b_B includes the moral hazard cost of the bankruptcy system. We are not aware of any estimates of that term.

[29] If there is no bankruptcy system ($p_B = 0$), then equation (2) simplifies to: $\frac{u'(c_H)}{\bar{u}'} = 1 + \frac{d \log m}{d \log b_H}$, the expression derived by Baily (1978).

reduces consumer bankruptcies by 8.1 percent. We take this to be our estimate of $\frac{d \log p_B}{d b_H}$. Assuming $b_H = 0.70$ implies that $\frac{d \log p_B}{d \log b_H} = -0.565$. Given the considerable uncertainty surrounding the coefficient of relative risk aversion, we calibrate the formula for a range of commonly used risk aversion values. The remaining details of the calibration are described in the appendix.

The results of the calibration are reported in Table 8. The first row presents the optimal level of insurance based on parameters meant to capture the current generosity of the bankruptcy system and the current bankruptcy rate. Depending on the coefficient of risk aversion, we compute the optimal health insurance benefit level to be between 72.7 percent and 83.7 percent. The second row of Table 8 presents the same estimates, but under the assumption that there exists no bankruptcy system ($b_B = 0$). In that case, the optimal health insurance benefit increases, since the social planner now uses health insurance as an imperfect substitute for a bankruptcy system. The optimal health insurance benefit rate is between 14 and 24 percent lower in row 1 as compared to row 2, suggesting that the imperfect substitutability between health insurance and bankruptcy can have a significant effect on the optimal health insurance benefit rate. By comparison, we compute that the optimal health insurance benefit rate is 15 percent lower when the coefficient of relative risk aversion moves from 4 to 2. Finally, row 3 of table 8 presents results which alter the economic environment so that there is no longer any risk of bankruptcy ($p_B = 0$). This would be the optimal health insurance benefit if a social planner ignored the risk of bankruptcy entirely. As expected, the estimated benefit level in row 3 is much lower than in rows 1 and 2.

In summary, Table 8 suggests that the optimal benefit level for health insurance depends critically on the interaction with bankruptcy, and whether the social planner recognizes that interaction. The table implies that policy recommendations ought to focus on the joint optimality of social insurance programs, rather than on one social insurance program in isolation. That conclusion may apply not only to bankruptcy and Medicaid, but to other forms of social insurance that are imperfect substitutes (for instance, disability insurance and unemployment insurance).

7 The Share of Bankruptcies Driven by Medical Costs

Previous researchers have claimed that medical costs are pivotal in between 17 to 54 percent of bankruptcies (Himmelstein et al. 2005; Dranove and Millenson 2006). This section develops a simple framework that translates our regression results into estimates directly comparable to such observational studies.

We decompose the overall probability of declaring bankruptcy, $P(B)$, into a conditional bankruptcy rate for the low-income population with health insurance, I, and without health insurance, $\neg I$:

$$P(B) = P(B|I)P(I) + P(B|\neg I)P(\neg I). \tag{3}$$

Suppose that the expansion of Medicaid increases the fraction of the population with health insurance by 10 percentage points (from $P(I)$ to $0.10 + P(I)$), and that this leads to a new bankruptcy rate, $\hat{\beta} \times P(B)$.[30] This leads to the following equation:

$$\hat{\beta} \times P(B) = P(B|I)\left(P(I) + 0.10\right) + P(B|\neg I)\left(\left(P(\neg I) - 0.10\right)\right). \tag{4}$$

Given estimates of $P(B)$, β, $P(I)$, and $P(\neg I)$, equations (3) and (4) form a system of two linear equations with two unknowns: $P(B|\neg I)$ and $P(B|I)$. We choose

[30] It is well documented that an increase in Medicaid eligibility does not translate into a one-for-one increase in health insurance coverage. Like many social insurance programs, the overall take-up rate of Medicaid is low, so many newly eligible households continue to remain uninsured. We consider nominally uninsured but Medicaid-eligible households "conditionally insured," meaning that if such households found themselves in the hospital then the hospital would enroll them in Medicaid.

$P(B) = 0.025$ based on Warren (2003) and $P(I) = 0.70$ based on tabulations from the CPS.[31] We use $\beta = 0.919$ based on our regression results. From equations (3) and (4) we calculate that $P(B|\neg I) = 0.040$ and $P(B|I) = 0.018$. This implies that—ceteris paribus—low-income households without health insurance are roughly two times more likely to file for bankruptcy than insured low-income households.

Universal health insurance for low-income families would simplify the overall bankruptcy rate in (3) to $P(B) = P(B|I)$. Consequently, the fraction of bankruptcies that can be attributed to a lack of health insurance is:

$$\frac{P(B) - P(B|I)}{P(B)} \approx 26\%.$$

This estimate is lower than the 54 percent reported by Himmelstein et al. (2005) and larger than the 17 percent reported by Dranove and Millenson (2006).

A key issue in comparing this estimate to those calculated by observational studies is that our estimates are based on families affected by Medicaid expansions. Out-of-pocket medical costs may be more or less important in the bankruptcy decision of higher-income families. Dranove and Millenson (2006) argue that most "medical bankruptcies" are filed by low-income families. In that case, our estimates can be interpreted as providing an upper bound on the overall importance of out-of-pocket

[31] Overall, roughly 1 percent of households file for bankruptcy in any given year, but bankruptcy risk is higher for low-income households, and bankruptcy filers are more likely to be drawn from the lower half of the income distribution (Warren, 2003). To verify this, we collected data on self-reported household income in the bankruptcy filings of a random sample of recent filers in the Southern District of Ohio. We find strong evidence that bankruptcy filers are more likely to be drawn from the lower half of the income distribution. Figure 7 presents kernel density plots of household income from (a) a sample of households from the 2003 current population survey, and (b) our sample of households filing for bankruptcy. The figure suggests that lower-income households constitute a disproportionate share of bankruptcy filers. Thus households on the margin of Medicaid eligibility have substantially higher probability of filing for bankruptcy than other households. Warren (2003) argue that bankruptcy risk is 2–3 times higher for low-income households, so we choose $P(B) = 0.025$. To estimate the share of low-income households that are uninsured, we calculate the share of uninsured households among households between 100 percent and 200 percent of the federal poverty line using the 1996 CPS.

medical costs on the consumer bankruptcy decision of the average family, and our estimates are more consistent with the work of Dranove and Millenson (2006) than with the work of Himmelstein et al. (2005).

8 Conclusion

This study estimates the effect of Medicaid expansions on personal bankruptcies. The results demonstrate a significant interaction between these two types of insurance: a 10 percentage-point increase in Medicaid eligibility would decrease bankruptcies by about 8 percent. Upon close inspection, these point estimates are economically large, but not implausible. Bankruptcies are disproportionately concentrated in low-income households on the margin of Medicaid eligibility. A 10 percentage point increase in Medicaid eligibility is itself an enormous expansion of social insurance. But in the 1990s, bankruptcies increased by roughly 5 percent each year. Our results therefore suggest that a massive expansion of Medicaid would prevent about one year of 1990s-era growth in consumer bankruptcies.

This paper suggests that medical costs are an important driver of bankruptcies, especially among low-income families. But medical costs cannot fully explain the large increase in consumer bankruptcies over the past thirty years. Between 1994 and 1999, the share of uninsured Americans increased by 7 percentage points (Short, 2001). Our regressions would predict a 7 percent increase in the number of bankruptcies over this period.[32] In reality, bankruptcies increased by 71 percent. Consequently, our estimates only explain roughly 10 percent of the overall increase in bankruptcies.[33]

Taken as a whole, our results suggest that Medicaid affects not only its benefi-

[32] Note that this prediction assumes that our point estimate for increases in Medicaid eligibility is also appropriate for predicting changes in consumer bankruptcies due to changes in the uninsured population.

[33] A similar point is made by Livshits et al. (2007). Canada also experienced an enormous increase in consumer bankruptcies during the 1980s and 1990s. But, during that time period, Canadians enjoyed universal access to health insurance.

ciaries, but also a dispersed group of creditors. Medicaid expansions appear to lead to greater transfers from debtors to creditors. As bankruptcies become less common following Medicaid expansions, lenders may charge lower prices to all other borrowers. The full extent of this pass-through remains an important area for future work.

Appendix

Assuming constant relative risk aversion ($u(c) = (u^{1-\alpha})/(1-\alpha)$), equation (2) becomes:

$$\frac{p_B(c-(1-b_H)m-(1-b_B)D)^{-\alpha}+(1-p_B)(c-(1-b_H)m)^{-\alpha}}{\bar{u}'} = 1+\frac{d\log m}{d\log b_H}+\frac{p_B b_B D}{p_H b_H m}\cdot\frac{d\log p_B}{d\log b_H} \qquad (5)$$

For the reasons described above, we choose $p_B = 0.025$ and $\frac{d\log m}{d\log b_H} = 0.2$. We use $p_H = 0.5$, and we choose $m = 9000$ and $c = 30000$ to roughly match average medical consumption and income among low-income households in MEPS and CPS. We use $D = 28000$ based on average debts discharged during bankruptcy as reported by Barron and Staten (1998). Finally, we assume that $b_B = 0.95$ to account for the non-zero costs of bankruptcy, mainly filing and legal fees. Given these assumptions and equation (5), the optimal health insurance benefits can be solved numerically using equation above.

References

Baily, M. N. (1978, December). Some aspects of optimal unemployment insurance. *Journal of Public Economics 10*, 379–402.

Barron, J. M. and M. E. Staten (1998). Personal bankruptcy: A report on petitioner's ability-to-pay. Discussion paper, Credit Research Center.

Chetty, R. (2006). A general formula for the optimal level of social insurance. *Journal of Public Economics 90*, 1879–1901.

Chetty, R. (2009). Sufficient statistics for welfare analysis: A bridge between structural and reduced-form methods. *Annual Review of Economics 1*, 451–488.

Cunningham, P. (2003, JUL-AUG). SCHIP making progress: Increased take-up contributes to coverage gains. *Health Affairs 22*(4), 163–172.

Currie, J. and J. Gruber (1996, December). Saving babies: The efficacy and cost of recent changes in the Medicaid eligibility of pregnant women. *The Journal of Political Economy 104*(6), 1263–1296.

Cutler, D. and J. Gruber (1996). Does public insurance crowd out private insurance? *Quarterly Journal of Economics 111*(2), 391–430.

DeLeire, T., L. M. Lopoo, and K. I. Simon (2007, February). Medicaid expansions and fertility in the United States. Working Paper 12907, National Bureau of Economic Research.

Dranove, D. and M. L. Millenson (2006). Medical bankruptcy: Myth versus fact. *Health Affairs 25*(2), W74–W83.

Fay, S., E. Hurst, and M. White (2002, June). The household bankruptcy decision. *American Economic Review 92*(3), 706–718.

Gropp, R., J. K. Scholz, and M. J. White (1997, February). Personal bankruptcy and credit supply and demand. *The Quarterly Journal of Economics 112*(1), 217–251.

Gross, D. and N. S. Souleles (2001, August). An empirical analysis of personal bankruptcy and delinquency. Working Paper 8409, National Bureau of Economic Research.

Gruber, J. (2000, August). Medicaid. Working Paper 7829, National Bureau of Economic Research.

Gruber, J. and H. Levy (2009). The evolution of medical spending risk. *The Journal of Economic Perspectives: A Journal of the American Economic Association 23*(4), 25–48. PMID: 19946978.

Gruber, J. and K. Simon (2008). Crowd-out ten years later: Have recent public insurance expansions crowded out private health insurance? *Journal of Health Economics 27*, 201–217.

Gruber, J. and A. Yelowitz (1999, December). Public health insurance and private savings. *Journal of Political Economy 107*(6), 1249–1274.

Himmelstein, D., D. Thorne, E. Warren, and S. Woolhandler (2009, August). Medical bankrutpcy in the United States, 2007: Results of a national study. *The American Journal of Medicine 122*(8), 741–746.

Himmelstein, D. U., E. Warren, D. Thorne, and S. Woolhandler (2005). Illness and injury as contributors to bankruptcy. *Health Affairs*, W65–73.

Hynes, R., E. Posner, and A. Malani (2004). The political economy of property exemption laws. *Journal of Law and Economics 47*(1), 19–43.

Keys, B. (2009). The credit market consequences of job displacement. Working Paper #09-08, National Poverty Center.

Livshits, I., J. MacGee, and M. Tertilt (2007, September). Accounting for the rise in consumer bankruptcies. Working Paper 13363, National Bureau of Economic Research.

Manning, W. G., J. P. Newhouse, N. Duan, E. B. Keeler, A. Leibowitz, and M. S. Marquis (1987). Health insurance and the demand for medical care: Evidence from a randomized experiment. *American Economic Review 77*(3), 251–277.

Nelson, J. P. (1999, October). Consumer bankruptcy and chapter choice: State panel evidence. *Contemporary Economic Policy 17*(4), 552–566.

Short, P. F. (2001). Counting and characterizing the uninsured. Technical Report 2, Economic Research Initiative on the Uninsured Working Paper Series.

Stavins, J. (2000, July/August). Credit card borrowing, deliquency, and personal bankruptcy. *New England Economic Review*, 15–30.

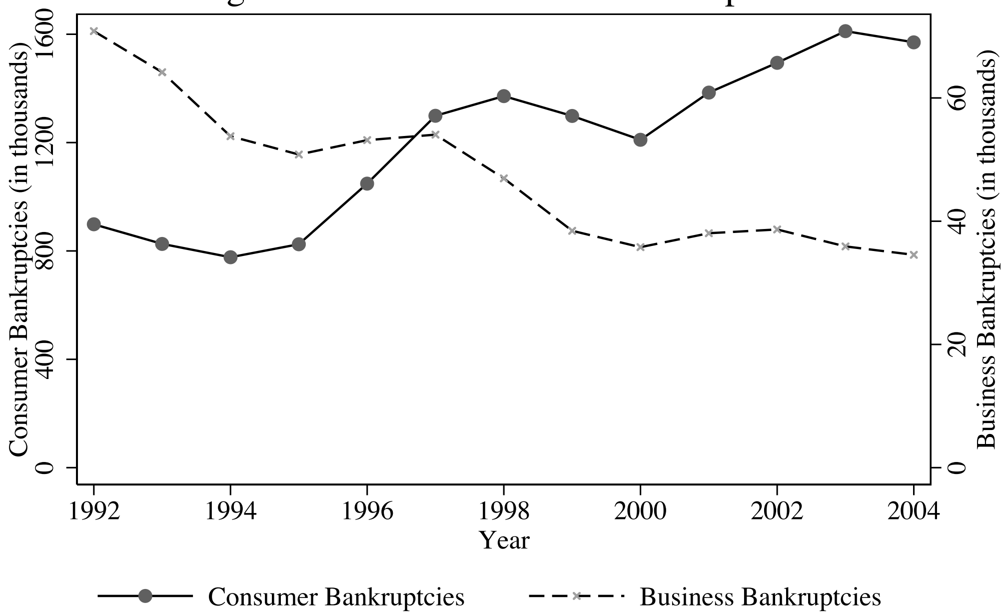
Sullivan, T., E. Warren, and J. L. Westbrook (1989). *As we forgive our debtors: Bankruptcy and consumer credit in America.* Oxford University Press.

Warren, E. (2003). Financial collapse and class status: Who goes bankruptcy? *Osgoode Hall Law Journal 41*(1), 115–146.

White, M. J. (1987). Personal bankruptcy under the 1978 bankruptcy code: An economic analysis. *Indiana Law Journal 63*(1), 1–53.

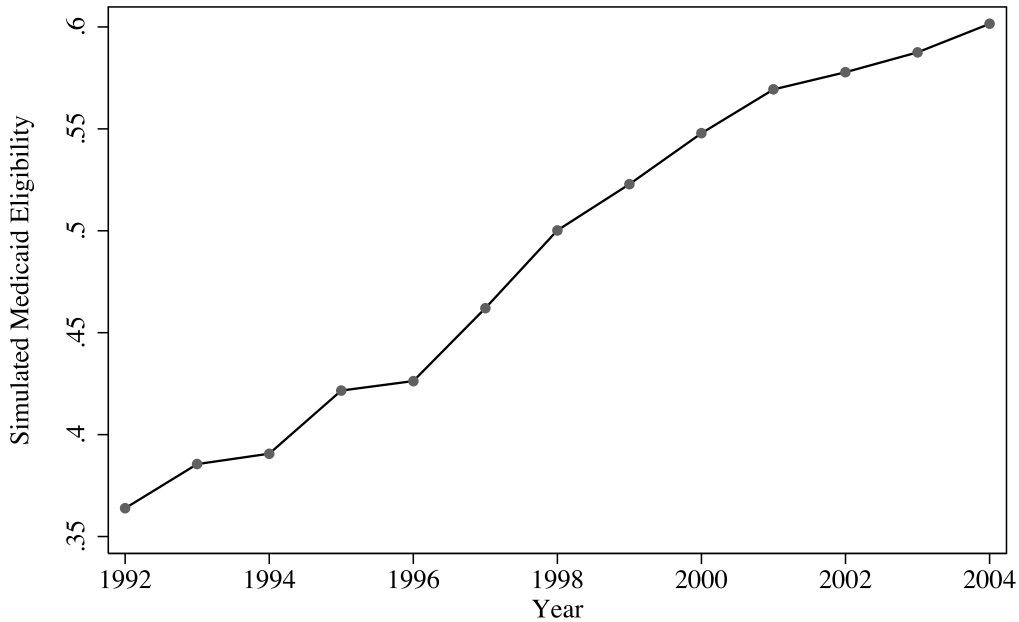
Zywicki, T. J. (2005). An economic analysis of the consumer bankruptcy crisis. *Northwestern University Law Review 99*(4), 1463–1542.

Figure 1: National Trend in Bankruptcies



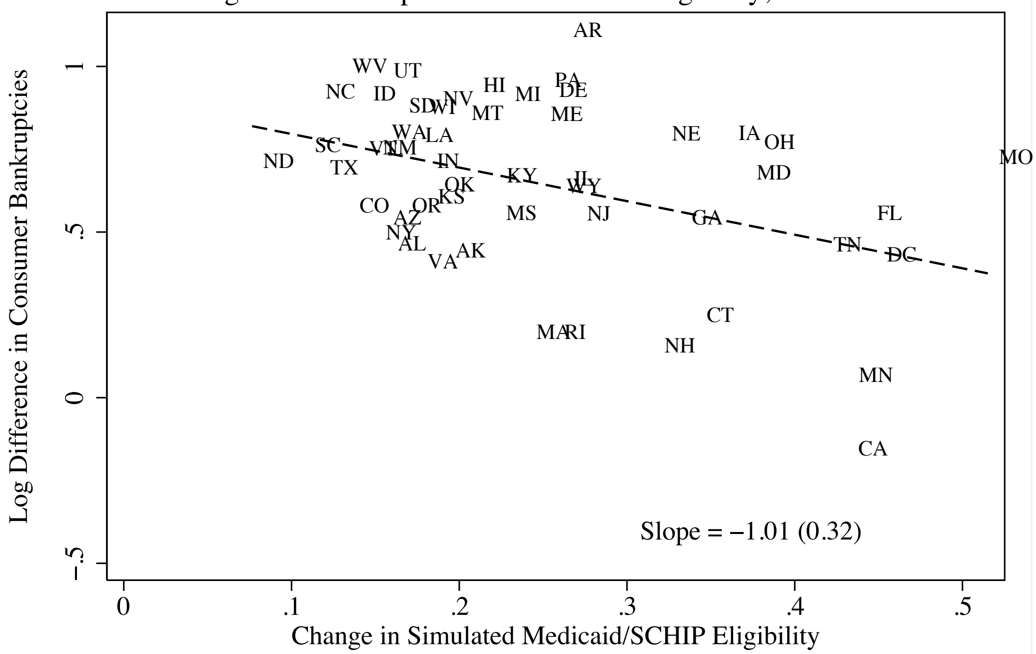
Source: Counts of bankruptcies from the annual census
published by the Administrative Office of the US Courts

Figure 2: Growth in Medicaid Eligibility, 1992–2004



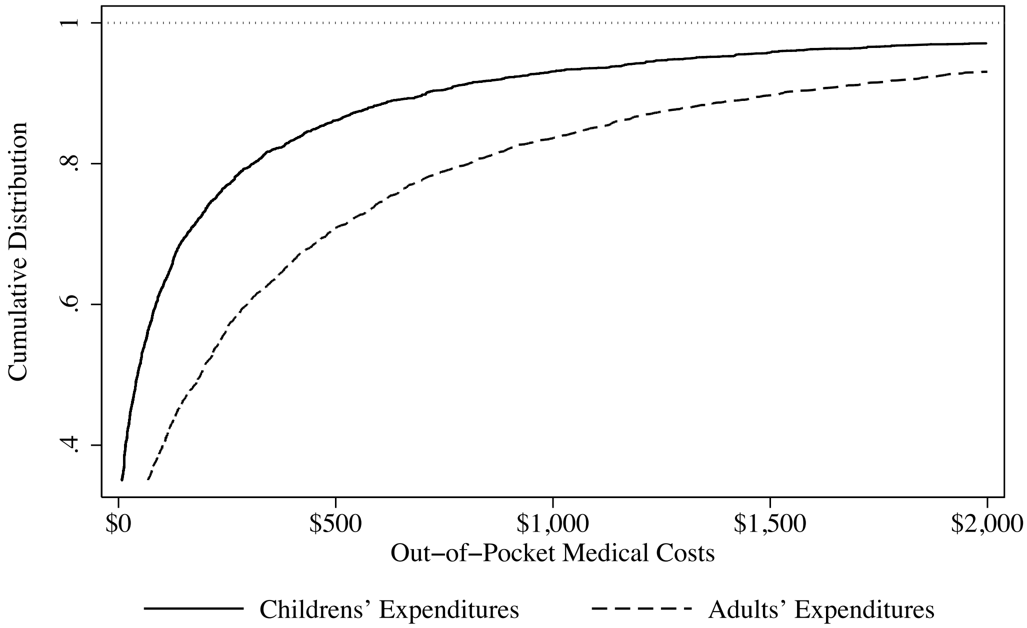
Source: Eligibility is based on our calculations from the
March Current Population Survey. See text for details.

Figure 3: Bankruptcies and Medicaid Eligibility, 1992–2004



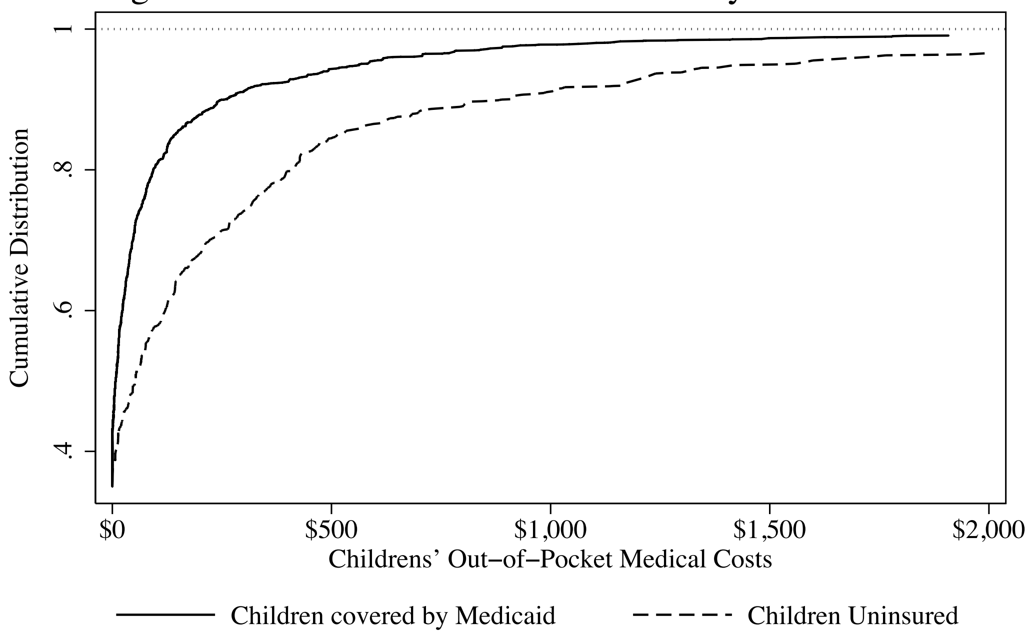
Source: Simulated eligibility calculated from the March CPS. See text for details.

Figure 4: Medical Spending for Children versus Adults



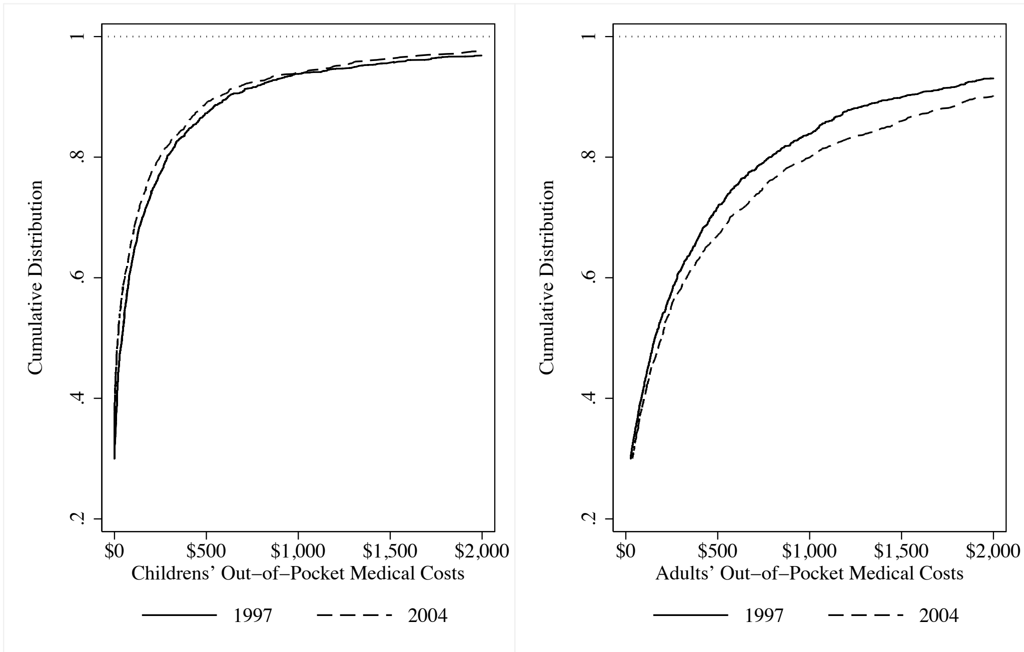
Source: Medical Expenditure Panel Survey

Figure 5: Out−of−Pocket Medical Costs by Insurance Status



Source: Medical Expenditure Panel Survey

Figure 6: Change in Medical Spending, 1997 to 2004



Source: Medical Expenditure Panel Survey

Figure 7: Kernel Density of Household Income

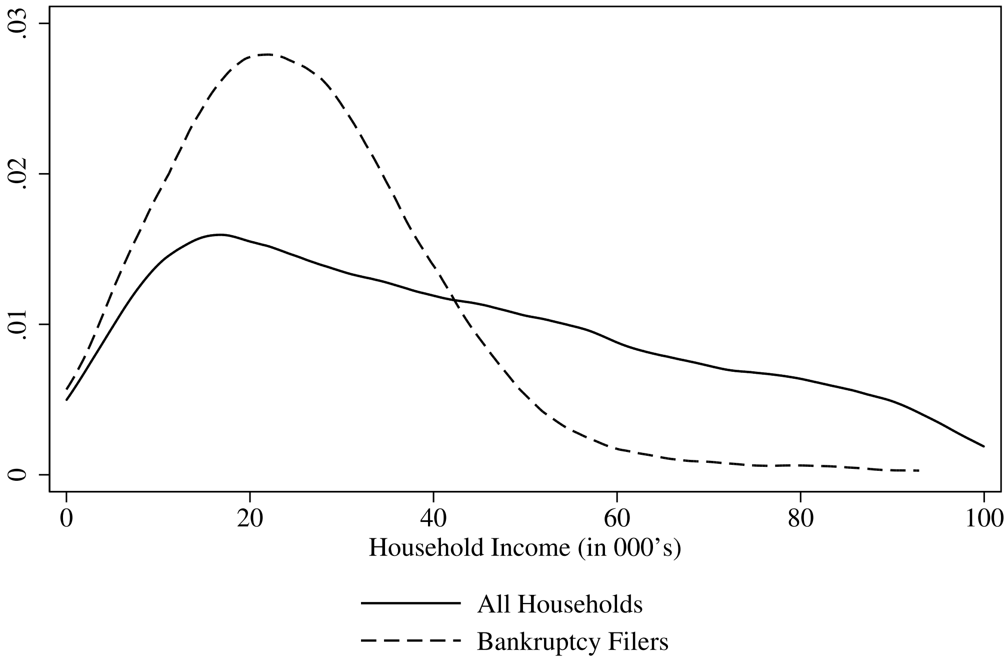


Table 1: Summary Statistics

	Consumer Bankruptcies			Simulated Medicaid Eligibility		
Year	Median	Mean	Standard Deviation	Median	Mean	Standard Deviation
			A. All States			
All Years	15,162	23,552	27,346	0.470	0.489	0.128
1992	13,212	17,615	22,798	0.336	0.364	0.052
1996	14,995	20,565	25,757	0.453	0.426	0.078
2000	16,105	23,732	25,565	0.515	0.548	0.114
2004	23,641	30,791	29,127	0.577	0.602	0.109
			B. Small Expansion States			
All Years	11,249	21,235	23,541	0.449	0.450	0.064
1992	7,196	14,421	17,874	0.441	0.406	0.051
1996	8,886	18,077	21,747	0.453	0.418	0.052
2000	11,569	21,362	22,959	0.507	0.474	0.059
2004	15,449	31,265	35,508	0.552	0.515	0.063
			C. Large Expansion States			
All Years	25,607	52,650	57,253	0.707	0.636	0.185
1992	18,020	45,847	57,766	0.336	0.360	0.054
1996	20,162	50,540	65,365	0.457	0.539	0.116
2000	25,607	52,007	58,933	0.730	0.765	0.058
2004	37,298	53,478	50,838	0.785	0.813	0.051

Notes: The sample consists of bankruptcy counts for all 50 states and the District of Columbia from 1992 through 2004; all observations are state-year. For the purposes of this table only, we define "small expansion states" as the five states with the smallest change in simulated eligibility between 1992 and 2004 (South Carolina, Texas, North Carolina, North Dakota, and West Virginia). The "large expansion states" are defined similarly (and are California, Missouri, Florida, Minnesota, and the District of Columbia).

Table 2: The Effect of Medicaid on Bankruptcy Declarations

	(1)	(2)	(3)	(4)
Dependent Variable:	Log Consumer Bankruptcies		Log Business Bankruptcies	
	OLS	IV	OLS	IV
Medicaid Eligibility	-0.264	-0.807	0.399	0.268
	(0.288)	(0.347)	(0.453)	(0.585)
	[0.365]	[0.024]	[0.382]	[0.649]
R^2	0.99		0.93	

Notes: *N* = 663. The sample consists of bankruptcy counts for all 50 states and the District of Columbia from 1992 through 2004; all observations are state-year. All specifications include state fixed effects and year fixed effects. Robust standard errors in parentheses account for correlation between observations within the same state; related *p*-values are in brackets.

Table 3: Alternative Specifications Investigating Timing and Trends

	(1) IV	(2) OLS	(3) OLS	(4) IV	(5) IV	(6) IV	(7) IV	(8) IV
Dependent Variable: Log Consumer Bankruptcies								
	Baseline	Reduced Form	Reduced Form with Lead and Lag	Region Trends	Region × Year Fixed Effects	State Trends	Sharp Expansion States Only	Sharp Expansion States Only w/ State Trends
Medicaid Eligibility	-0.807			-0.794	-0.826	-0.349	-0.988	-0.903
	(0.347)			(0.334)	(0.407)	(0.424)	(0.275)	(0.301)
	[0.024]			[0.021]	[0.048]	[0.415]	[0.002]	[0.008]
Simulated Medicaid Eligibility		-0.493	-0.339					
		(0.209)	0.251					
		[0.022]	[0.182]					
Simulated Medicaid Eligibilty, 2-year Lag			-0.259					
			(0.160)					
			[0.113]					
Simulated Medicaid Eligibilty, 2-year Lead			0.102					
			(0.189)					
			[0.591]					
N	663	663	459	663	663	663	221	221

Notes: In all specifications except column 2 and 3, Medicaid eligibility is predicted with simulated Medicaid eligibility as an instrumental variable. In all columns except 3, 7, and 8, the sample consists of bankruptcy counts for all 50 states and the District of Columbia from 1992 through 2004. In column 3, only the years 1994 through 2002 are included because of the inclusion of 2-year lead and 2-year lag. In columns 7 and 8 only states with at most one "sharp expansion" in Medicaid between 1992 and 2004 are included; see text for details. All specifications include state fixed effects and year fixed effects. Robust standard errors in parentheses account for correlation between observations within the same state; related *p* -values are in brackets.

Table 4: Alternative Specifications that Control for
Other Determinants of Consumer Bankruptcies

	Dependent Variable: Log Consumer Bankruptcies									
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Medicaid Eligibility	-0.807	-0.708	-0.753	-0.736	-0.728	-0.798	-0.709	-0.830	-0.806	-0.586
	(0.347)	(0.335)	(0.334)	(0.333)	(0.326)	(0.345)	(0.327)	(0.361)	(0.348)	(0.300)
	[0.024]	[0.040]	[0.029]	[0.032]	[0.030]	[0.025]	[0.035]	[0.026]	[0.025]	[0.056]
25th Percentile of Log Wage Distribution		-0.428								-0.172
		(0.163)								(0.143)
		[0.011]								[0.234]
10th Percentile of Log Wage Distribution			-0.210							-0.048
			(0.091)							(0.067)
			[0.025]							[0.474]
Share on Food Stamps				1.396						0.275
				(0.638)						(0.500)
				[0.033]						[0.584]
Share below poverty level					1.177					0.106
					(0.559)					(0.421)
					[0.040]					[0.801]
Share Divorced						0.781				1.197
						(1.034)				(0.932)
						[0.454]				[0.205]
Unemployment Rate							4.429			3.288
							(1.526)			(1.252)
							[0.005]			[0.011]
Log of business bankruptcies								0.084		0.069
								(0.049)		(0.045)
								[0.094]		[0.130]
Share Self Employed									-0.387	0.174
									(0.984)	(0.874)
									[0.696]	[0.843]

Notes: $N = 663$. In all specifications Medicaid eligibility is predicted with simulated Medicaid eligibility as an instrumental variable. The sample consists of bankruptcy counts for all 50 states and the District of Columbia from 1992 through 2004; all observations are state-year. All specifications include state fixed effects and year fixed effects. Robust standard errors in parentheses account for correlation between observations within the same state; related *p*-values are in brackets.

Table 5: State Bankruptcy Exemption Levels as Potential Confounders

	Dependent Variable: Log Consumer Bankruptcies								
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Medicaid Eligibility	-0.844	-0.869	-0.863	-0.896	-0.854	-0.860	-0.877	-0.862	-0.851
	(0.359)	(0.350)	(0.358)	(0.348)	(0.352)	(0.357)	(0.349)	(0.348)	(0.349)
	[0.023]	[0.016]	[0.020]	[0.013]	[0.019]	[0.020]	[0.015]	[0.017]	[0.018]
Indicator for Unlimited Homestead Exemption		0.122		0.145	0.103		0.123		
		(0.082)		(0.083)	(0.076)		(0.077)		
		[0.142]		[0.086]	[0.180]		[0.113]		
Homestead Exemption (in millions)		-0.084		-0.075					
		(0.095)		(0.095)					
		[0.378]		[0.433]					
Indicator for Unlimited Property Exemption			0.102	0.119		0.088	0.097		
			(0.075)	(0.087)		(0.076)	(0.083)		
			[0.183]	[0.178]		[0.251]	[0.245]		
Property Exemption (in millions)			-0.295	-0.389					
			(0.204)	(0.240)					
			[0.154]	[0.112]					
Log Homestead Exemption					-0.143		-0.119		
					(0.191)		(0.188)		
					[0.456]		[0.531]		
Log Property Exemption						-0.392	-0.498		
						(0.324)	(0.349)		
						[0.232]	[0.160]		
Indicator for {Unlimited Homestead OR Unlimited Property Exemption}								-0.089	
								(0.093)	
								[0.344]	
Maximum of Homestead and Property Exemption								0.114	0.099
								(0.074)	(0.069)
								[0.130]	[0.160]
Log Maximum of Homestead and Property Exemption									-0.167
									(0.182)
									[0.366]

Notes: N = 650. In all specifications Medicaid eligibility is predicted with simulated Medicaid eligibility as an instrumental variable. The sample consists of bankruptcy counts for all 50 states from 1992-2004 (DC is excluded due to lack of data on homestead and property exemptions). All observations are state-year. All specifications include state fixed effects and year fixed effects. Standard errors, adjusted to allow for an arbitrary variance-covariance matrix within each state over time, are in parentheses and p -values are in brackets.

Table 6: Short-run versus Long-run Effects

Dependent Variable: Log Consumer Bankruptcies

	(1) Baseline	(2) 3-Year Averages	(3) 1992, 1996, 2000, 2004 Panel
Medicaid Eligibility	-0.807	-0.846	-0.865
	(0.347)	(0.355)	(0.436)
	[0.024]	[0.021]	[0.053]
N	663	204	204

Notes: In all specifications Medicaid eligibility is predicted with simulated Medicaid eligibility as an instrumental variable. The baseline sample consists of bankruptcy counts for all 50 states and the District of Columbia from 1992 through 2004; remaining columns report results using alternative samples. All observations are state-year. All specifications include state fixed effects and year fixed effects. Robust standard errors in parentheses account for correlation between observations from the same state; related *p* -values are in brackets.

Table 7: Effect of Medicaid on Bankruptcy Declarations,
Within-State Evidence from Court Dockets

Dependent Variable: Log Consumer Bankruptcies by State, based on Selected Zip Code Groups									
	(1a)	(1b)	(1c)	(2a)	(2b)	(2c)	(3a)	(3b)	(3c)
	First Tercile	Second Tercile	Third Tercile	First Tercile	Second Tercile	Third Tercile	First Tercile	Second Tercile	Third Tercile
Panel A: Bankruptcies stratified by share of zipcode population younger than 17									
Medicaid Eligibility	- 0.427	- 0.217	- 0.882	- 0.738	- 0.423	- 1.014	- 0.554	- 0.635	- 1.179
	(0.257)	(0.305)	(0.487)	(0.316)	(0.299)	(0.454)	(0.480)	(0.375)	(0.618)
	[0.112]	[0.486]	[0.085]	[0.030]	[0.173]	[0.037]	[0.262]	[0.106]	[0.071]
Panel B: Bankruptcies stratified by share of households in zipcode with less than $40,000 in income									
Medicaid Eligibility	- 0.461	- 0.454	- 1.760	- 0.660	- 0.597	- 1.785	- 0.798	- 0.959	- 2.679
	(0.282)	(0.236)	(0.882)	(0.301)	(0.289)	(0.921)	(0.448)	(0.366)	(1.680)
	[0.117]	[0.069]	[0.060]	[0.040]	[0.052]	[0.067]	[0.090]	[0.016]	[0.127]
Region-specific				X	X	X			
State-specific trends							X	X	X

Notes: N = 273. The dependent variable is the logarithm of bankruptcy counts by state and year for selected groups of zip codes. Zip codes are grouped based on the share of their population under age 17 (panel A), or by share of households with less than $40,000 in annual income (panel B). Both variables for stratification are measured in the 1990 census. Sample includes data between 1992 and 2004. All specifications include state fixed effects and year fixed effects. Robust standard errors in parentheses account for correlation between observations from the same state; related p -values are in brackets.

Table 8: Optimal Level of Insurance, b_H

	Coefficient of Relative Risk Aversion		
	2	3	4
Average Bankruptcy Risk (p_B = 0.025), and Current Bankruptcy System (b_B = 0.95)	72.7%	79.5%	83.7%
Average Bankruptcy Risk (p_B = 0.025), and No Bankruptcy System (b_B = 0.00)	95.4%	97.2%	97.9%
No Bankruptcy Risk (p_B = 0.0)	38.8%	57.9%	67.9%

Notes: This table presents results from a calibration exercise which determines the optimal health insurance benefit rate, b_H, given other parameters. See accompanying text and appendix for details.

Appendix Table 1: First Stage Regressions

Dependent Variable: Consumer Bankruptcies	(1) Baseline	(2) Region Trends	(3) Region × Year Fixed Effects	(4) State Trends
Simulated Medicaid Eligibility	0.611	0.623	0.632	0.467
	(0.047)	(0.038)	(0.047)	(0.038)
	[0.000]	[0.000]	[0.000]	[0.000]
R^2	0.931	0.935	0.941	0.958
First Stage F-statistic	166.56	268.80	177.95	147.25

Notes: N = 663. The sample consists of bankruptcy counts for all zip codes in districts for which we could obtain bankruptcy records, 1992-2004. All observations are state-year. All specifications include state fixed effects and year fixed effects. Standard errors, adjusted to allow for an arbitrary variance-covariance matrix within each state over time, are in parentheses and *p* -values are in brackets.

Appendix Table 2: Reverse Regressions to Investigate Policy Endogeneity

	Dependent Variable: Simulated Medicaid Eligibility									
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
3rd Lag of Log Consumer Bankruptcies	0.082	0.084	0.082	0.076	0.081	0.081	0.082	0.083	0.080	0.081
	(0.067)	(0.070)	(0.069)	(0.067)	(0.068)	(0.071)	(0.068)	(0.067)	(0.067)	(0.075)
	[0.232]	[0.236]	[0.242]	[0.258]	[0.239]	[0.258]	[0.232]	[0.219]	[0.233]	[0.284]
25th Percentile of Log Wage Distribution		0.028								0.028
		(0.054)								(0.048)
		[0.603]								[0.569]
10th Percentile of Log Wage Distribution			0.003							-0.016
			(0.028)							(0.026)
			[0.900]							[0.543]
Share on Food Stamps				-0.163						-0.039
				(0.312)						(0.316)
				[0.603]						[0.901]
Share below poverty level					-0.183					-0.162
					(0.217)					(0.216)
					[0.402]					[0.456]
Share Divorced						-0.020				0.041
						(0.401)				(0.397)
						[0.961]				[0.918]
Unemployment Rate							-0.231			-0.198
							(0.331)			(0.316)
							[0.490]			[0.533]
Log of business bankruptcies								0.010		0.011
								(0.011)		(0.012)
								[0.356]		[0.363]
Share Self Employed									0.286	0.205
									(0.355)	(0.343)
									[0.424]	[0.554]
R^2	0.894	0.894	0.894	0.894	0.894	0.894	0.894	0.894	0.894	0.895

Notes: *N* = 510. The sample consists of bankruptcy counts for all 50 states and the District of Columbia between 1995 through 2004; all observations are state-year. All specifications include state fixed effects and year fixed effects. Robust standard errors in parentheses account for correlation between observations within the same state; related *p*-values are in brackets.