

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

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Abstract

Anecdotal evidence and several observational studies suggest that out-of-pocket medical costs are pivotal in a large fraction of consumer bankruptcies. 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 the personal bankruptcy rate by 8.4 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. We conclude with simple calibration exercises, which suggest that out-of-pocket medical costs are pivotal in roughly 26 percent of personal bankruptcies among low-income households.

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1 Introduction

Bankruptcy protection is a legal procedure designed to forgive debtors their debt. Once undertaken by few debtors (Zywicki, 2005), bankruptcy has become common over the past two decades. 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), for instance, calculates that 8.5 percent of American households have filed for bankruptcy.

This trend has motivated research on factors that induce households to declare bankruptcy. One such factor is the burden of out-of-pocket medical costs. It is often argued that a large fraction of consumer bankruptcies are driven by the costs of health care. This conjecture has been widely publicized and has even motivated legislation to prevent “medical bankruptcies.”¹ A recent 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 evidence regarding the relative importance of medical costs in the decision to declare bankruptcy. The few studies that have attempted to quantify the contribution of medical costs to bankruptcy rely mostly on interviews with members of households that have recently filed for bankruptcy. Such interviews cannot credibly isolate whether bankruptcy filers who did experience high medical costs would have otherwise declared bankruptcy.

To our knowledge, this paper is the first to use plausibly exogenous variation—in this case, expansions of publicly provided health insurance—to document the role of adverse shocks in consumer bankruptcies. In the 1990s, states expanded access to

¹For example, the American Association of Retired Persons has publicized anecdotal evidence on medical bankruptcy as part of its political campaign, “Divided We Fail.”

publicly-provided health insurance by expanding eligibility for Medicaid and through the State Children’s Health Insurance Program (SCHIP). Medicaid and health insurance through SCHIP dramatically decrease the medical costs faced by households. Using cross-state variation in these expansions from 1992 to 2004, we find that Medicaid and SCHIP eligibility also reduce bankruptcy risk. In our preferred specification, we calculate that a 10 percentage point increase in eligibility for publicly-provided insurance reduces the personal bankruptcy rate by 8 percent.

Although we find evidence of an interaction between Medicaid and consumer bankruptcy, we do not conclude that *most* bankruptcies are driven by medical costs, a claim that has been made by other researchers. We employ a calibration exercise to translate our regression results into estimates of the share of bankruptcies driven by medical costs. We estimate that medical costs are pivotal in roughly 26 percent of bankruptcies by low-income households. This share is much smaller than estimates previously put forward in observational studies.

The first contribution of this paper is to isolate the bankruptcy-related benefits of Medicaid. In that way, we can estimate the relative importance of medical costs in the consumer bankruptcy decision. The second contribution of this paper involves the normative implications of that finding. We demonstrate substitution between two types of social insurance: bankruptcy and Medicaid.² We present a theoretical model that examines the interaction between Medicaid and the consumer bankruptcy decision.³ The model suggests that the joint optimality of both programs must take this interaction into account. When we use our empirical results to calibrate the model, we estimate that the optimal health insurance benefit rate is 17 percent higher

²In this sense, the study is similar to studies by Finkelstein and McKnight (2008), documenting the financial benefits of Medicare and by Gruber (1997) regarding the consumption-smoothing benefits of unemployment insurance.

³For simplicity, we refer to both Medicaid and SCHIP simply as “Medicaid,” even though SCHIP provides health insurance to children through programs that are technically distinct from Medicaid.

than would be suggested by conventional models, which focus only on consumption-smoothing benefits and moral hazard costs. In principle, our model applies not just to Medicaid and bankruptcy, but also to other forms of imperfectly substitutable social insurance.

The remainder of the paper proceeds as follows. The subsequent section discusses the state of research on personal bankruptcy. Section 3 develops a model of the interaction between bankruptcy and Medicaid and discusses the normative implication of such an interaction. Section 4 describes state Medicaid expansions and discusses our empirical strategy. Section 5 presents our main results. Section 6 estimates the share of bankruptcies driven by medical costs. Section 7 calibrates the model with our empirical results. Section 8 concludes.

2 Previous Research on the Determinants of Consumer Bankruptcy

A large literature has 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 document that households are forward-looking and optimally choose whether or not to file for bankruptcy based on the expected 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 creates an ex-ante moral hazard problem.

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 (PSID) 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. A study by Himmelstein et al. (2005), for example, 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 cited “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 (for instance, Sullivan et al. 1989).

A concern with such observational studies, 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.

A concern with both strands of research is that the studies do not employ quasi-experimental variation in the determinants of bankruptcy. This empirical challenge cannot be overcome by exploiting changes in the laws that govern bankruptcy, since such laws rarely change.⁴ The state laws that govern homestead exemptions, for example, have rarely been modified since their inception (Gropp et al., 1997). 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 medical costs.

3 Theoretical Implications of the Interaction between Medicaid and Bankruptcy

This section offers a simple model of the interaction between bankruptcy and Medicaid.⁵ The goal of the model is to provide sufficient statistics (Chetty, 2009) to allow us to calibrate the optimal health insurance benefit rate when health insurance would affect the probability of filing for bankruptcy.

The 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, $v(m)$.⁶

⁴A recent exception is the Bankruptcy Abuse Prevention and Consumer Protection Act, passed by Congress in 2005.

⁵Our model is based on the one-period optimal insurance problem analyzed by Chetty (2006).

⁶Note that the shape of the $v(m)$ function determines the ex-post moral hazard in health consumption.

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 explicitly 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 bankruptcy system dissolves a share b_B of that debt.⁷ 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, so that: $c_H = W - \tau - (1 - b_H) \cdot m$. 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: $c_{BH} = W - \tau - (1 - b_H) \cdot m - (1 - b_B) \cdot D$.

⁷In 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 .

Under these assumptions, the agent solves the following problem:

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

The social planner takes the agent's actions as given and maximizes V^* subject to the resource constraint $\tau = p_H b_H m + p_B(e, m)b_B D$.⁸ Optimal health insurance benefits must satisfy:

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

where \bar{u}' is the agent's expected marginal utility of consumption. Equation (1) is analogous to the formula for optimal insurance derived by Baily (1978).⁹ 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 (1) 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

⁸The 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 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.

⁹If there is no bankruptcy system ($p_B = 0$), then equation (1) simplifies to: $\frac{u'(c_H)}{\bar{u}'} = 1 + \frac{d \log m}{d \log b_H}$, the expression derived by Baily (1978).

consumption (the first elasticity on the right-hand side of equation 1), 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 bankruptcy rate to the generosity of health insurance. Below, we estimate that this elasticity is negative, Medicaid expansions reduce the bankruptcy rate. Based on our model, this negative interaction suggests a larger health insurance benefit rate than implied by standard calculations. In section 7, we calibrate equation (1) with our empirical results in order to explore the implications of this model for the joint optimality of both insurance programs.

4 Empirical Strategy and Data

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

4.1 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 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 and Simon 2008 and Gruber 2000 for more details on the Medicaid program). Many states expanded eligibility for parents in conjunction with their SCHIP expansions. 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 Medicaid eligibility from 1992 through 2004.

Overall, roughly 20 percent of all U.S. households became eligible for Medicaid during this time.

These expansions may have affected the financial standing of adults through several mechanisms. First, the expansions may have lowered the financial burden on parents by providing health insurance for their children. The expansions also affected adults directly, providing coverage regardless of age for those under certain income thresholds. But the expansions may have also affected uninsured adults who were not directly made eligible for Medicaid. If the expansions decreased the amount of uncompensated care provided by hospitals, then hospitals may have been more willing to provide free care to patients who were uninsured but not eligible for Medicaid. In this way, expansions targeted only at children would have affected the financial resources of unrelated adults.

To demonstrate the potential effects of the Medicaid expansions on consumer finances, 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. Figure 3 plots the distribution of out-of-pocket medical costs for 2 groups of households: those with at least one family member eligible for Medicaid and those with no family members eligible. For this cross-section, the figure demonstrates that Medicaid beneficiaries face a dramatically lower risk of large out-of-pocket medical costs. Roughly 2 percent of the uninsured spend more than \$5,000 in out-of-pocket medical costs, while less than 0.2 percent of Medicaid beneficiaries spend more than \$5,000 in out-of-pocket medical costs. Such a cross-sectional pattern does not conclusively demonstrate a causal relationship. Unfortunately, the MEPS sample is too small to construct an instrumental variable for Medicaid participation. Nevertheless, figure 3 provides suggestive evidence that Medicaid substantially reduces financial risk, especially in the right tail. Our regressions below test whether

this potential drop in financial risk ultimately lowers the probability of bankruptcy.

4.2 Data

Our investigation into bankruptcy and public insurance requires accurate measures of both types of insurance. For the former, 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 and has been used in related studies (see, for example, Fay et al. 2002.) The census is composed of simple counts of cases for each bankruptcy district since the 1980s. 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.¹⁰

We construct measures of public insurance eligibility from the 1992–2004 March Current Population Survey (CPS).¹¹ First, we calculate whether each surveyed household is eligible for Medicaid in their 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 the state-year’s 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.¹²

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

¹⁰The excluded bankruptcy districts are those in the Virgin Islands, Puerto Rico, Northern Mariana Islands, and Guam.

¹¹The 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 structural change.

¹²We are grateful to Kosali Simon for computer code that constructs these two variables.

during our sample period.¹³ For the “small expansion states,” bankruptcy counts more than doubled, growing from an average of 14,336 in 1992 to 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.

4.3 Empirical Strategy

Figure 4 summarizes our approach and main results. The figure plots for each state the difference in log consumer bankruptcies between 1992 and 2004 against the change in simulated Medicaid eligibility over that time period. The figure demonstrates that states with larger Medicaid expansions 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. Figure 4 suggests that a 10 percentage point increase in Medicaid eligibility reduces consumer bankruptcies by roughly ten percent. In what follows, we use a regression framework to rigorously test this pattern.

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

$$\log(c_{it}) = \alpha_0 + \alpha_i + \alpha_t + \beta M_{it} + \varepsilon_{it}, \quad (2)$$

where c_{it} denotes the number of consumer bankruptcies in state i and year t , M_{it} denotes the fraction of the population eligible for Medicaid, and ε_{it} represents unobserved state-year shocks that affect the number of consumer bankruptcies.

Simply estimating equation (2) with ordinary least squares (OLS) would lead to biased estimates of β . Adverse economic shocks will lead to more consumer bankruptcies and to more households qualifying for Medicaid. Instead, we use simulated

¹³The 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.

Medicaid eligibility as an instrumental variable (IV) for actual Medicaid eligibility.¹⁴ Simulated Medicaid eligibility is correlated with actual Medicaid eligibility (the t -statistic for simulated eligibility from our first stage regression is 12.78), but is assumed not to be correlated with adverse economic shocks. This identifying assumption requires that—absent changes in Medicaid eligibility—state bankruptcy rates would have evolved similarly over time. We begin by estimating equation (2) using instrumental variables under this assumption. We then investigate the validity of this assumption in several ways, exploring our methodology’s robustness to state trends in bankruptcy rates and to time-varying control variables.

5 The Aggregate Effect of Medicaid on Bankruptcies

Table 2 presents our main results. Column 1 shows the OLS relationship between Medicaid eligibility and state consumer bankruptcy filings. This relationship is moderately negative and not statistically significant at conventional levels. Column 2 reports the IV estimates; the magnitude of the point estimate implies that a 10 percentage point increase in Medicaid eligibility reduces consumer bankruptcies by 8.4 percent. The magnitude of the IV estimate is over 3 times as large as the OLS estimates. That pattern is to be expected; unobserved, adverse shocks are positively correlated with both bankruptcies and actual Medicaid eligibility. As a result, the omitted variables bias is positive.

The remainder of table 2 reports the results of a falsification test. One would expect Medicaid to have little impact on business bankruptcies; few businesses are

¹⁴Simulated 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).

both nearly bankrupt and have many employees 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 relationship between Medicaid eligibility and consumer bankruptcies, but no relationship for business bankruptcies.

We turn next to specification tests designed to explore the robustness of these findings. Table 3 reports results of several alternative specifications. Column 1 reproduces the baseline IV results from table 2. Column 2 presents reduced-form estimates that test 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. The results suggest that the contemporaneous effect of Medicaid eligibility on consumer bankruptcies is not simply a proxy for future changes.

An immediate concern with the baseline specification is that state bankruptcies may follow unobserved, area-specific trends correlated with Medicaid expansions. The remainder of table 3 addresses such a concern by testing whether the effect of Medicaid expansions can be distinguished from a linear time-trend. Column 3 presents results that include a linear time-trend for each of nine census regions. Such region time-trends have little effect on our estimates or precision. Column 4 includes region-year fixed effects, and the results are also very similar to the baseline results. Finally, column 5 presents the results of a more stringent test: including state-specific linear time-trends. Relative to the baseline specification, the magnitude of the point estimate declines substantially and the standard errors increase only slightly.

Strictly interpreted, column 5 suggests a smaller interaction between Medicaid and bankruptcy, although the estimate's confidence interval does not rule out the

estimates in the previous columns. The point estimate implies that a 10 percentage point expansion of Medicaid would lead to a 3.4 percent decrease in bankruptcies, but this estimate is statistically indistinguishable from zero. With only eleven years of data for each state, we view this specification to be quite demanding. Moreover, many states rolled out their Medicaid eligibility expansions over time, making eligibility itself well approximated by a state-specific linear trend.¹⁵

Some states, however, had either no significant Medicaid expansions during our sample period or had 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.¹⁶ Column 6 presents the baseline specification restricted to these 23 states; the coefficient on Medicaid eligibility is similar in magnitude and precision to the baseline result. Column 7 adds state-specific linear trends to this subsample. For these states, the point estimate is not significantly 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 smoothly over time. Nevertheless, the results in column 5 raise the concern that Medicaid expansions may be correlated with unobserved trends within each state. For that reason, we focus next on potential confounders.

Table 4 reports the results of our baseline estimates, controlling for a variety of consumer bankruptcy determinants. Column 1 reproduces our baseline IV estimates,

¹⁵Also 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.” As such, one would expect a time-trend to absorb much of the effect of SCHIP expansions.

¹⁶We 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.

in which simulated Medicaid eligibility is used as an instrument for actual Medicaid eligibility, and fixed effects are included for state and year. Subsequent columns control for proxies of bankruptcy risk: the 25th percentile of the log wage distribution, the 10th percentile of the log wage distribution, average log earnings, and the unemployment rate. In all cases, higher income or employment is associated with fewer consumer bankruptcies. But the coefficient on Medicaid eligibility remains statistically significant and changes little in magnitude. Column 6 presents results that include a control for counts of business bankruptcies. Business bankruptcies appear to be weakly predictive of consumer bankruptcies, but also have little effect on the precision of the Medicaid eligibility estimates.¹⁷

A final concern with our baseline specification is that the effect of Medicaid may involve complex adjustment dynamics, which would ultimately cause the short-run effect of Medicaid expansions to differ from the long-run effect. For example, if bankruptcy rates require several years to adjust to changes in public insurance then the previous fixed-effects regressions would not capture the full, long-run effect. Table 5 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. 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.

¹⁷The state-specific exemption levels for consumer bankruptcy change infrequently. Nevertheless, we obtained records on state bankruptcy exemption levels for a subset of our sample: 1995 through 2004 based on the work of Hynes et al. (2004). For this sub-sample of years, our baseline IV point estimate is less precisely measured than in our full sample: -0.637 with a standard error of 0.434. When we control for the state personal property exemption for married couples, the estimate changes little: -0.609 with a standard error of 0.429.

As a final empirical exercise, we explore one mechanism that might drive the interaction between Medicaid and bankruptcy. Medicaid reduces the probability that its enrollees will acquire medical debt, and this effect of Medicaid could drive the interaction with the bankruptcy system. We explore this mechanism directly by studying the effect of Medicaid on debts accrued to local health care providers. We do so with data on uncompensated care from the American Hospital Association (AHA) annual census of U.S. hospitals.¹⁸ For 1994 through 1999, we observe the total amount of bad debt (unpaid medical bills) and charity care (care for which the patient was not billed). Unfortunately, only five years of these data are available.

Table 6 estimates the effect of Medicaid eligibility on aggregate hospital bad debt and charity care using the state-level AHA data. Column 1 reports results of our baseline specification using only the years for which we have AHA data. The results are similar to the baseline specification for our full sample. Columns 2 through 4 replace the dependent variable with total hospital bad debt, total hospital charity care, and total uncompensated care (the sum of bad debt and charity care). Although none of the results in these columns are significant at conventional levels, the point estimates are uniformly negative and the magnitudes are economically large. For instance, in column 4, the coefficient on Medicaid eligibility suggests that a 10 percentage point increase in Medicaid eligibility reduces hospital uncompensated care by 5.2 percent. Hospitals provide 40 billion dollars of uncompensated care each year (Hadley and Holohan, 2004); thus a 5.2 percent reduction suggests that Medicaid transfers a large amount of money back to hospitals. We interpret these uncompensated care results as suggestive and broadly consistent with our baseline results.¹⁹

¹⁸We are grateful to Damon Seils and Kevin Schulman for assistance with these data.

¹⁹The results also suggest that the incidence of Medicaid expansions may fall at least partially on the hospitals themselves.

6 The Share of Bankruptcies Driven by Medical Costs

Researchers have found that medical costs are pivotal in between 17 to 54 percent of bankruptcies (Himmelstein et al. 2005; Dranove and Millenson 2006), depending on the definition of qualifying medical costs. This section offers 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, $\neg I$:

$$P(B) = P(B|I)P(I) + P(B|\neg I)P(\neg I). \quad (3)$$

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, $\hat{\beta} \times P(B)$.²⁰ This leads to the following equation:

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

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 $P(B) = 0.025$ based on our aggregate bankruptcy statistics and $P(I) = 0.70$ based on tabulations from the CPS.²¹ We use $\beta = 0.916$ based on our regression results.

²⁰It 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.

²¹Overall, roughly 1 percent of households file for bankruptcy in any given year, but bankruptcy risk is higher for low-income households; Warren (2003) suggests that bankruptcy risk is 2–3 times

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 less critical 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 medical costs on the consumer bankruptcy decision of the average family.

Bankruptcy filers are more likely to be drawn from the lower half of the income distribution, so Medicaid-eligible households are not atypical among bankruptcy filers. To confirm that this is the case, 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. Consistent with the work of Warren (2003), we find strong evidence that bankruptcy filers are more likely to be drawn from the lower half of the income

higher for low-income households. 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.

distribution. Figure 5 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 bankruptcy risk than other households.

7 Optimal Insurance Calibration

Last, we use our empirical results to estimate the elasticity of bankruptcy risk to health insurance benefits, $\frac{d \log p_B}{d \log b_H}$, in order to calibrate equation (1). We find that expanding Medicaid eligibility by 10 percentage points reduces consumer bankruptcies by 8.4 percent. We take this to be our estimate of $\frac{d \log p_B}{d b_H}$; assuming $b_H = 0.70$ as above implies that $\frac{d \log p_B}{d \log b_H} = -0.587$. We assume constant relative risk aversion, and we calibrate the formula for a range of commonly used risk aversion values. The details of the calibration are described in the appendix.

The results of the calibration are reported in table 7. When the bankruptcy system is ignored ($p_B = 0$), the optimal health insurance benefit rate is 64.9 percent, assuming a coefficient of relative risk aversion of 3. Accounting for the substitution elasticity estimated above, the optimal insurance benefit rate rises to 76.1 percent. In this way, the interaction between Medicaid and bankruptcy counteracts roughly one third of the moral hazard cost of expanding health insurance.

8 Conclusions

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, since 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.

Our estimates do not suggest that medical costs are responsible for the massive rise in consumer bankruptcies. From 1994 to 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.²² In reality, bankruptcies increased by 71 percent. Consequently, our estimates explain roughly 10 percent of the overall increase in bankruptcies. As pointed out by Livshits et al. (2007), Canada also experienced an enormous increase in consumer bankruptcies over the 1980s and 1990s. But, during that time period, Canadians enjoyed universal access to health insurance, suggesting that medical costs cannot explain the large increase in consumer bankruptcies. We conclude that medical costs are an important driver of bankruptcies, especially among low-income families, but that medical costs are unlikely to be the primary cause of the overall rise in consumer bankruptcies.

Taken as a whole, our results suggest that Medicaid affects not only its beneficiaries, 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.

²²This 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.

Appendix

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

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

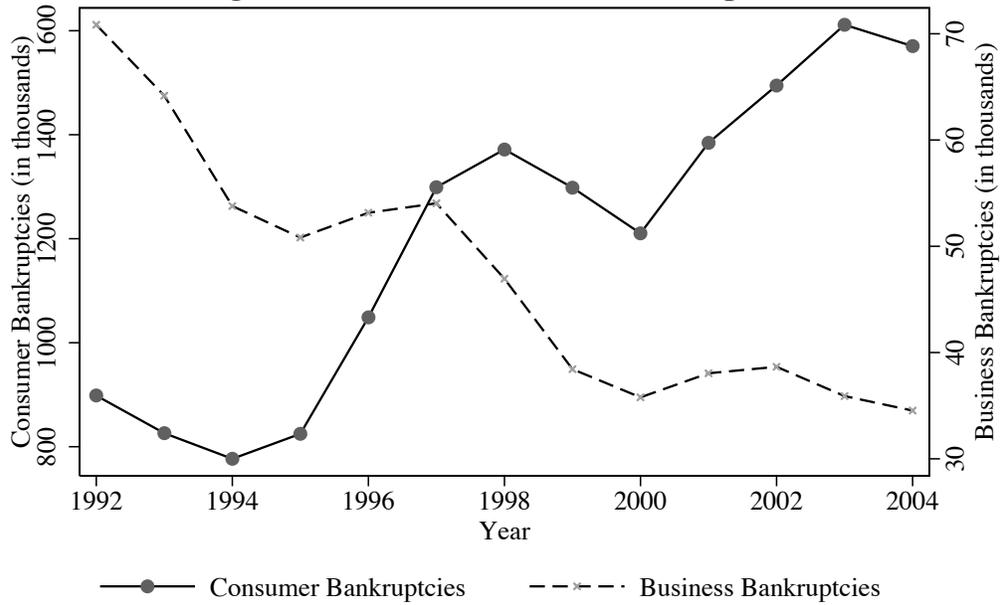
For the reasons described above, we choose $p_B = 0.02$ and $\frac{d \log m}{d \log b_H} = 0.2$. The terms $p_H = 0.5$, $m = 9000$, and $c = 25000$ match numbers used in calibrations by Finkelstein et al. (2008), based on calculations from the MEPS. The estimate of $D = 28000$ was chosen based on average debts discharged during bankruptcy as reported by Barron and Staten (1998). Finally, we assume that $b_B = 0.95$ to account for the costs of bankruptcy, mainly filing and legal fees. Given these assumptions and equation (5), calculating the optimal health insurance benefits (b_B), as reported in table 7, is straightforward.

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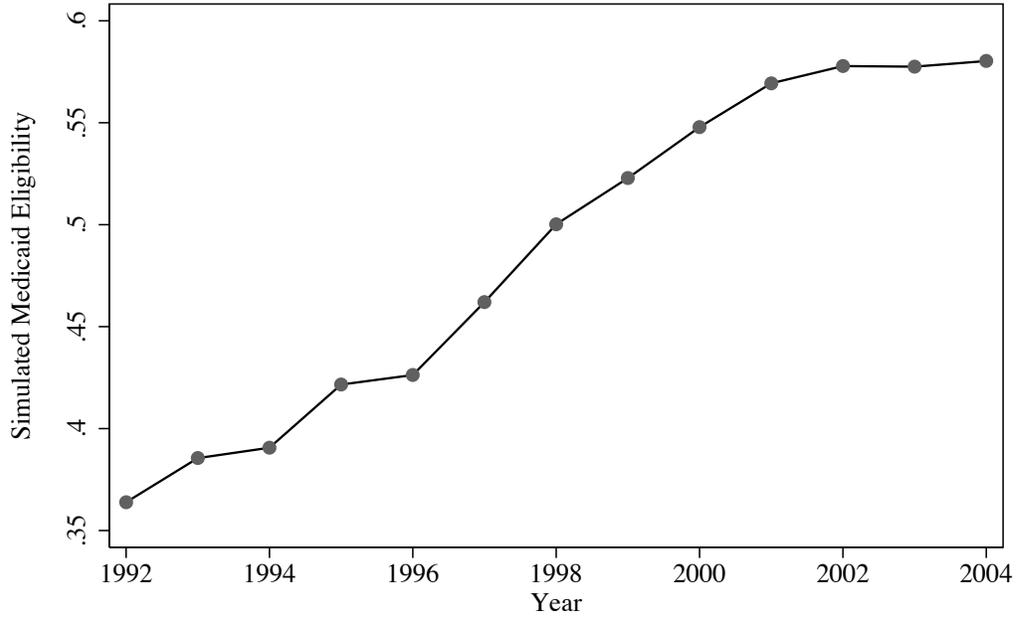
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Figure 1: National Trend in Bankruptcies



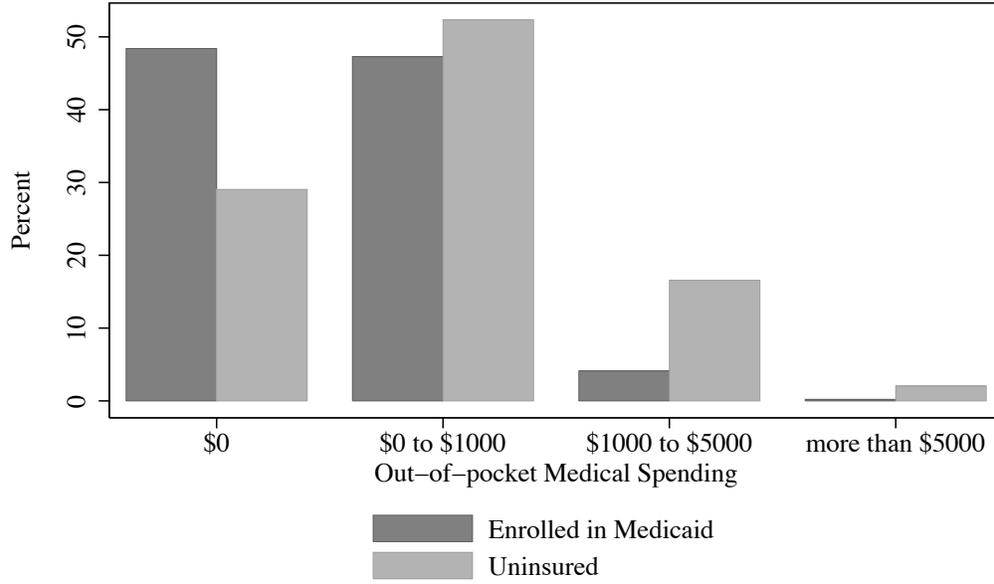
Source: Counts of bankruptcies from the annual census published by the Administrative Office of the US Courts

Figure 2: Growth in Medicaid Eligibility, 1992–2004



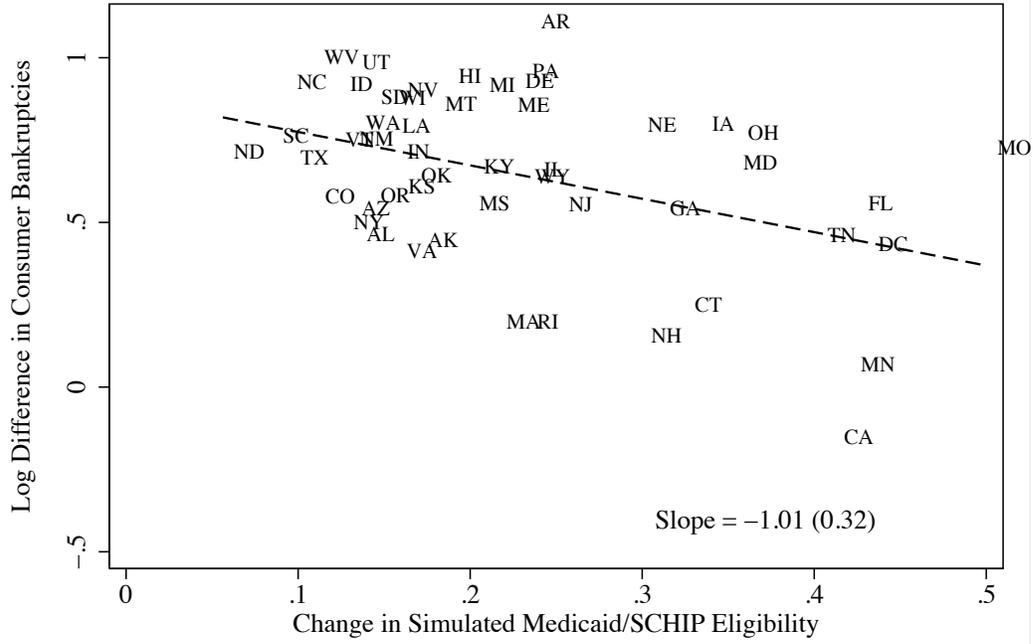
Source: Eligibility is based on our calculations from the March Current Population Survey. See text for details.

Figure 3: Out-of-Pocket Medical Spending, Uninsured versus Medicaid Recipients



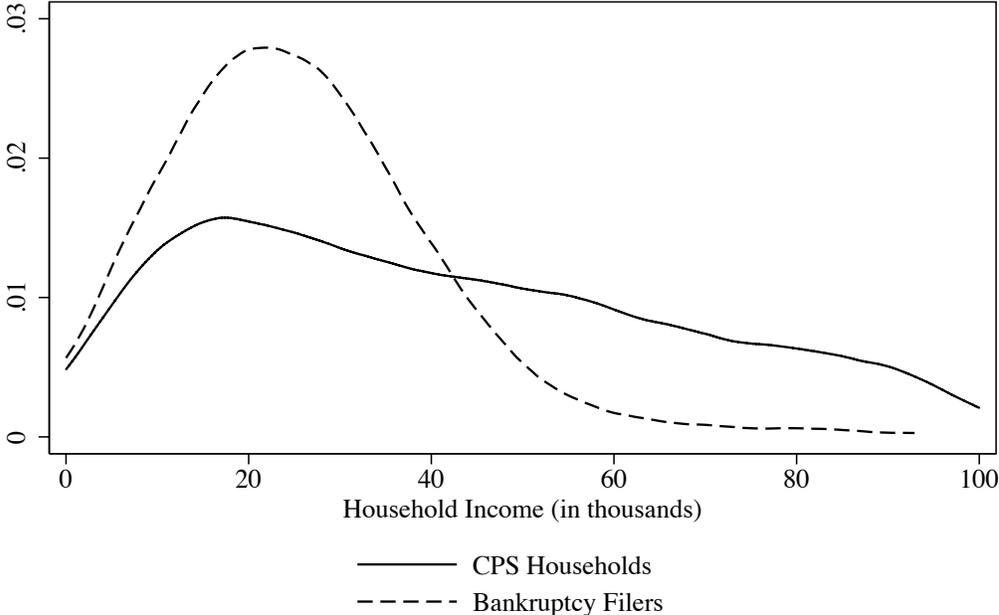
Source: Medical Expenditure Panel Survey. See text for details.

Figure 4: Bankruptcies and Medicaid Eligibility, 1992–2004



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

Figure 5: Estimated Density of Household Income



Source: CPS March Supplement and PACER Bankruptcy Records. See text for details.

Table 1: Summary Statistics

Year	Consumer Bankruptcies			Simulated Medicaid Eligibility		
	Median	Mean	Standard Deviation	Median	Mean	Standard Deviation
<u>A. All States</u>						
All Years	15,162	23,552	27,346	0.468	0.487	0.126
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	0.552	0.580	0.111
<u>B. Small Expansion States</u>						
All Years	11,175	21,027	23,658	0.444	0.438	0.068
1992	7,196	14,336	17,937	0.441	0.399	0.060
1996	8,886	17,976	21,824	0.453	0.411	0.061
2000	11,569	21,100	23,154	0.507	0.459	0.072
2004	15,449	30,872	35,795	0.529	0.490	0.065
<u>C. Large Expansion States</u>						
All Years	25,607	51,942	54,150	0.653	0.636	0.181
1992	18,020	46,320	55,615	0.339	0.381	0.062
1996	20,162	51,711	63,092	0.457	0.540	0.115
2000	25,607	49,773	55,567	0.802	0.748	0.130
2004	37,298	51,585	44,378	0.832	0.771	0.131

Notes: The sample consists of bankruptcy counts for the 50 states and DC from 1992-2004; 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

Dependent Variable:	(1)	(2)	(3)	(4)
	Consumer Bankruptcies		Business Bankruptcies	
Regression:	OLS	IV	OLS	IV
Medicaid Eligibility	-0.244 (0.295) [0.412]	-0.839 (0.356) [0.022]	0.391 (0.447) [0.386]	0.266 (0.600) [0.660]
R ²	0.99		0.93	
N	663		663	

Notes: The sample consists of bankruptcy counts for all 50 states and DC from 1992-2004; all observations are state-year. All dependent variables are in logs. 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 3: Alternative Specifications Involving Time Trends

Dependent Variable: Log Count of Consumer Bankruptcies							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	Baseline	Reduced Form with Lead and Lag	Region Trends	Region × Year Fixed Effects	State Trends	Sharp Expansion States	Sharp Expansion States w/ State Trends
Medicaid Eligibility	-0.839 (0.356) [0.022]		-0.819 (0.341) [0.020]	-0.855 (0.417) [0.046]	-0.341 (0.407) [0.406]	-1.023 (0.284) [0.002]	-0.886 (0.310) [0.011]
Simulated Medicaid Eligibility		-0.335 0.251 [0.189]					
Simulated Medicaid Eligibility, 2-year Lead		0.086 (0.190) [0.651]					
Simulated Medicaid Eligibility, 2-year Lag		-0.260 (0.161) [0.112]					
N	663	459	663	663	663	221	221

Notes: In all specifications except column (2), Medicaid eligibility is predicted with simulated Medicaid eligibility as an instrumental variable. The sample consists of bankruptcy counts for all 50 states and DC from 1992-2004; all observations are state-year. All specifications include state fixed effects and year fixed effects. The first column adds annual state unemployment rate and average household income to the baseline specification. 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 4: Alternative Specifications that Control for
Other Determinants of Consumer Bankruptcies

Dependent Variable: Long Difference in Log Count of Consumer Bankruptcies, 1992-2004							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Medicaid Eligibility	-0.839 (0.356) [0.022]	-0.735 (0.344) [0.038]	-0.783 (0.344) [0.027]	-0.676 (0.318) [0.039]	-0.738 (0.336) [0.033]	-0.861 (0.370) [0.024]	-0.631 (0.318) [0.052]
25th Percentile of Log Wage Distribution		-0.434 (0.163) [0.010]					-0.164 (0.140) [0.248]
10th Percentile of Log Wage Distribution			-0.213 (0.090) [0.022]				-0.057 (0.069) [0.416]
Average Log Earnings				-0.749 (0.239) [0.003]			-0.434 (0.241) [0.078]
Unemployment Rate					4.454 (1.528) [0.005]		3.365 (1.271) [0.011]
Log of Business Bankruptcies						0.084 (0.049) [0.094]	0.065 (0.044) [0.143]
N	663	663	663	663	663	663	663

Notes: 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 DC from 1992-2004; all observations are state-year. All specifications include state fixed effects and year fixed effects. The first column adds annual state unemployment rate and average household income to the baseline specification. 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 5: Short-run versus Long-run Effects

Dependent Variable: Log Count of Consumer Bankruptcies			
	(1)	(2)	(3)
	Baseline	3-Year Averages	1992, 1996, 2000, 2004 Panel
Medicaid Eligibility	-0.839 (0.356) [0.022]	-0.869 (0.364) [0.021]	-0.949 (0.475) [0.051]
N	663	204	204

Notes: 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 DC from 1992-2004; all observations are state-year. All specifications include state fixed effects and year fixed effects. All specifications include the state unemployment rate and the log of average earnings in the state as time-varying controls. 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: The Effect of Medicaid on Uncompensated Care

	(1)	(2)	(3)	(4)
Dependent Variable:	Consumer Bankruptcies	Bad Debt	Charity Care	Total Uncompensated Care
Medicaid Eligibility, M_{jt}	-0.901 (0.397) [0.028]	-0.137 (0.217) [0.531]	-0.940 (0.721) [0.199]	-0.515 (0.435) [0.232]
N	306	306	306	306

Notes: In all specifications Medicaid eligibility is predicted with simulated Medicaid eligibility as an instrumental variable. All dependent variables are in logs. 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 7: Optimal Insurance Calibration

	Coefficient of Relative Risk Aversion		
	2	3	4
Ignoring Bankruptcy System ($p_B = 0$)	49.1%	64.9%	73.3%
Including Bankruptcy System	67.1%	76.1%	81.3%

Notes: Results are from calibrating equation (1) and give the optimal health insurance benefit rate; see accompanying text and Appendix for more details on parameters used to calibrate the optimal insurance equation.