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ABSTRACT

This paper provides empirical evidence of Medicaid crowd out of demand for private long-term care insurance. Using data on the near- and young-elderly in the Health and Retirement Survey, our central estimate suggests that a $10,000 decrease in the level of assets an individual can keep while qualifying for Medicaid would increase private long-term care insurance coverage by 1.1 percentage points. These estimates imply that if every state in the country moved from their current Medicaid asset eligibility requirements to the most stringent Medicaid eligibility requirements allowed by federal law â€“ a change that would decrease average household assets protected by Medicaid by about $25,000 â€“ demand for private long-term care insurance would rise by 2.7 percentage points. While this represents a 30 percent increase in insurance coverage relative to the baseline ownership rate of 9.1 percent, it also indicates that the vast majority of households would still find it unattractive to purchase private insurance. We discuss reasons why, even with extremely stringent eligibility requirements, Medicaid may still exert a large crowd-out effect on demand for private insurance.

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Introduction

Expenditures on long-term care, such as home health care and nursing homes, accounted for 8.5 percent of all health care spending in the United States in 2004 (Congressional Budget Office, 2004). These long-term care expenditures are projected to triple in real terms over the next few decades, in large part due to the aging of the population (Congressional Budget Office, 1999). Because over one-third of Medicaid expenditures are already devoted to long-term care (U.S. Congress, 2004), there is rising concern among policy makers about the fiscal pressure that further growth in long-term care expenditures will place on federal and state budgets in the years to come, and growing interest in stimulating the market for private long-term care insurance. For example, in a much-publicized press release issued in October 2004, the National Governors Association announced that states spent nearly as much money on Medicaid in fiscal year 2003 as they did on K-12 education, and expressed concern that Medicaid is putting a “squeeze” on state budgets going forward (National Governors Association, 2004).

The market for private long-term care insurance is currently quite limited. Only about 10 percent of the elderly have private long-term care insurance (Brown and Finkelstein, 2004a). Because these policies tend to be quite limited in scope, only 4 percent of total long-term care expenditures are paid for by private insurance (Congressional Budget Office, 2004). By contrast, in the health care sector as a whole, 35 percent of expenditures are covered by private insurance (National Center for Health Statistics, 2002).

Medicaid provides public long-term care insurance in the form of a payer-of-last resort. It covers long-term care expenditures only after the individual has met asset and income eligibility tests, and after any private insurance policy held by the individual has paid any benefits it owes. In this paper we explore how changes in Medicaid’s means-tested eligibility thresholds might affect demand for private long-term care insurance.

We use data from the 1996, 1998 and 2000 waves of the Health and Retirement Survey to study the effect of Medicaid asset protection rules on private long-term care insurance coverage among individuals aged 55 to 69. To investigate Medicaid’s impact, we draw on the substantial variation across individuals in the amount of assets that can be protected from Medicaid based on their state of residence, marital status and asset holdings. Due to the potential endogeneity of asset holdings to these Medicaid rules, we predict assets based on demographic characteristics of the individual.

We find statistically significant evidence that more generous Medicaid asset protection is associated with lower levels of private long-term care insurance coverage. Our central estimate is that a $10,000 increase in the amount of assets an individual can protect from Medicaid is associated with a decrease in private long-term care insurance coverage of 1.1 percentage points. This implies, for example, that if all states were to adopt the most stringent asset eligibility requirements allowed by federal law in 2000 – $16,824 for a married couple and $2,000 for a single individual – and thereby decrease average protected assets by about $25,000**,** overall demand for private long-term care insurance would rise by 2.7 percentage points. While such an increase is large relative to the existing ownership rate in our sample of near-elderly and young elderly of 9.1 percent, it suggests that the vast majority of these individuals would remain uninsured.

Our empirical findings complement recent simulation-based estimates of the impact of Medicaid on private long-term care insurance demand (Brown and Finkelstein, 2004b). Like our empirical estimates, these simulation results also suggest that changes in Medicaid's asset disregards are unlikely to have a substantial effect on private long-term care insurance demand. At the same time, however, Brown and Finkelstein (2004b) estimate that Medicaid may be able to explain the lack of private insurance purchases for at least two-thirds of the wealth distribution, even if there were no other factors limiting the size of the market. This is because Medicaid imposes a substantial implicit tax on private long-term care insurance; for example, they estimate that about 60 to 75 percent of the expected present discounted value benefits that a median wealth individual would receive from a typical private long-term care insurance policy are redundant of benefits that Medicaid would have provided had the individual not purchased private insurance. Changes in Medicaid's asset disregards, however, do not have a large effect on this implicit tax. Together, the empirical and simulation results underscore the importance of understanding the mechanism behind the crowd out effect of a particular public program in considering the likely impact of potential reforms to the public program on private demand.

The rest of the paper proceeds as follows. Section one provides background information on long-term care expenditure risk and the nature of existing public and private insurance coverage for this risk. It also briefly reviews the insights from simulation estimates of how Medicaid affects private long-term care insurance demand. Section two presents the data and empirical framework. Section three presents our crowd-out estimates. Section four uses these crowd-out estimates to simulate the likely effects of changes in Medicaid means-testing thresholds. Section five concludes.

1. Background on long-term care insurance and Medicaid crowd out

Long-term care represents a significant source of financial uncertainty for elderly households. Although most 65 year olds will never enter a nursing home, of those who do enter a nursing home, 12 percent of men and 22 percent of women will spend more than 3 years there; one-in-eight women who enter a nursing home will spend more than 5 years there (Brown and Finkelstein, 2004b). These stays are costly. On average, a year in a nursing home costs $50,000 in 2002 for a semi-private room, and even more for a private room (MetLife, 2002).

Very little of this expenditure risk is covered by private insurance. According to the 2000 Health and Retirement Survey, among those individuals aged 60 and over, only 10.5 percent own private long-term care insurance. Moreover, Brown and Finkelstein (2004a) estimate that the typical purchased policy covers only about one-third of EPDV long term care expenditures. As a result, only about 4 percent of long-term care expenditures are paid for by private insurance, while about one-third are paid for out of pocket (Congressional Budget Office, 2004); by contrast in the health sector as a whole, private insurance pays for 35 percent of expenditures and only 17 percent are paid for out of pocket (National Center for Health Statistics, 2002). Medicaid pays for about 35 percent of long-term care expenditures (Congressional Budget Office, 2004).[1]

[1] This leaves a remaining one-quarter of expenses that are covered by Medicare. However, this apparently large Medicare share is somewhat misleading. About half of Medicare long-term care spending consists of Medicare's home health care benefit, which is a genuine long-term care service. However, the other half comes from Medicare's coverage of short-term, skilled nursing home facilities following an acute hospital stay; this is not the custodial nursing home care that accounts for the vast majority of nursing home days and is covered by private long-term care

An extensive theoretical literature has proposed a host of potential explanations for the limited size of the private long-term care insurance market. These explanations include both factors that constrain supply and factors that limit demand. Norton (2000) provides a useful overview of the various potential explanations.

On the supply side, market function may be impaired by such problems as high transactions costs, imperfect competition, asymmetric information, or dynamic problems with long-term contracting. There is evidence consistent with the existence of many of these supply-side failures in the private long-term care insurance market. Finkelstein and McGarry (2006) provide evidence of asymmetric information in the market. There is also evidence of dynamic contracting problems arising both from the difficulty of insuring the aggregate risk of rising medical costs (Cutler, 1996) and from dynamic adverse selection as individuals who learn that they are better risks than expected drop out of the market (Finkelstein et al., 2005). Brown and Finkelstein (2004a) present evidence that premiums are marked up about 18 cents per dollar of premium above actuarially fair levels; this markup appears to reflect a combination of transaction costs and imperfect competition.

On the demand side, several different factors that may constrain the private insurance market have been suggested. Limited consumer rationality – such as difficulty understanding low-probability high-loss events (Kunreuther, 1978) or misconceptions about the extent of public health insurance coverage for long-term care – may play a role. Demand may also be limited by the availability of imperfect but cheaper substitutes, such as financial transfers from children, unpaid care provided directly by family members in lieu of formal paid care, or the public insurance provided by the means-tested Medicaid program (Pauly, 1990; Brown and Finkelstein, 2004b).

There is evidence that these demand side factors are likely to be important in understanding the limited size of the private market. Brown and Finkelstein (2004a) suggest that the loads on policies – and whatever market failures produce them – are unlikely to be sufficient to explain the limited market size.

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insurance and by Medicaid, and is somewhat misleadingly included in long term care spending estimates. (Congressional Budget Office 2004, US Congress, 2000).

They note that the average load on a typical private policy is about 50 cents on the dollar higher for men than women, yet ownership patterns are extremely similar by gender, a fact that cannot be explained solely by the within-household correlation in ownership patterns. This suggests an important role for demand side factors such as Medicaid.

Brown and Finkelstein (2004b) provide more direct evidence of a crowd out effect of Medicaid. They develop and calibrate a utility-based model of an elderly, life cycle consumer's demand for private long-term care insurance and compare demand under various counterfactual assumptions regarding the nature of private insurance and of the Medicaid program. Their simulations suggest that given the current structure of Medicaid, even if actuarially fair, comprehensive private insurance policies were to be available, at least two-thirds of the wealth distribution would still not purchase this insurance. They show that the mechanism behind this large estimated Medicaid crowd out effect stems from the fact that a large portion of private insurance benefits are redundant of benefits that Medicaid would have provided in the absence of private insurance, a phenomenon that they label the Medicaid "implicit tax". For a male (female) at the median of the wealth distribution, they estimate that 60 percent (75 percent) of the benefits from a private policy are redundant of benefits that Medicaid would otherwise have paid.

The Medicaid implicit tax stems from two features of Medicaid's design that results in private insurance reducing expected Medicaid expenditures. First, by protecting assets against negative expenditure shocks, private insurance reduces the likelihood that an individual will meet Medicaid's asset-eligibility requirement. Second, Medicaid is a secondary payer when the individual has private insurance. This secondary payer status means that if an individual has private insurance, the private policy pays first, even if the individual's asset and income levels make him otherwise eligible for Medicaid; Medicaid then covers any expenditures not reimbursed by the private policy.[2]

[2] Understanding Medicaid's implicit tax also helps explain the ostensibly puzzling finding that men and women purchase private insurance in very similar proportions, despite substantially higher loads on male policies. Since women have much higher expected lifetime long-term care utilization, the expected proportion of long-term care expenditures paid for by Medicaid is higher for women than men of the same asset levels, and thus the Medicaid implicit tax on private insurance is higher for women than for men. Indeed, Brown and Finkelstein (2004b) show

Brown and Finkelstein (2004b) estimate that changes in Medicaid's asset disregards would not have a substantial effect on the Medicaid implicit tax, and thus, would not make private long term care insurance desirable for most of the wealth distribution. Specifically, they simulate the likely effect of a policy that has been adopted in several states which makes the Medicaid asset disregards less stringent if the individual purchases private insurance. They estimate that such a policy would not however, have much effect on the implicit tax or on private insurance demand because, even in the absence of any asset eligibility requirements – i.e. complete asset protection for individuals – Medicaid still imposes a substantial implicit tax on private insurance through its status as a secondary payer.

This paper extends the simulation analysis in Brown and Finkelstein (2004b) to examine empirically how the amount of assets that Medicaid allows an individual to keep while receiving Medicaid coverage for long term care expenses affects demand for private long-term care insurance. Our empirical estimates of the crowd-out effect of Medicaid on private long-term care insurance demand are also related to a sizeable empirical literature that has investigated the extent of Medicaid's crowd out of acute private health insurance among working families. The estimates from this literature range in magnitude, but at the upper end suggest that up to half of the increase in public insurance coverage from increased Medicaid eligibility is offset by reductions in private insurance coverage (see Gruber, 2003 for a review of this literature).

To our knowledge, only two other empirical papers have examined the impact of Medicaid on private long term care insurance demand. Sloan and Norton (1997) compare private long-term care insurance holdings in the 1992 and 1994 HRS and the 1993 AHEAD across individuals in states with different Medicaid income eligibility limits. They find evidence that higher Medicaid income eligibility limits are associated with lower probability of owning long-term care insurance in the AHEAD data (ages 70+) but not in the HRS data (ages 51 – 64); they do not examine the effect of asset limits. Kang et al (2004) use the 1992 through 1998 waves of the HRS to examine the effect of Medicaid asset and income tests on

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that the *net* loads on polices – i.e. the load on the net benefits from the private policy, which omits any benefits paid by the private policy that Medicaid would otherwise have paid – are quite similar for men and women.

private insurance coverage, using variation in individual financial resources and state Medicaid eligibility limits. They find evidence consistent with a crowd out effect of less stringent Medicaid asset eligibility limits, but not evidence of an effect of Medicaid income limits on long-term care insurance coverage.

Our paper builds on this earlier work in two important dimensions. First, we limit our attention to data from 1996 and later waves of the HRS since prior survey waves utilized a confusing question to ascertain long-term care insurance coverage, resulting in substantial under-reporting (coverage rates are about one-fifth of what other surveys from that time suggest) and, more generally, extremely poor data quality; see Finkelstein and McGarry (2006) (Appendix A) for more details on these data issues. Second, both papers utilize differences in state Medicaid rules to identify the impact of Medicaid on long-term care insurance demand; however, there are other potentially important determinants of the demand for long-term care insurance that vary by state, such as the price and quality of nursing homes. Our empirical approach allows us to surmount this concern, as we discuss in more detail below.

2. Data and Empirical Approach

2.1 Data and Summary Statistics on Long-Term Care Insurance Coverage

We use data from the Health and Retirement Survey (HRS), a nationally representative sample of the elderly and near-elderly. We use a restricted access version of the HRS that allows us to identify the individual's state of residence. Our analysis uses data from the 1996, 1998 and 2000 waves of the HRS. The 1996 wave consists exclusively of individuals from the original HRS cohort (individuals born 1931 to 1941). The 1998 and 2000 waves also include individuals from the adjacent, younger cohort (born 1942 to 1947), and the adjacent older cohort (born 1924 – 1930); these are known respectively as the "War Baby" cohort and the "Children of the Depression" (CODA) cohort. We limit the analysis to individuals aged 55 to 69 in each wave. As discussed, we do not use data from waves prior to 1995 due to data issues with the measurement of long-term care insurance coverage; we exclude the 1995 AHEAD wave because individuals in this wave are outside our age range.

We limit our analysis to individuals aged 55 to 69 to focus on the decisions of individuals who are in the prime buying ages for long-term care insurance (HIAA, 2000). Once purchased, the policy is intended to be a “lifetime” policy; indeed, subsequent annual premiums are constant in nominal terms, so that policy payments are quite front-loaded. As a result, it is important to examine the effect of Medicaid rules that were in effect when an individual might be considering the purchase of private long-term care insurance. For this reason, we particularly wish to exclude individuals aged 70 and over from the analysis. Such individuals may well have been making their purchase decisions in the mid to late 1980s, during which Medicaid eligibility rules were substantially different than they are today. Crucially for our empirical strategy, which relies on the differential treatment of married and single individuals within different states, these rules would not have varied within state by marital status prior to 1989.[3] The current structure of Medicaid eligibility rules was adopted with the Medicare Catastrophic Coverage Act of 1988, which was implemented in 1989 (Stone, 2002).

Because of the panel nature of the data, we observe many individuals multiple times over the waves. Our full sample consists of 28,100 observations on 12,402 unique individuals. We account for the multiple observations of the same individuals in the error structure in our regression analysis. We do not, however, directly exploit the panel nature of the data and the changes in Medicaid eligibility rules for specific individuals over time due to changes in martial status or – more commonly – changes in state rules. We believe the use of such changes provide a questionable form of identification since it is unclear under which set of rules the individual made the (lifetime) purchase of long-term care insurance. Indeed, as we discuss in more detail below, our preferred specification limits the analysis to the sub-sample of individuals who did not change marital status between 1996 and 2000 and who live in one of the 30 states which have not had any real changes to their Medicaid asset allowances between 1991 and 2000 (see Appendix A for details). We refer to this sub-sample as the “Constant Medicaid rules sub-sample” because the individuals faced constant Medicaid rules over our time period. They represent an arguably cleaner sample on which to analyze the crowd-out effects of Medicaid as there is considerably less

[3] Consistent with this, using our empirical strategy we find statistically insignificant effects of current (1996 – 2000) Medicaid rules on long-term care insurance coverage for individuals who are 70 and older (mean age of 79) and who therefore may have been at the prime buying age under a very different set of rules (results not reported). We also show in the sensitivity analysis below that the crowd-out effects we estimate in our 55-69 year old sample are stronger at younger ages within this range.

uncertainty about what rules were in effect when the individuals bought (or considered buying) long-term care insurance.

Table 1 presents some summary statistics for both the full sample (column 1) and the constant Medicaid rules sub-sample (column 2). All statistics are based on using household weights. We focus on the summary statistics for the sub-sample in column 2, although the results are generally similar. The long-term care insurance coverage rate is 9.1 percent. This is comparable to the rates found in other surveys for similar age ranges (see e.g. HIAA, 2000). Just over 70 percent of the sample is married, just under half is male, and about two-fifths are retired.

The average long-term care insurance coverage rate masks important variation across sub-groups in their long-term care insurance holdings. Table 2 therefore presents summary statistics on long-term care ownership rates separately by various covariates. Once again, column 1 presents the results for the full sample, and column 2 presents results for the constant Medicaid rules sub-sample. Coverage rates are similar by gender, and higher for married individuals than single individuals (10.0 percent vs. 7.1 percent). Coverage rates are higher among 62 to 69 year olds (10.4 percent) than among 55 to 61 year olds (8.1 percent). Coverage rates also vary across states; the inter-quartile range in long-term care insurance coverage rates across states ranges from 0.06 to 0.12 (not shown).

The pattern of coverage by net worth is most dramatic. Less than 4 percent of the sample in the bottom quartile owns long-term care insurance, compared to 15 percent in the highest quartile of net worth. In fact, long-term care insurance coverage rates increase monotonically by wealth decile, from 0.03 percent in the bottom decile to 0.17 percent in the top. The wealth profile likely reflects the fact that the means-tested eligibility requirements of Medicaid make it a better substitute for private insurance for lower wealth individuals.

2.2 Overview of Medicaid rules and our empirical approach

We focus our analysis on the impact on private long-term care insurance demand of the amount of protected financial assets that an individual can keep while still receiving Medicaid reimbursement for long-term care utilization. Below, we show that other Medicaid rules such as the minimum allowable income retention for the community spouse or the treatment of the community house's spouse upon its sale or her death do not appear to affect insurance coverage, and do not affect our estimate of the effect of the asset rules on insurance coverage.

Medicaid financial asset disregards exhibit substantial variation across individuals based on an individual's marital status, state of residence, and asset holdings. Our empirical strategy, broadly speaking, is to control for any direct effects of marital status, state, and assets holdings on long-term care insurance demand, and then to identify the impact of Medicaid on long-term care insurance demand using the variation in Medicaid generosity that exists across higher interactions of these three variables (i.e. assets by marital status, state by marital status, and assets by state, as well as assets by state by marital status). Thus for example, we use both differences across states in the amount of assets protected for married individuals relative to single individuals, and differences across states in the amount of assets protected for individuals of different asset levels to identify the impact of Medicaid's asset protection rules on demand for private long-term care insurance. Throughout the analysis, we use predicted assets to deal with the potential endogeneity of assets to Medicaid spend down rules (Coe, 2005).

Medicaid asset rules for single individuals are relatively simple and uniform across states and particularly within states: they do not vary with the assets of the individual (as long as the individual has assets of more than the protected amount). The modal state rule (used by nearly 70 percent of states) allowed single individuals receiving Medicaid coverage for nursing home care to retain no more than $2,000 in financial wealth. The remaining states had asset limits ranging from $1,500 to $6,500.

In contrast, the amount of assets a community spouse is allowed to keep when her spouse goes into a nursing home exhibits substantial variation across states at a given household asset level, from a minimum of $16,824 to a maximum of $84,120 in 2000. Moreover, the amount of assets a community spouse can keep varies with household assets, and this difference across states is highly non-monotonic in the level of household assets.

For married households with assets below the minimum amount that federal law requires be kept when one spouse is in a nursing home ($16,824 in 2000), there is no difference across states in Medicaid asset disregards. For most states, there is also no difference in the amount of assets the married couple can keep if their assets are more than double the maximum amount that federal law allows to be kept when one spouse is in a nursing home (which puts the asset amount at $168,240 in 2000). However, for married households within this range – which corresponds to roughly the 20th to 60th percentile of the asset distribution for married households in the relevant age range in the 2000 HRS– there are substantial differences across states in the amount of assets that a married household can keep under Medicaid.

By way of illustration, Figure 1graphs the difference in the amount of assets a community spouse can keep as a function of total household financial assets in the two most common sets of state rules. Under the most common set of rules – which is used in 26 states – the community spouse is allowed to keep all of their assets up to the federally allowed maximum protected assets ($84,120 in 2000) after which they face a 100% marginal tax rate on all further assets. In the second most common set of rules – which is used in another 15 states – the community spouse is allowed to keep all of her assets up to the federally allowed minimum protected assets ($16,824 in 2000), faces a 100 percent marginal tax rates on all assets between this federal minimum and two times the minimum, faces a 50 percent marginal tax rate on all assets between twice the federal minimum and twice the federal maximum, and a 100 percent marginal tax rate on all assets above twice the federal maximum.

As seen in Figure 1, the difference in amount of protected assets that a community spouse with a given amount of assets faces varies non-monotonically with assets. Using the asset distribution for married households in our age range in the 2000 HRS, we estimate that moving from the most common set of state rules to the next most common would on average allow a married household to keep $21,715 more in assets when one spouse entered a nursing home, which represents 29 percent of average financial assets in this range. The maximum difference in the amount of assets that a household would be able to keep is $42,060 and occurs for household with assets of $84,120. The minimum difference in protected assets is 0, and occurs for household with assets of less than $16,824 or more than $168,240. If we were instead compared these most common state rules (which are also the most generous in terms of the amount of protected assets allowed for married couples) to the least generous state rules (used by 3 states) the maximum difference in the amount of assets the household would be able to keep would rise to $67,296 (which would occur in households with $84,120 or more in assets).

To sum up, we exploit several key sources of variation in the amount of protected assets to identify the impact of Medicaid asset protection on demand for private long term care insurance. These include: differences across states in the average asset disregards for married and single individuals, differences across married individuals of different asset levels in different states, and differences across married and single individuals of different asset amounts, as well as higher order interactions between state of residence, marital status, and assets. In all cases, we control for any direct effects of asset levels, marital status, or state of residence on the probability an individual has private long term care insurance. For interested readers, Appendix A provides considerably more detailed information on how the Medicaid eligibility rules vary across states by marital status and asset level.

2.3 Econometric Framework

Temporarily ignoring several econometric concerns (that we will address below), a natural starting point would be to estimate the following equation:

$$\mathrm{LTCI}_{\mathrm{ist}} = \beta_1 \mathrm{Protected}_{\mathrm{ist}} + \beta_2 \mathrm{Married}_{\mathrm{ist}} + \alpha_s + X'_{ist}\eta + \varepsilon_{ist} \qquad (1)$$

In this estimating equation, the dependent variable $\mathrm{LTCI}_{\mathrm{ist}}$ is a binary indicator for whether individual *i* in state *s* and year *t* owns long-term care insurance, $\mathrm{Married}_{\mathrm{ist}}$ is an indicator variable for whether the individual is married and α_s represents a full set of state fixed effects.[4] The vector of covariates (X) consist of indicator variables for education categorized by highest degree achieved (less than high school, high school, some college, college degree or more), gender, occupation, industry, number of children up to 5, Hispanic heritage, race, retired, age, wave, and cohort; in addition; in addition, (X) includes interactions of each of the education categories with all of the other control variables.

[4] We specify equation (1) as a linear probability model because it allows us to handle instrumental variables most flexibly; as we discuss in more detail below, we are concerned about endogeneity of the right hand side variable *Protected* and therefore estimate equation (1) by instrumental variables. We have confirmed, however, that the marginal effects from Probit specifications evaluated at the mean yield nearly identical results to the linear probability model specified in equation (1).

The main covariate of interest is *Protected*, which we measure in units of $10,000. *Protected* measures the amount of financial assets that a particular household is allowed to keep and still qualify for Medicaid reimbursement. A higher level of *Protected* corresponds to a more generous (less means tested) Medicaid program. The mean (median) amount of *Protected* assets in our sample is $36,345 ($18,152) with a standard deviation of $36,135.

Protected varies across households depending on state of residence (s), marital status (m), and household assets according to the following formula:

$$Protected_{ims} = \begin{cases} \text{Assets}_{ims} & \textit{if} \quad \text{Assets}_{ims} \leq \text{Minimum}_{ms} \\ \text{Minimum}_{ms} + .5 * (\text{Assets}_{ims} - \text{Minimum}_{ms}) & \textit{if} \quad \text{Minimum}_{ms} < \text{Assets}_{ims} < \text{Maximum}_{ms} \\ \text{Maximum}_{ms} & \textit{if} \quad \text{Assets}_{ims} \geq \text{Maximum}_{ms} \end{cases} \quad (2)$$

The state sets the level for the minimum and maximum amount of assets protected by the Medicaid program, within the constraints imposed by Federal law. We calculate $Protected_{ims}$ for each individual in our sample, based on their assets, marital status, and the specific state Medicaid rules detailed in Appendix A.

By including a dummy for whether the individual is married and a full set of state fixed effects we control, respectively, for any fixed differences across married and single individuals or across individuals in different states in their demand for private long term care insurance. The covariates (X) are designed to control for demographics that may directly affect insurance demand, perhaps through their effect on asset levels or perhaps through other means. (We do not directly control for assets in equation (1) because of its potential endogeneity, although we have verified that controlling flexibly for net worth decile does not in fact affect the results). *Protected* is therefore identified off of two-way and three-way interactions between state, marital status, and assets.

We note that the state fixed effects allow us to control flexibly for a number of other potentially important determinants of demand for private long-term care insurance. They condition out any differences across states in the price and quality of nursing homes, which may affect demand for long-term care insurance. They also condition out any differences across states in the Medicaid program that may influence insurance coverage but are the same for married and single individuals within a state or individuals of different asset levels within a state. These include, for example, the Medicaid rules regarding the nature and extent of coverage provided for home health care, and the Medicaid reimbursement rates relative to private payer rates in the state. Our estimates therefore focus precisely on the impact of Medicaid eligibility rules for nursing home coverage on long-term care insurance demand.

A potential concern with estimating equation (1) is that – as equation (2) makes clear – *Protected* is a function of assets, and therefore savings decisions, which may themselves be affected by Medicaid rules. Thus assets may be endogenous to insurance purchase decisions. Indeed, there is empirical evidence that the savings of the elderly appear to respond to the incentives embodied in Medicaid's rules for eligibility for coverage for long-term care expenditures (Coe, 2005). This is consistent, more generally, with the evidence that savings decisions are affected by the incentives provided by means tested public insurance programs (see e.g. Hubbard, Skinner and Zeldes, 1995 and Gruber and Yelowitz, 1999).

To address the potential endogeneity of assets to Medicaid rules, we calculate predicted assets for each household based on a reduced form prediction model that uses only plausibly exogenous demographic characteristics to predict asset accumulation. Specifically, we estimate:

$$Log(Assets)_{\text{ist}} = X'_{ist}\delta + \upsilon_{ist} \qquad (3)$$

We estimate the asset equation in logs because the highly skewed nature of the asset distribution results in a much better fit in predicting log assets than assets. We define assets to be Medicaid-taxable assets; these are the same as net worth for single individuals, but exclude housing wealth from net worth for married individuals, since housing wealth is not treated as a Medicaid-taxable asset for married individuals. As covariates we include the same set of covariates used in X in equation (1) that we described above. We also include a marital status dummy since savings behavior may well differ across single and married individuals. Note that we do not use state dummies – or state Medicaid rules – in predicting wealth. The goal of equation (3) is not to develop the best prediction model of assets but to isolate the portion of assets that can be explained by plausibly exogenous demographic characteristics rather than asset accumulation decisions that are themselves endogenous to the state Medicaid rules. We estimate equation (3) using the full data sample, and household weights. Estimation of the prediction equation (3) yields an R-squared of 0.24.

Using the results of equation (3), we generate predicted assets for each individual in the sample. We then use predicted assets – instead of actual assets – as well as the individual's state of resident and marital status to calculate the amount of assets that would be protected by the Medicaid program. We refer to these protected assets calculated using predicted rather than actual assets as *Protected_Hat.* Thus, *Protected_Hat* represents the amount of assets the Medicaid program would disregard if the household's actual assets were as predicted by their characteristics. By contrast, *Protected* denotes the amount of assets the Medicaid program would protect based on their actual (potentially endogenous) assets. Like *Protected*, *Protected_Hat* is measured in units of $10,000. The mean (median) value of *Protected_Hat* in our sample is $43,121 ($39,929), with a standard deviation of $34,348.

In the results reported below, we estimate equation (1) by instrumental variables, instrumenting for *Protected* with *Protected_Hat*. In all of our regression estimates we use the HRS household weights. We adjust the standard errors to allow for an arbitrary variance-covariance matrix in the error term within each state. To take account of the sampling variation in the predicted variable *Protected_Hat* (Murphy and Topel, 1985) we also report standard errors from a non-parametric bootstrap. Specifically, we bootstrap the prediction equation (equation 2) and for each iteration of the bootstrap, calculate predicted assets, use these to calculate *Protected_Hat*, and then estimate equation (1) using *Protected_Hat* as an instrument for *Protected* on the drawn sample; we run 200 iterations of the bootstrap. In practice, the standard errors are not affected much by this procedure; we report both sets of standard errors in the results below.

Because we are using multiple waves of the HRS, we calculate *Protected* using the state rules and individual demographics in effect in the year in which the interview takes place. As mentioned above, 21 states, affecting 32 percent of the sample, experience real changes in the community spouse asset disregards between 1991 and 2000. In addition, about 5 percent of our sample changes marital status over the waves 1996 – 2000 In principle, these changes in state rules and marital status over time provide us with a fourth source of variation in Medicaid asset protection rules faced by an individual. We do not, however, believe that such changes in Medicaid asset protection are a particularly clean or useful source of variation, as it is unclear for these individuals which Medicaid asset protection rules were in effect – and thus the relevant rules – when the individual was considering whether to purchase long-term care insurance. Although we report estimation results for the full sample, our preferred specification limits the sample to the approximately three-fifths of the original sample (17,623 observations consisting of 7,923 unique individuals) who did not change martial status between 1996 and 2000 and are from states whose Medicaid rules did not change in real terms since 1991. Our estimates of crowd-out become larger and more precise in this sub-sample, which is consistent with greater measurement error in the full sample in the relevant Medicaid rules in effect when an individual is making his long term care insurance coverage decision.

Finally, it is worth noting a potential limitation to our approach is that we are using current predicted assets, while what matters for the Medicaid asset tax is the assets an individual has at the time of nursing home entry. This will bias against finding an effect of Medicaid. In practice, however, the relatively low rates at which the elderly appear to spend down their assets over their retirements suggest that this may not be too great of a problem (see e.g. Hurd 1989, Hurd 2002, and Mitchell and Moore 1997). .

3. Crowd-out estimates

Table 3 reports the main results from estimating equation (1) by instrumental variables, using *Protected_Hat* to instrument for *Protected*. The first column shows the results for the whole sample. The coefficient on *Protected* is -0.0056, and is statistically significant at the 10 percent level. The point estimate suggests that a $10,000 increase in the amount of assets an individual can retain while qualifying for Medicaid is associated with a 0.56 percentage point decline in long-term care insurance coverage.

The remaining columns report analysis when the sample is limited to individuals who face constant Medicaid rules. While we lose almost two-fifths of our observations due to these data cuts, we believe this sub-sample will provide a cleaner estimate of the impact of Medicaid on long-term care insurance coverage. Consistent with this view, column 2 indicates that the estimated effect of Medicaid on long-term care insurance demand is larger (and more statistically significant) in the constant Medicaid rules sub-sample than in the full sample. The point estimate on *Protected* rises to -0.109, and is statistically significant at the 5 percent level. This suggests that a $10,000 increase in the amount of assets a household can hold and still be eligible for Medicaid is associated with a 1.1 percentage point decline in the probability of holding long-term care insurance. The results in column 2 constitute our preferred specification, and we use these results for our central estimate.

The remaining columns of Table 3 explore the sensitivity of our central estimate to using different sources of variation to identify the effect of Medicaid protected asset rules on long-term care insurance demand. As discussed above, variation in *Protected_hat* comes from the two-way interaction of predicted assets with state, the two-way interaction of predicted assets with marital status, the two-way interaction of marital status with state, and the three way interaction of marital status, predicted assets, and state. To investigate whether each of these sources of variation yields similar results, columns (3) through (6) show the results in which we control one by one for various sources of variation, and therefore identify only off of the others. Specifically, in column (3) we add a control for predicted assets interacted with marital status, in column (4) we add controls for predicted assets interacted with state dummies, and in column (5) in which a we add controls for marital status interacted with state dummies. Finally, in column (6) we include controls for all two-way interactions (predicted assets by marital status, predicted assets by state, and married by state) so that the only variation used to identify *Protected_Hat* is the three-way interaction of state by marital status by predicted assets. Although the analysis often loses power when various sources of identifying variation are eliminated, the results indicate that the coefficient on *Protected* always remains negative and roughly of the same magnitude as the -0.011 in the baseline specification; it various from -0.0093 to -0.017 depending on the specification. The fact that all the sources of variation yield similar estimates increases our confidence in the empirical strategy and our baseline estimates.

Table 4 reports results from a number of additional sensitivity analyses. Column 1 replicates the IV estimates from our preferred specification (Table 3, column 2). One potential concern is that two other aspects of Medicaid vary across state by marital status and may also affect insurance demand: the treatment of income and estate recovery practices.[5]. Multi-collinearity in various Medicaid program rules' generosity could produce a misleading estimate of the impact of Medicaid asset rules. Moreover, the impact of these other features of Medicaid on long-term care insurance demand are of independent interest. Column 2 therefore adds two variables to control for these two features. The variable "Income" measures the amount of income (in units of $10,000) the household is allowed to keep and still qualify for Medicaid; this varies across states and within state by marital status. "Liens" is an indicator variable for whether a state will put a lien on a house when one spouse is in the nursing home in order to recoup expenses upon the death of the community spouse. This practice means that the house is no longer a bequeathable asset for married couples and the house is only a temporarily protected asset; there is no change for single households since the house is not a protected asset for them in any state. Appendix A describes the state income and housing ("liens") rules in more detail. The results in column 2 of Table 5 show the expected positive coefficient on "Liens", but the positive coefficient on "Income" is the opposite of what was expected. Neither coefficient is statistically significant, and an F-test indicates that they are not jointly significant (not shown). Perhaps most importantly, inclusion of these variables does little to change the parameter of interest, the coefficient on *Protected.*

As discussed previously, the variation in our variable of interest *Protected* occurs mostly in the range of 20^{th} to 60^{th} percentile of the asset distribution of married individuals (see Figure 1). Therefore, column 3 shows the results limiting the sample to this (albeit endogenous) range; as expected, the point estimate increases in absolute value. However, even with a doubling of the point estimate – to -0.0223 (standard error = 0.0130), the results still imply that even if all of the states decreased the amount of protected

[5] Of course, many other aspects of Medicaid vary across state – such as reimbursement rates for nursing homes and whether and how much coverage is provided for home care. An advantage of our strategy is that because we do not use cross-state differences in Medicaid to identify its effects, we purge these differences and are able to focus on the effect of one particular Medicaid parameter of interest.

assets to the minimum allowable under federal law in 2000, the vast majority of individuals in our sample would remain without private insurance.

Columns (4) and (5) reports the results from doing the analysis separately for younger ages (55-61) and older ages (62-69), respectively. The sample specification suggests the effect is stronger on younger ages – which may be because these individuals are more likely to be buying during the time of the analysis and thus the state rules in effect at that time are more likely to be the relevant one. Columns (6) and (7) reports results for, respectively, those with a high school education or less and those with some college or more; the results are substantively and statistically indistinguishable.

Finally, we have verified (in results not reported) that estimation of the reduced form OLS – in which *Protected* is replaced by *Protected_Hat* on the right hand side of equation (1) – yields qualitatively similar results to the instrumental variables estimation of equation (1), in which *Protected* is instrumented for with *Protected_Hat.* The coefficients on this reduced form estimation tend to be somewhat smaller (although still statistically significant) than the instrumental variables version; for example, a reduced form estimation of our preferred specification (shown in Table 3, column 2) yields a coefficient on *Protected_Hat* of -0.0052 (statistically significant at the 5 percent level) compared to the IV estimate of -0.109. This is consistent with the introduction of measurement error in using *Protected_Hat* instead of *Protected* to measure the Medicaid rules faced by a given household. By contrast, estimating equation (1) with *Protected* rather than *Protected_Hat* on the right hand side results in a positive coefficient; this suggests that the issue of the potential endogeneity of assets to the Medicaid rules is in fact quantitatively important for our estimates.

4. Simulated effects of potential Medicaid reforms

The preceding analysis suggests a statistically significant crowd-out effect of Medicaid on demand for private long-term care insurance. Our central estimate suggests that a $10,000 increase in the amount of assets a household can hold and still be eligible for Medicaid is associated with a 1.1 percentage point decline in the probability of holding long-term care insurance. Relatedly, these findings suggest that increasing the stringency of Medicaid's means testing – i.e. decreasing the amount of protected assets –

would induce greater demand for private insurance. In this section we provide a gauge of the magnitude of these crowd-out estimates by exploring their implications for the effect of potential Medicaid reforms in the size of the private long-term care insurance market.

In particular, we consider what our estimates imply for how much long-term care insurance holdings would increase if all of the states decreased the amount of protected assets to the minimum allowable under federal law in 2000. These minimum federally allowable asset protection laws were $16,824 for a married couple, and $2,000 for a single individual. Currently, only 3 states (Arkansas, DC and Oregon) have rules this stringent, while 26 states have the most generous asset protection rules allowed by federal law. Given current state rules and the distribution of assets in the data, we estimate that this change would decrease the average protected assets for an individual in our sample by a little under $25,000. Our point estimates in column 4 therefore imply that this decrease in asset protection would be associated with an increase in private long term care insurance coverage by about 2.7 percentage points. This represents about a 30 percent increase in coverage relative to the current coverage rate of 9.1 percent. However, it suggests that the vast majority of the individuals in our sample would remain without private insurance.

We also considered what our results imply for private long-term care insurance coverage if the minimum federally allowed asset protect laws were reduced by half (to $8,421 for a married couple and $1,000 for a single individual) and all states were to set their asset protection laws at their new minimum. Given the current state rules and the distribution of assets in the data, we estimate that this (out-of-sample) change would decrease the average protected assets for an individual in our sample by almost $30,000, or by $5,000 more on average than the previous reform we considered. Our crowd-out estimates imply that a $30,000 decline in the amount of protected assets would be associated with a 3.3 percentage point increase in private long-term care insurance coverage rates. While this represents a more than one-third increase over current insurance coverage rates, it would still leave over 85 percent of individuals in the sample without private insurance.

These empirical findings are broadly consistent with the simulation-based evidence in Brown and Finkelstein (2004b). They find that recent policy reforms adopted in several states to allow individuals who purchase private insurance to qualify for Medicaid coverage while retaining substantially more assets would have relatively little effect on the implicit tax that Medicaid imposes on private insurance, and hence little effect on demand for private insurance. Importantly, however, Brown and Finkelstein (2004b) estimate that Medicaid's implicit tax has a large crowd-out effect on private insurance demand. Changes in asset protection by themselves however do not affect this implicit tax much because as long as Medicaid remains a secondary payer, even without any asset limits to Medicaid eligibility a large portion of private insurance benefits are redundant of what Medicaid would otherwise have paid. (By the same token, removing the secondary payer status without changing the Medicaid asset limits similarly leaves a large Medicaid implicit tax). Our empirical findings, coupled with the simulation-based evidence in Brown and Finkelstein (2004b) thus underscore the importance of understanding not just the size of the crowd out effect, but also the *mechanism* behind it in considering the likely impact of potential reforms to the public program on private demand.

5. Conclusion

Long-term care is a large, and largely uninsured, potential expense facing the elderly. Medicaid serves as the insurer of last resort. As the baby boomers age, long-term care expenditures are expected to rise substantially, and with them Medicaid expenditures. This will put increasing pressure on state and federal budgets. As a result, increasing attention is focused on how public policy can stimulate the private long-term care insurance market.

This paper looks empirically at the effects of the Medicaid program on private long-term care insurance demand. We draw on the substantial variation in the level of assets that an individual can protect from Medicaid based on an individual's state of residence, marital status and asset holdings to identify the impact of Medicaid on private long-term care insurance demand. Our estimates suggest that more generous Medicaid asset protection is associated with less private long-term care insurance coverage. Our central estimate implies that a $10,000 increase in the amount of assets a household can retain while qualifying for Medicaid coverage of long-term care expenditures is associated with a 1.1 percentage point reduction in long-term care insurance coverage.

Although our findings point to a crowd-out effect of Medicaid asset protection on long-term care insurance demand, they also suggest that even large scale reductions in Medicaid asset protection are unlikely to stimulate private insurance coverage among most of the elderly population. We estimate that, if all states were to adopt the most stringent asset eligibility requirements allowed by federal law in 2000 – $16,824 for a married couple and $2,000 for a single individual – and thereby decrease average protected assets by about $25,000**,** overall demand for private long-term care insurance would rise by 2.7 percentage points, leaving almost 90 percent of the elderly still without private insurance.

These empirical findings complement recent simulation research (Brown and Finkelstein, 2004b) which also suggest that changes in Medicaid's asset protection rules would do little to address the lack of private long-term care insurance among most of the elderly. At the same time, Brown and Finkelstein (2004b) find that Medicaid may have a substantial crowd-out effect on long-term care insurance demand through the large implicit tax it places on the benefits from private long-term care insurance policies. Changes in Medicaid asset rules appear to not have much affect on this implicit tax, which may explain the simulation and empirical evidence that changes in Medicaid asset rules do not appear to have much effect on demand for private insurance. Together, these findings raise the important question of whether it is feasible to design Medicaid in a way that reduces the implicit tax it places on private insurance, and thus the constraints it appears to place on private insurance demand.

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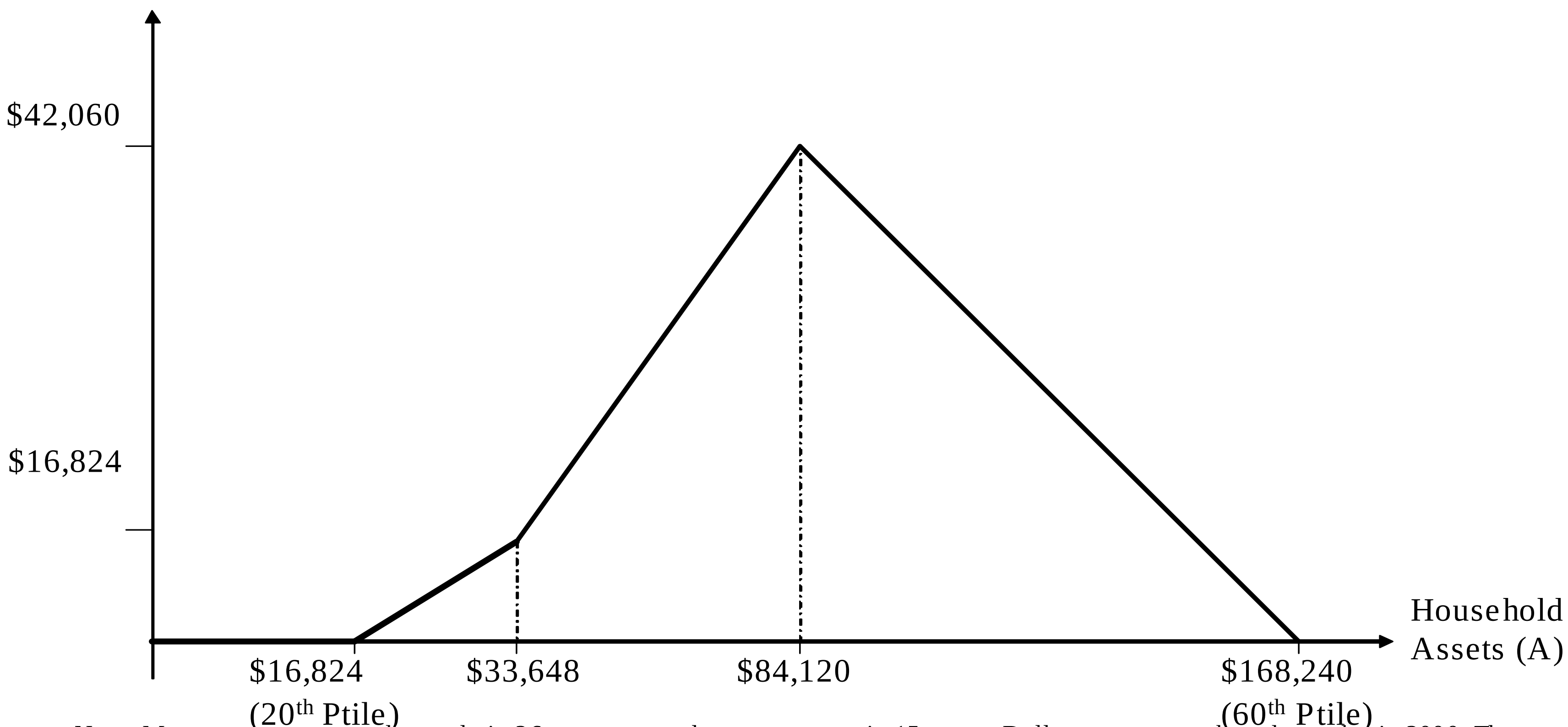
Figure 1:Difference in Protected Assets for Community Spouse

(Most common state rules – second most common state rules)



Note: Most common state rules apply in 26 states, second most common in 15 states. Dollar amounts are based on rules in 2000. The rules (and affected states) are described in more detail in Appendix A (see "case 1A" and "case 2A" respectively). Percentiles denotes the percentile of the asset distribution for married households in the 2000 HRS.

Table 1: Summary Statistics

	(1)	(2)
	Full Sample	Constant Medicaid rules sub-sample
Percent with LTC Insurance	9.6	9.1
Average Age	61.5	61.6
Percent Married	70.2	71.4
Percent Male	48.1	48.1
Percent Retired	40.2	39.8
Average number of children	3.3	3.3
Household Networth (in Thousands)		
Mean	385	367
25^{th} Percentile	50	48
Median	157	153
75^{th} Percentile	387	391
N	28,1000	17,623

Note: "Full Sample" consists of individuals aged 55-69 in the 1996, 1998 or 2000 HRS who report their marital status and long-term care insurance coverage. "Constant Medicaid Rules" sub-sample is restricted to individuals who did not experience changes in marital status during our data (1996 – 2000) and who are in one of the 30 states that did not have real Medicaid asset rule changes between 1991-2000. Net worth is the HRS imputed value for net worth. It includes net financing worth, housing wealth, and defined contribution pension values, but does not include DB pension or Social Security wealth. All statistics are calculated using household weights.

Table 2: Long-Term Care Insurance Ownership Rates

	(1)	(2)
Sample	Full Sample	Constant Medicaid rules sub-sample
Entire Sample	9.6	9.1
Males	9.7	9.0
Females	9.4	9.3
Singles	7.8	7.1
Marrieds	10.3	10.0
		8.1
age 55-61	8.9	
age 62-69	10.1	10.4
		9.9
wave = 1996	10.5	
wave = 1998	9.3	9.0
wave = 2000	9.1	8.6
		3.4
Net Worth, Bottom Quartile	4.1	
Net Worth, 2nd Quartile	7.7	7.1
Net Worth, 3rd Quartile	10.2	10.0
Net Worth, Top Quartile	15.1	14.8

Note: Full Sample consists of individuals aged 55-69 in the 1996, 1998 or 2000 HRS who report their marital status and long-term care insurance coverage. "Constant Medicaid Rules" sub-sample is restricted to individuals who did not experience changes in marital status during our data (1996 – 2000) and who are in one of the 30 states that did not have real Medicaid asset rule changes between 1991-2000. Net worth is the HRS imputed value for net worth. It includes net financing worth, housing wealth, and defined contribution pension values, but does not include DB pension or Social Security wealth. All statistics are calculated using household weights.

Table 3: Main crowd out estimates

	Full Sample	Constant Medicaid Rules Sub-Sample				
	(1)	(2)	(3)	(4)	(5)	(6)
Protected	-0.0056*	-0.0109**	-0.0093	-0.0103*	-0.0172**	-0.0153
	(0.0031)	(0.0048)	(0.0060)	(0.0050)	(0.0064)	(0.0112)
	[0.0045]	[0.0048]	[0.0052]	[0.0055]	[0.0053]	[0.0079]
Additional Controls			Asset HAT * married	Asset HAT * state	Married * State	All two-way interactions
Observations	24,841	15,576	15576	15576	15576	15576

Notes: Table reports the results of estimating equation (1) by instrumental variables (IV); *Protected_Hat* is used as an instrument for *Protected.* All regressions use household weights. *Protected* and *Protected_Hat* are measured in units of $10,000. “Constant Medicaid rules sub-sample” restricts the sample to individuals whose marital status did not change between 1996 and 2000 and who are in states whose real Medicaid asset disregards did not change between 1991 and 2000. Columns (3) through (6) add additional controls to the specification in column (2), as indicated in the row labeled “additional controls”. Specifically, column 3 adds an interaction between predicted assets and marital status, Column 4 adds interactions between predicted assets and state dummies, Column 5 adds interactions between marital status and state dummies, and column 6 adds all three of these sets of two-way interactions in the previous three columns. Standard errors allowing for an arbitrary variance covariance matrix within each state are (in parentheses); standard errors calculated using a non-parametric bootstrap are [in square brackets] (see text for more details). ***, **, * denotes statistical significance at the 1 percent, 5 percent and 10 percent level respectively, based on the standard errors (in parentheses). All regressions include indicator variables for marital status, state fixed effects, education categorical dummies, gender, occupation, industry, number of children up to five, age, retired, HRS cohort, race, Hispanic, and HRS wave, as well as education categorical dummies interacted with the indicators for marital status, gender, occupation, industry, number of children up to five, age, retired, HRS cohort, race, Hispanic and HRS wave.

Table 4: Sensitivity Analysis: Alternative specifications and sub-samples

	Baseline	Other Medicaid Parameters	20^{th} to 60^{th} pctile of Asset distribution	age 55-61	age 62-69	High School Education or Less	Some College or More
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
PROTECTED	-0.0109**	-0.0114**	-0.0223*	-0.0136**	-0.0043	-0.011**	-0.012*
	(0.0048)	(0.0055)	(0.0130)	(0.0053)	(0.0085)	(0.0052)	(0.006)
	[0.0048]	[0.0049]	[0.0127]	[0.0067]	[0.0068]	[0.0059]	[0.010]
INCOME		0.0363					
		(0.4285)					
LIENS		0.0087					
		(0.0175)					
Observations	15576	15576	6029	8552	7024	9,452	6,124

Note: Table reports the results of estimating equation (1) by instrumental variables (IV) on the constant Medicaid rules sub-sample; *Protected_Hat* is used as an instrument for *Protected.* All regressions use household weights. *Protected* and *Protected_Hat* are measured in units of $10,000. Column 1 replicates the baseline results (from Table 3, column 2). Column 2 adds controls for other Medicaid rules, specifically the amount of income (in units of $10,000) that a household can keep and still qualify for Medicaid ("Income"), and an indicator variable for whether a state will put a lien on a house when one spouse is in a nursing home in order to recoup expenses upon the death of the community spouse ("Liens"). Column 3 limits the sample to individuals in the 20^{th} to 60^{th} percentile of the asset distribution of married individuals (which is the range in which the variation in *Protected* is largest). Columns 4 and 5 look separately at individuals aged 55-61 and 62-69. Columns 6 and 7 look separately at individuals with a high school education or less and individuals with some college education or more. Standard errors allowing for an arbitrary variance covariance matrix within each state are (in parentheses); standard errors calculated using a non-parametric bootstrap are [in square brackets] (see text for more details). ***, **, * denotes statistical significance at the 1 percent, 5 percent and 10 percent level respectively, based on the standard errors (in parentheses). All regressions include indicator variables for marital status, state fixed effects, education categorical dummies, gender, occupation, industry, number of children up to five, age, retired, HRS cohort, race, Hispanic, and HRS wave, as well as education categorical dummies interacted with the indicators for marital status, gender, occupation, industry, number of children up to five, age, retired, HRS cohort, race, Hispanic and HRS wave.

Appendix A: Overview of Medicaid Rules

This appendix discusses in some detail the rules that govern financial eligibility for Medicaid. We focus primarily on the rules regarding the amount of financial assets that an individual or couple is permitted to keep and still receive Medicaid reimbursement for nursing homes. We label this value *Protected* in our empirical work in the main body of the paper; it is our key variable of interest. We also briefly discuss the rules regarding Medicaid income disregards and Medicaid treatment of housing wealth; we briefly explore the impact of these variables in the sensitivity analysis presented in Table 5.

Medicaid rules vary considerably across states. Within each state, the rules pertaining to a single individual who goes into a nursing home differ from those pertaining to married individuals who go into a nursing home whose spouse remains in the community. Since the differential rules within state by marital status form the main source of our empirical identification strategy, we discuss these differences in some detail.

Appendix Table A1 provides a summary of the various Medicaid rules for nursing home expenses for single and for married people in each state as of 2000. We now discuss each briefly in turn. At the end of this section, we turn to a discussion of state rule changes over the 1991 to 2000 period.

<u>Rules for single individuals in 2000</u>

Medicaid rules for single individuals are relatively simple and uniform across states.

Maximum amount of retainable assets: They can keep no more financial wealth than the Medicaid-specified asset limit. Anything above these must go toward paying for the care. In 2000, the modal asset limit was $2,000 (which nearly 70 percent of states used). The remaining states have an asset limit that ranges from $1,500 to $6,500.

Maximum amount of retainable monthly income: They can keep no more monthly income than the Personal Needs Allowance (PNA). In 2000, these ranged from $30 to $77 per month.
Rules regarding housing wealth: They must sell their house (and use the proceeds to pay for that care), unless there is a chance of recovery or a dependent child living in the house.

<u>Rules for married individuals (one spouse in NH, one in community)</u>[6]

Treatment of financial assets:

When one spouse enters a nursing home, total household financial (non-housing) assets (A) are attributed evenly between the two spouses. From this even attribution, the spouse that goes into a nursing home is allowed to keep only the Medicaid-specified income and asset limits for single individuals (i.e. $30-$77 a month of income and approximately $2,000 of assets).

The main source of variation used in our empirical work is the amount of assets that the community spouse is allowed to keep. This amount depends on the total amount of household financial assets (A) and the state rules regarding the minimum and maximum assets that the community spouse is allowed to keep (STATEMIN and STATEMAX respectively). A community spouse whose share of the assets is below the state minimum allowable (STATEMIN) is allowed to take assets from the nursing home spouse to top-up their asset level up to STATEMIN.

In setting the minimum and maximum amount of assets the community spouse can retain, the states are constrained to set a minimum (STATEMIN) that is at least as high as the federal

[6] For the less common case when both spouses need nursing home care, they are essentially treated as two single individuals in terms of the treatment of assets, income and housing. The one exception is that some states set a lower threshold for the amount of assets the couple can keep ($3,000 combined instead of $2,000 each).

minimum (FEDMIN) and to set the maximum (STATEMAX) no higher than the Federal maximum (FEDMAX).[7]

In all states therefore, married couples with combined assets of less than the Federal Minimum ($16,824 in 2000) face the same treatment of their assets (they are allowed to keep all of them). Furthermore, in all but 5 states, married couples with combined assets of more than twice the Federal maximum (i.e. $168,240 in 2000) are allowed to keep the same amount ($84,120).[8] The differential treatment across states of married couples' assets therefore occurs mainly for couples with assets between the Federal Minimum and twice the Federal Maximum. In 2000, these limits were $16,824 and $168,240 respectively, and correspond to the 20th and 58th percentile of financial assets for married households in the 2000 HRS. Married households in this "affected range" have average assets of $74,266.[9]

For married couples with assets between the Federal Minimum and twice the Federal Maximum, there is substantial variation across states in the amount of assets they are allowed to retain. This variation arises from where states choose to set their State Minimum and State Maximum allowable retainable assets.

Figure A1 compares the amount of assets that the community spouse can keep under the two most common state rules. In the first, which is used in 26 states, the state sets both the minimum and maximum allowable assets (STATEMIN and STATEMAX) to the Federal Maximum. Under these rules, the household faces a 0 percent marginal tax rate on its assets until it reaches the state minimum amount of allowed retainable assets, at which point it faces a 100 percent marginal tax rate on all further assets. We refer to this as "Case 1A" and illustrate it with the dark line in Figure A1.[10]

The second most common set of state rules (which apply in 15 states), set the state minimum equal to the federal minimum allowable retainable assets, and the state maximum equal to the federal maximum allowable retainable assets. This is shown by the dashed line in Figure A1. In this case, the marginal tax rate faced on assets goes from 0%, to 100%, to 50% and then back to 100%, as shown in the figure.[11]

Figure 1 graphs the *difference* in the amount of assets a community spouse can keep (as a function of total household financial assets) if they are in a Case 1A state relative to a Case 2A As is readily apparent, this difference is non-monotonic in the couples' assets.

The difference in the amount of allowable retained assets is quite substantial. For example, if all married couples were to move from Case 1A states to Case 2A states, the average difference in the amount a household is allowed to keep would be $8,277 in the whole sample, which is approximately 2 percent of average assets. For those in the "affected range" (20th to 58th percentile of assets i.e. between $16,824 and $168,240), this move would on average allow them to keep $21,715 more in assets, which represents 29 percent of average financial assets in this range. The maximum change in the amount of assets that the household would be able to keep is $67,296 and occurs for households with assets of $84,120.

Finally, we note that while we have focused on the two major types of state rules, there are 10 other states whose rules differ from those in Cases 1A and 2A. For the sake of brevity we do not discuss these rules in detail; they are summarized in Table A1. Like the two more common cases

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[7] In 2000, the FEDMIN and FEDMAX were, respectively, $16,824 and $84,120 (Stone 2002). They are indexed to the CPI but have otherwise remained unchanged between the 1991 and 2000 period.

[8] The five state exceptions are AR, DC, NY, OR, and SC.

[9] These are summary stats on the 2000 HRS, using financial assets. The sample is limited to married households aged 60-70. The cut points in the distribution are identical for the age 75+ sample (20th-58th percentile).

[10] The 26 states in this category are: AK, AZ, CA, CO, FL, GA, HI, IL, KY, LA, MA, MD, ME, MI, MS, MO, ND, NE, NV, OK, SD, VT, WA, WI, WV, and WY.

[11] The 15 states in this category are: CT, ID, IN, KS, MT, NC, NH, NJ, OH, PA, RI, TN, TX, UT, and VA.

discussed above, the difference across states in the treatment of assets is non-monotonic in the couples' assets across these other cases as well (relative to each other or the two more common cases). It is also non-trivial in magnitude.

Treatment of Income: Income is split based on the "name on the check" rule, rather than evenly between the two spouses as is the case for assets, in all but 2 states. The institutionalized individual is allowed to keep the same amount of income as a single household (defined above as the PNA). The community spouse is allowed to keep an unlimited amount of income if it is in her name. If the community spouse's income is below a minimum amount known as the community spouse income limit (CSIL), then she is eligible to keep enough of her institutionalized spouse's income to bring her total income up to that limit. This minimum income amount varies across states from approximately $1,407 to $2,133 per month as shown in Table A1.

Treatment of Housing: The house of a community spouse is left out of the calculation of assets or income and it is completely protected for the community spouse during her lifetime. It may be of unlimited value. However, about one-quarter of states will put a lien on this house, which allows the state to collect money from the sale of the house to reimburse them for their Medicaid outlays upon the sale of the house or the death of the community spouse. We refer to such states as LIEN states. Enforcement of estate recovery practices varies across states (Sabatino and Wood, 1996).

<u>Changes in Rules.</u>

All of the preceding discussion applies to the state rules in 2000. For purposes of the empirical work, it is important to know whether states changed their rules at some point, as individuals might have purchased or considered purchasing long-term care insurance under a different set of rules. As discussed in the text, there was a major change in rules in 1988 (effective in 1989), which motivated our focus on an age group who was likely to be of buying age after 1988. We also tracked down information on rule changes between 1991 and 2000. There is no central database for state-specific Medicaid eligibility rules. We compiled a timeline for these state-specific rule changes by collecting a variety of different sources that covered the different years; where sources disagreed, we telephoned the relevant agency in the state to ascertain the correct information. We were unable to obtain state-specific information between 1989 and 1991.[12]

There were no major changes in the Medicaid rules for single individuals during this time period. However, for married individuals, 21 states changed their assets protection rules for the community spouse between 1991 and 2000 (see Appendix Table A1). In addition, 27 states changed the allowable income limit for institutionalized individuals (PNA). Finally, over our period 13 states introduced estate recovery plans (LIENS).

[12] Our specific sources for state Medicaid rules from 1991- 2000 are: Bruen et al (1998); Congressional Research Service (1993); Horvath (1997); Kassner and Shirley (1998); the National Association of Medicaid Directors; Price (1996); Sabatino and Wood (1996); Schwab (1998); Stone (2002), and telephone calls to particular states.

Table A1: Medicaid Eligibility Parameters by State, 2000

State	Rules for single individuals		Rules for married individuals with community spouse			
	Max amount of retainable assets	Max amount of retainable monthly income (PNA)	Min amount of retainable assets for community spouse (STATEMIN)	"Min" amount of retainable monthly income for community spouse (CSIL)	Is lien put on community spouse's house?	No change asset rules for married individuals (1991-2000)
AK	2,000	75	84,120	2,103	0	1
AL	2,000	30	25,000	1,407	1	0
AR*	2,000	40	16,824	1,407	0	1
AZ	2,000	76.80	84,120	2,103	0	1
CA	2,000	35	84,120	2,103	1	1
CO	2,000	50	84,120	1,407	1	0
CT	1,600	52	16,824	1,407	1	1
DC*	2,600	42	16,824	2,103	0	1
DE	2,000	70	25,000	1,407	1	0
FL	2,000	35	84,120	2,103	0	1
GA	3,000	30	84,120	2,103	0	1
HI	2,000	30	84,120	2,103	1	1
IA	2,000	30	24,000	2,103	0	0
ID	2,000	30	16,824	1,407	1	1
IL	2,000	30	84,120	2,103	1	1
IN	1,500	50	16,824	1,407	0	1
KS	2,000	30	16,824	1,407	0	1
KY	2,000	40	84,120	2,103	0	1
LA	2,000	38	84,120	2,103	0	1
MA	6,500	60	84,120	1,407	1	0
MD	2,500	40	84,120	2,049	1	0
ME	3,000	40	84,120	2,103	0	0
MI	2,000	60	84,120	2,103	0	0
MN	3,000	67	23,774	1,407	1	1
MO	2,000	30	16,824	1,407	0	1
MS	2,000	44	84,120	2,103	0	1
MT	2,000	40	16,824	2,103	1	1
NC	2,000	30	16,824	2,103	0	1
ND	3,000	40	84,120	2,103	0	1
NE	4,000	50	84,120	1,407	0	0
NH	2,500	50	16,824	2,103	0	1
NJ	4,000	35	16,824	1,407	0	1
NM	2,000	45	31,290	1,407	0	1
NV	2,000	35	84,120	2,103	0	0
NY*	3,600	50	74,820*	2,103	1	0
OH	1,500	40	16,824	1,407	0	1
OK	2,000	50	84,120	2,103	0	0
OR*	2,000	30	16,824	1,407	0	1
PA	2,400	30	16,824	1,407	0	1
RI	4,000	50	16,824	1,407		1

SC*	2,000	30	66,480	1,662	0	0
SD	2,000	30	84,120	1,407	0	0
TN	2,000	30	16,824	1,407	0	1
TX	2,000	45	16,824	2,103	0	1
UT	2,000	45	16,824	1,407	0	1
VA	2,000	30	16,824	1,407	0	1
VT	2,000	47.66	84,120	1,407	0	1
WA	2,000	41.62	84,120	1,407	0	1
WI	2,000	45	84,120	1,875	1	1
WV	2,000	50	84,120	1,407	0	0
WY	2,000	30	84,120	2,103	0	1

Source: Stone 2002, Sabatino and Wood 1996, and authors' corrections based on telephone conversations with particular states where other sources disagreed with those listed here.

Notes:

- For all states, the maximum amount of assets the community spouse is allowed to keep (STATEMAX) is the same (and equal to FEDMAX) unless the state is denoted with a * in which case STATEMAX=STATEMIN.
- Treatment of housing for single individuals is not described in the table since it is the same in all states (see text).
- PNA stands for Personal Needs Allowance
- CSIL stands for Community Spouse Income Limit.
- In 2000, FEDMIN and FEDMAX were $16,824 and $84,120, respectively.

Figure A1:Medicaid Marginal Tax

Married Households

STATEMAX=FEDMAX