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Why Did the SSI-Disabled Program Grow So Much?
Disentangling the Effect of Medicaid

Aaron S. Yelowitz
University of California, Los Angeles

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Abstract

The participation rate for working-age adults in the Supplemental Security Income (SSI) program increased by 37 percent from 1987 to 1993. This paper examines the role of public health insurance provided through Medicaid on the SSI participation decision. I use the rapid growth in Medicaid expenditure across states and over time as a proxy for its value. The estimation is complicated by the easing of standards for determining disability. If the marginal individual who entered SSI under these easier standards was healthier than the average participant, then average Medicaid expenditure would fall. Thus, conventional OLS estimates could lead to a spurious negative correlation between average Medicaid expenditure and SSI participation. I therefore apply two-stage least squares (TSLS) to estimate Medicaid's effect, using Medicaid expenditure for blind and elderly SSI recipients, and adult and child AFDC recipients as instruments for disabled Medicaid expenditure.

The TSLS estimates indicate that rising Medicaid expenditure significantly increased the SSI participation for whites, but had little effect on African Americans. Among whites, the rising value of Medicaid explains one-third of the growth in SSI participation.

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1. INTRODUCTION

From 1984 to 1993, the disabled Supplemental Security Income (SSI) population grew at an annual average rate of 9.2 percent.[1] This study asks whether the availability of public health insurance through the Medicaid program contributed to the caseload growth. I focus on the SSI participation behavior of adults between the ages of 18 and 64 using Current Population Survey (CPS) data spanning the calendar years 1987 to 1993.

Although every state offers Medicaid to disabled SSI recipients in some form, the value of the insurance varies. Each state has considerable leeway in the scope of health care services and access to care from Medicaid. Following the recent empirical approaches of Blank (1989) and Winkler (1991), who estimate the effect of Medicaid on female-headed households, I proxy for Medicaid's value with the average Medicaid expenditure for disabled SSI recipients in each state and year. If increased Medicaid expenditure reflects a greater valuation, then this proxy should result in increased SSI participation.

An important endogeneity problem emerges in using average Medicaid expenditure of disabled recipients to predict SSI participation. Beyond changes in medical prices, scope of services, and access to care, variation in expenditures may reflect changes in the underlying health status for disabled SSI recipients. Since the criteria in evaluating disability are somewhat subjective, the standards may change. If the standards become less strict, then relatively healthy people will enter the SSI rolls, thereby lowering the average expenditure.[2] Thus, the SSI rolls may increase at the same time that

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[1]U.S. House of Representatives (1994).

[2]To illustrate, the Social Security Administration conducted several outreach activities for SSI. If these outreach programs attracted relatively healthy new recipients, then the average health status may have improved. See U.S. House of Representatives (1993) for a discussion of these activities.

average Medicaid expenditure falls, yielding a spurious negative correlation that is driven by changes in health status. To correct this bias, I therefore apply two-stage least squares (TSLS) with four instruments: the average Medicaid expenditure for SSI blind and elderly, and the average Medicaid expenditure for AFDC children and adults. The instruments reflect variation in the generosity of the state's Medicaid package (access to and quality of care, medical prices, and scope of services), and they are not correlated with changing definitions of disability. Since these four groups are well-defined and the criteria for Medicaid eligibility are much more objective, average Medicaid expenditure reflects the true variation in Medicaid's value.

The results support the preceding story. The ordinary least squares (OLS) estimates yield small and imprecise estimates of Medicaid expenditure on SSI participation. In contrast, the TSLS estimates yield results that are more than four times as large. Not only are the instruments extremely powerful in explaining average Medicaid expenditure for the disabled in the first stage, but most of the models presented also pass overidentification tests. I conclude that the rising value of Medicaid contributed greatly to the increase in the SSI rolls in the late 1980s and early 1990s. Medicaid explains as much as one-third of the SSI growth. Moreover, the effects are concentrated in the white population and not the African-American population.

The paper is organized as follows. Section 2 describes some background on SSI and Medicaid, and reviews the economic importance of Medicaid for other populations. This section also discusses the practical problems that previous research has encountered in isolating Medicaid's effect. Section 3 presents some theoretical considerations. The institutional detail is incorporated into a budget constraint, and implications for SSI participation are discussed. Section 4 presents a descriptive analysis of the CPS data. This section documents the very different trends in SSI and Medicaid participation by race and gender for working-age adults. Section 5 presents results from OLS and TSLS. I provide tests on the validity of the instruments, and perform some sensibility checks. Section 6 concludes and presents further extensions.

2. BACKGROUND

Background on the SSI-Disabled and Medicaid Programs

SSI was introduced in 1974, replacing state-run programs for the needy aged, blind, and disabled. In 1993 alone, $23.5 billion was spent on SSI cash benefits for these groups. While the number of elderly and blind participants remained stable, the number of disabled SSI participants increased from 2.9 million in 1988 to 4.0 million in 1993.

A poor adult must be disabled to qualify for SSI. For purposes of eligibility, disabled individuals are defined as those "unable to engage in any substantial gainful activity by reason of a medically determined physical or mental impairment expected to result in death or that has lasted, or can be expected to last, for a continuous period of at least 12 months."[3] While this definition may appear to be quite objective, eligibility standards, especially for mental impairments, have changed due to legislative, regulatory, and judicial action.[4]

Besides receiving a monthly cash supplement, SSI provides the disabled adult with a second valuable benefit: Medicaid coverage. Each state's Medicaid program offers its own package of covered medical services that fall within broad federal guidelines. Federal law requires states to offer eight mandatory services and allows them to offer up to thirty-one optional services.[5] The 15 percent of all

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[3]U.S. House of Representatives (1993).

[4]Ross (1995).

[5]Required coverage includes inpatient and outpatient hospital services, rural health clinic services, federally qualified health center services, laboratory and X-ray services, nursing facility services for individuals under age 21, family planning services, physicians' services, home health services for any individual entitled to nursing facility care, nurse-midwife services, and services of certified nurse practitioners.

Medicaid beneficiaries who are disabled account for a far greater share of Medicaid's costs than non-disabled recipients. The average spending on blind and disabled beneficiaries amounted to $9,226 per beneficiary in 1993.[6]

Disabled adults are excluded from most public health insurance except through SSI. A notable exception to this is section 1619 of the SSI law, which is intended to remove some of the work disincentives for the disabled. Section 1619(a) provides continuation of cash benefits even if earnings exceed the "substantial gainful activity" level, as long as the disabling condition has not improved. Under section 1619(b), disabled individuals can continue to be eligible for Medicaid even if their earnings take them past the SSI income limit. These provisions turn out to be quite minor, however. In September 1992, just 48,000 of the 2.6 million disabled adults between the ages of 18 and 64 participated in either the 1619(a) or 1619(b) program.[7] For these provisions to be applicable, an individual must still initially qualify for and participate in SSI. So the provisions are not really an avenue off SSI.

Prior Studies of Medicaid and Welfare Participation

Although Medicaid was introduced thirty years ago and program costs have been soaring, only recently has it garnered much academic interest. One reason Medicaid's effect on SSI participation has been ignored is because the behavioral elasticities of the blind, elderly, and disabled were believed to be extremely small. In addition, estimating Medicaid's impact on welfare participation is complicated by the high correlation between eligibility for Medicaid and cash benefits for the disabled.

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[6]This figure is in nominal dollars. The expenditure numbers on the disabled throughout will include Medicaid spending in intermediate-care facilities and skilled nursing homes. It is important to include this component because access to these facilities is, indeed, a part of Medicaid's value. While only a small portion of the population will become institutionalized, it is also true that only a small portion will use any particular Medicaid service.

[7]U.S. House of Representatives (1993).

Though there are few existing examinations of Medicaid and SSI, several studies have looked at the impact of Medicaid on AFDC participation and work effort. The earlier studies found that Medicaid had a surprisingly small effect on the welfare and work choices for female heads, but more recent work has found larger effects.

Blank (1989) uses cross-sectional variation in average Medicaid expenditure, which varies tremendously across states, to examine AFDC participation.[8] Using data from the 1980 National Medical Care Utilization and Expenditure Study (NMCUES), she finds that health problems significantly increase AFDC participation, but that program rules do not. The insignificant effect of the Medically Needy (MN) program is not surprising because eight of the thirty MN states in her sample had an income eligibility level below the maximum AFDC payment level. What is surprising is the robustness of the finding that the state-specific Medicaid insurance value did not affect AFDC participation.

Moffitt and Wolfe (1992) construct an individual-specific valuation of health insurance to surmount Medicaid's collinearity with AFDC eligibility. They note that a Medicaid variable constructed from a state-specific average may badly measure Medicaid value for any particular family. Linking the 1984 Survey of Income and Program Participation (SIPP) and 1980 NMCUES, they construct a "heterogeneity" index for Medicaid's value based on different health characteristics of the woman and her family. This index yields enormous variation in Medicaid. Using this variation, they find sizable effects of Medicaid on labor market outcomes.

I examined expansions in Medicaid eligibility targeted toward young children from 1988 to 1991 (Yelowitz 1995a). These expansions linked Medicaid eligibility to the federal poverty line rather

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[8]Winkler (1991) examines both AFDC participation and labor supply using the 1986 CPS. In her model, she cashes out Medicaid at the market value for each state, in a fashion similar to Blank (1989). She finds that Medicaid generally has a modest, but statistically significant, impact on labor force participation, but no effect on hours of work or AFDC participation.

than to a state's AFDC income eligibility limit, thus offering an incentive to leave welfare. I found that these reforms significantly decreased AFDC participation and increased labor force participation. Among female-headed households, the effects were largest for previously married women, but were negligible for never-married women.

Very little evidence exists on the interaction of Medicaid and SSI. I examined recent changes in the Medicaid program on the SSI participation for a different group, senior citizens (Yelowitz 1995b). By using the implementation of a buy-in program for Medicare in the 1980s (which offered a substitute for the cost-sharing provisions of Medicaid), I found significant interactions: Medicaid has a bigger impact on exits from SSI for the elderly than it has on exits from AFDC for female heads.

3. THEORETICAL CONSIDERATIONS

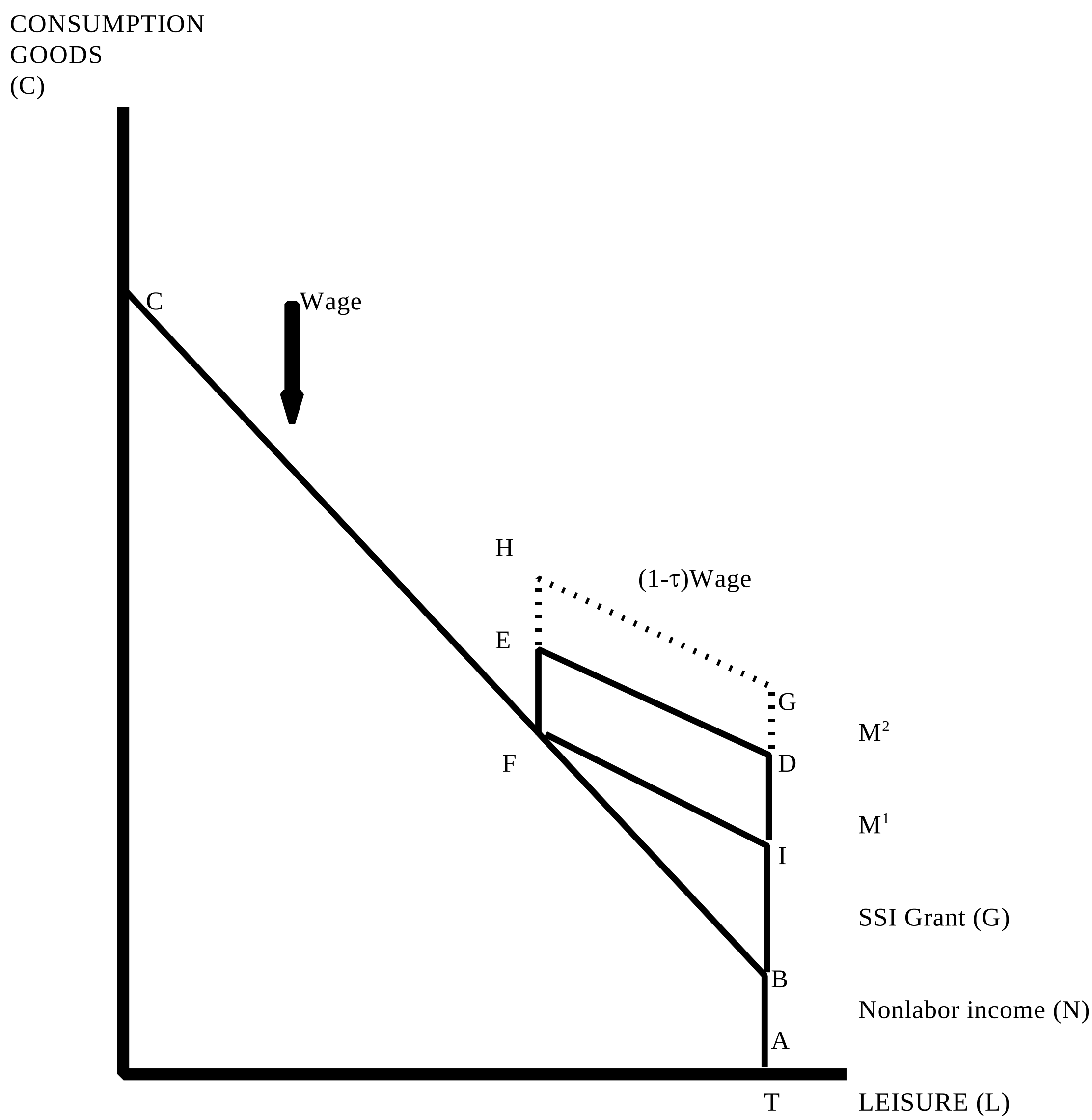
This section briefly outlines several ways that Medicaid's value influences SSI participation. The individual maximizes a utility function, U(C,L), which is a function of consumption goods (C) and leisure (L). The price of consumption goods (P_C) is normalized to $1 per unit, while the price of leisure is simply the wage rate (W). He is given a time endowment (T), which he can allocate between work and leisure. He may also receive nonlabor income (N), for instance from the earnings of his spouse. Therefore, his full budget constraint is initially defined as:

(1) $P_C C + WL = WT + N.$

In Figure 1, this is represented as the segment ABC. Given this budget constraint, the consumer maximizes his utility.

By introducing the SSI system into the model, the government changes the budget constraint. The program offers a grant (G), which was $669 per month for a married couple in 1994, and

Figure 1


CONSUMPTION
GOODS
(C)
C
Wage
H
(1-τ)Wage
E
G
M²
F
D
M¹
I
SSI Grant (G)
B
Nonlabor income (N)
A
T
LEISURE (L)

reduces this grant for earning income in the labor market. This reduction, known as the "benefit reduction rate" (τ), is 50 percent on earned income. Therefore, the net wage falls to $(1-\tau)W$ along the initial part of the budget constraint. The budget constraint with SSI cash benefits is characterized by the segment AIFC.

The final institutional feature is Medicaid. Broadly speaking, Medicaid is received when the individual is on SSI, and is entirely lost after leaving SSI. This discrete drop in benefits is known as the "Medicaid notch"—the design of the program creates a portion of the budget constraint where no utility-maximizing person should choose to be.

Variation in the value of Medicaid changes the budget constraint. Consider an individual who lives in a state where Medicaid is valued at some small amount, M^1—this can be thought of as the dollars the family would have to spend on medical expenses in the absence of insurance. His budget constraint now is represented by ADEFC in Figure 1. Consider a second individual who lives in a different state that has the same SSI grant but a more generous Medicaid program, so that the value is M^1+M^2. In this state, the budget constraint is represented by AGHFC. Relative to the first individual, the second individual would be more likely to participate in SSI.

This model could also be amended to include the stigma of program participation by adding an argument to the utility function, $U=U(C,L,P)$, where the act of SSI participation (P) lowers utility. As Moffitt (1983) explains, virtually all U.S. transfer programs have many eligible people who do not participate. By increasing Medicaid's value from M^1 to M^1+M^2, SSI participation could increase for two reasons. First, a person who was initially located somewhere along the segment FC (i.e., initially ineligible) may now find that he receives higher utility on segment GH, and therefore cuts back on his hours of work. Second, some who are eligible and initially located on segment BF (i.e., nonparticipating eligibles) now find Medicaid's value high enough that it outweighs the stigma cost of welfare.

Two implicit assumptions deserve mention. First, the model assumes that the individual does not have access to private health insurance. Clearly, the importance of Medicaid and its effect on SSI participation should be much more important for individuals without other health insurance opportunities, such as coverage through an employer (it is possible for Medicaid to still have some effect, however, if the scope of services differs from the private package). Second, the model does not account for the effects of health status on labor supply and SSI participation.[9] Poor health has at least three different effects on the budget constraint. It lowers the wage that the individual can receive in the labor market either by limiting the type of job and hours of work he can take, or by lowering his productivity. Poor health also changes preferences toward work and leisure: at any bundle, the marginal rate of substitution rises with poor health. Finally, poor health increases the value of Medicaid—since expected utilization is higher, the benefits are worth more than before. Unless accurate proxies for the individual's health status can be found, models relying on variation in Medicaid's value generated by health status may mistakenly attribute changes in preferences, productivity, and wages to Medicaid.

4. DESCRIPTIVE ANALYSIS

I use the 1988–94 CPS March Annual Demographic File, which provides retrospective information on family income, health insurance coverage, and program participation from 1987 to 1993 for the noninstitutionalized population. Because only a small fraction of the adult population participates in SSI-disabled, a large data set is essential to observe trends. Therefore, the CPS is perhaps a more useful household data set than others. I begin the analysis with the 1988 CPS because several additional questions on health insurance coverage were added, making later surveys less

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[9]See Wolfe and Hill (1995) for a model that explicitly accounts for the effect of health on the labor supply decisions and welfare participation decisions of single women.

comparable to earlier ones.[10] I end the analysis with the 1994 CPS because the last data on Medicaid expenditure (the key independent variable) is for fiscal year 1993.

Table 1 shows the sequential selection criteria and the number of observations eliminated from each screen for each CPS year. I use about one-third of the roughly 1.05 million observations contained in the 1988–94 CPS files. The nine most important exclusions are: being over the age of 64, being under the age of 18, living in Arizona, having imputed information on SSI or Medicaid receipt, having an imputed spouse number, being a woman under the age of 45, being a race other than African American or white, living in a single-parent household, and having more related children than own children in a family.

The motivation behind these exclusions deserves some explanation. First, I restrict my attention to working-age adults who would be unlikely to collect Medicaid from a program other than SSI-disabled. Thus, I exclude single-parent households with children under age 18 because the mother and children may be eligible for Medicaid under AFDC. Second, I eliminate women between the ages of 18 and 44 from my sample. For this group, pregnancy is the primary health insurance expense, and other reforms in Medicaid from 1984 onward could bias the results for SSI participation.[11] Third, I follow Winkler (1991) in excluding Arizona from the analysis. Arizona had a Medicaid demonstration project for part of the period I examine, and data on average Medicaid expenditures is not available for all years.

Table 2 presents summary statistics for the variables used in the analysis for the entire population, SSI nonrecipients, and SSI recipients. Since only 4,058 of the 345,453 observations are SSI recipients, the means of demographic variables for the nonrecipients closely match those of the

[10]These questions specifically dealt with the health insurance status of children in the household. Survey respondents were effectively asked twice about the health insurance coverage of children.

[11]See Currie and Gruber (1994) for an analysis of these Medicaid pregnancy expansions.

TABLE 1

Sample Selection Criteria, Current Population Survey: March Annual Demographic File

	1988	1989	1990	1991	1992	1993	1994
Initial Number of Observations	155,980	144,687	158,079	158,477	155,796	155,197	150,943
Over 64	18,610	17,740	18,902	19,043	18,954	19,074	18,574
Under 18	43,032	39,482	43,281	43,762	42,700	42,901	42,337
Lived in Arizona	1,091	1,045	1,078	1,057	993	974	1,017
Imputed disability status	287	280	367	291	274	414	486
Imputed SSI receipt	463	447	427	469	333	354	185
Imputed SSI value	74	78	86	91	103	96	207
Imputed Medicaid	1,188	1,067	1,208	1,378	1,429	1,504	1,523
Imputed veteran status	495	418	503	524	508	471	N/A
Imputed age	280	190	199	142	212	187	308
Imputed marital status	1,007	900	432	360	272	311	501
Imputed spouse number	1,212	1,606	2,309	2,223	969	902	133
Imputed sex	172	166	157	160	140	159	195
Imputed education	443	328	284	231	302	201	271
Imputed race	41	38	53	36	33	34	628
Women under age 45	31,077	28,520	31,789	31,693	31,323	30,611	29,592
AFDC participants	276	223	266	266	297	305	300
Not African American or white	1,952	1,820	2,148	2,290	2,381	2,624	2,741
Imputed wage/salary income	548	505	561	514	461	434	634
Imputed worker's comp income	112	93	106	141	114	95	140
Imputed veterans benefit	84	78	86	79	69	64	57
Imputed disability income	95	79	97	81	105	105	92
Female head with child present	1,164	1,045	1,222	1,234	1,244	1,254	1,236
Male head with child present	804	745	885	852	928	969	913
Related children in family	1,232	1,058	1,249	1,298	1,247	1,262	1,360
Final Number of Observations	50,241	46,736	50,384	50,262	50,405	49,902	47,523

TABLE 2

Summary Statistics, 1987–1993

Variable Name	Entire Sample		Nonrecipients		SSI recipients	
SSI Participation	0.0117	(0.0002)	0.0000	(0.0000)	1.0000	(0.0000)
Medicaid participation	0.0230	(0.0003)	0.0124	(0.0002)	0.9117	(0.0044)
Medicare participation	0.0248	(0.0003)	0.0218	(0.0003)	0.2789	(0.0070)
Annual Medicaid benefit for SSI-disabled	$8,161	(6.30)	$8,163	(6.34)	$7,928	(56.61)
Annual Medicaid benefit for SSI blind	$7,313	(11.89)	$7,306	(11.93)	$7,947	(130.11)
Annual Medicaid benefit for SSI elderly	$7,771	(7.64)	$7,769	(7.67)	$7,912	(76.52)
Annual Medicaid benefit for AFDC child	$1,054	(0.61)	$1,054	(0.62)	$1,074	(5.91)
Annual Medicaid benefit for AFDC adult	$1,861	(0.87)	$1,861	(0.88)	$1,856	(7.99)
Annual SSI cash benefit	$7,143	(3.56)	$7,158	(3.58)	$5,909	(26.99)
Section 209(b) state	0.2462	(0.0007)	0.2461	(0.0007)	0.2543	(0.0068)
Unemployment rate	0.0616	(0.0000)	0.0615	(0.0000)	0.0640	(0.0003)
State's labor force participation	4,530,601	(6,727)	4,528,145	(6,765)	4,737,266	(63,439)
Respondent's age	42.26	(0.0222)	42.19	(0.0223)	47.83	(0.2041)
African American	0.0764	(0.0004)	0.0747	(0.0005)	0.2198	(0.0065)
Resides in central city	0.2104	(0.0006)	0.2090	(0.0006)	0.3299	(0.0073)
Education<9	0.0655	(0.0004)	0.0616	(0.0004)	0.3881	(0.0076)
9≤ Education<12	0.1008	(0.0005)	0.0992	(0.0005)	0.2331	(0.0066)
Education=12	0.3737	(0.0008)	0.3749	(0.0008)	0.2703	(0.0069)
Education>12	0.4599	(0.0008)	0.4641	(0.0008)	0.1084	(0.0048)
Currently married	0.6645	(0.0008)	0.6696	(0.0008)	0.2331	(0.0066)
Number of own children under age 6	0.1929	(0.0008)	0.1947	(0.0009)	0.0384	(0.0040)
Number of own children aged 6 to 17	0.4237	(0.0014)	0.4272	(0.0014)	0.1286	(0.0083)
Male	0.7506	(0.0007)	0.7530	(0.0007)	0.5510	(0.0078)
Veteran	0.2053	(0.0006)	0.2069	(0.0006)	0.0694	(0.0039)
Private health insurance coverage in own name	0.5907	(0.0008)	0.5973	(0.0008)	0.0401	(0.0030)

Source: Results from author's tabulation of the March 1988–94 Current Population Survey Annual Demographic File.

Notes: Standard errors in parentheses. Total number of observations is 345,453, of whom 4,058 are SSI recipients. All dollar amounts are in 1990 dollars.

entire population. Over the entire sample, SSI participation is 1.17 percent, while Medicaid participation is nearly double that number, 2.30 percent. Even with the exclusion of single-parent households, some families may have access to Medicaid from sources other than SSI-disabled. Part of the gap between the two participation rates could be the result of the existing Medically Needy and General Assistance programs. Moving to the final column, more than 90 percent of SSI recipients also report Medicaid coverage. There are at least two reasons why Medicaid participation may not be complete for SSI recipients. First, the survey respondent might report Medicaid receipt only if he actually went to the hospital. Second, because a number of states require a second application for Medicaid, the respondent may not apply for benefits until he becomes sick. This table also shows Medicare participation averages 27.9 percent for SSI recipients and 2.2 percent for nonrecipients. Since an SSI recipient is much more likely to participate in the Disability Insurance (DI) program than a nonrecipient, a prolonged SSI spell can result in Medicare coverage. A nonrecipient can also qualify for DI and thereby qualify for Medicare.

The next seven rows in Table 2 illustrate state-level policy variables characterizing the Medicaid and SSI programs.[12] The real Medicaid expenditure per disabled SSI recipient is more than $400 higher than for an elderly SSI recipient, and more than $800 higher than a blind SSI recipient. The real Medicaid expenditure also exceeds the maximum annual SSI grant (including the state supplements) by more than $1,000. There are large differences in the benefit levels between SSI recipients and nonrecipients: nonrecipients live in states with higher Medicaid expenditure and substantially higher SSI grants. On the surface, these differences in average expenditure on Medicaid and average SSI benefits would suggest that higher benefits reduce participation. However, other omitted factors, such as attitudes toward welfare participation, vary across states and are surely

[12]All of these variables were obtained from various editions of U.S. House of Representatives, *The Green Book: Overview of Entitlement Programs*. See the Data Appendix for details.

correlated with benefit levels. Two of the instrumental variables—Medicaid expenditure for AFDC children and adults—are nearly identical in value for recipients and nonrecipients. Although both of these levels are substantially smaller than the other Medicaid variables, they turn out to be highly significant in explaining Medicaid expenditure for the disabled. Finally, around one-quarter of the sample live in a 209(b) state, meaning that the SSI participant must file another application in order to receive Medicaid.

The next two rows show the means of economic variables included in the model. Recipients live in states with higher unemployment and large labor force pools. A study by Stapleton, Coleman, and Dietrich (1995) found that changes in the unemployment rate accounted for 10 percent of the growth in SSI applications from 1988 to 1992.

Finally, Table 2 displays several demographic characteristics that are included in the regression analysis. On average, SSI recipients are five years older than nonrecipients. Participants are much more likely to be African American. In addition, SSI recipients are far less educated: 38 percent did not even enter high school, and another 23 percent did not complete high school. In contrast, only 16 percent of nonrecipients did not receive at least a high school diploma. SSI recipients are less likely to be married, be male, have children, or be a veteran. Finally, there are noticeable differences in the take-up (and presumably, availability) of private insurance coverage. Less than 5 percent of SSI recipients had coverage in their own name, compared with 60 percent of nonrecipients.

Table 3 breaks out the trends in SSI participation from 1987 to 1993, for the entire sample and for several demographic groups. For the entire sample, the SSI participation rate rose steadily, from 0.98 percent in 1987 to 1.35 percent in 1993. Perhaps the most striking feature of this table is that the level of participation for the African-American population is more than three times as high as for the white population. The trend for blacks, however, shows no consistent pattern—the

TABLE 3

Trends in SSI and Medicaid Participation over Time

	Entire Sample		African American		White		Men, 18–64		Men, 45–64		Women, 45–64	
SSI Participation												
1987	0.985%	(0.044)	3.078	(0.280)	0.813	(0.041)	0.727	(0.043)	0.992	(0.091)	1.768	(0.118)
1988	1.067	(0.047)	3.846	(0.327)	0.845	(0.044)	0.777	(0.046)	1.080	(0.099)	1.958	(0.129)
1989	1.073	(0.045)	2.812	(0.265)	0.928	(0.044)	0.776	(0.045)	1.114	(0.096)	1.988	(0.125)
1990	1.155	(0.047)	3.559	(0.300)	0.959	(0.045)	0.813	(0.046)	0.956	(0.089)	2.201	(0.131)
1991	1.319	(0.050)	3.511	(0.293)	1.133	(0.049)	0.992	(0.050)	1.346	(0.105)	2.310	(0.134)
1992	1.268	(0.050)	3.346	(0.288)	1.093	(0.048)	0.942	(0.050)	1.178	(0.098)	2.225	(0.131)
1993	1.355	(0.053)	3.550	(0.305)	1.171	(0.051)	1.012	(0.053)	1.239	(0.102)	2.340	(0.136)
Medicaid Participation												
1987	1.966	(0.061)	5.972	(0.384)	1.638	(0.058)	1.706	(0.066)	1.907	(0.126)	2.756	(0.146)
1988	1.919	(0.063)	6.015	(0.404)	1.592	(0.060)	1.608	(0.067)	1.946	(0.133)	2.872	(0.155)
1989	2.070	(0.063)	5.830	(0.376)	1.756	(0.060)	1.715	(0.066)	2.220	(0.135)	3.160	(0.157)
1990	2.397	(0.068)	6.881	(0.411)	2.031	(0.065)	2.065	(0.073)	2.314	(0.138)	3.411	(0.163)
1991	2.495	(0.069)	6.234	(0.385)	2.179	(0.067)	2.177	(0.074)	2.288	(0.137)	3.461	(0.163)
1992	2.601	(0.071)	6.177	(0.386)	2.299	(0.069)	2.253	(0.076)	2.349	(0.137)	3.622	(0.166)
1993	2.666	(0.073)	6.391	(0.404)	2.355	(0.072)	2.282	(0.079)	2.384	(0.141)	3.768	(0.171)
Correlation in Year-to-Year Changes of SSI and Medicaid Participation												
	-0.06		0.56		0.08		-0.02		-0.37		-0.08	

Source: Results from the March 1988–94 Current Population Survey.
Standard errors are in parentheses.

participation rate falls from 3.08 percent in 1987 to 2.81 percent in 1989, and then rises to 3.55 percent in 1993. The trend for whites is clearer: the SSI participation rate increased by more than one-third, from 0.81 percent to 1.17 percent, despite varying economic conditions. The different trends may help to explain the different findings for whites and African Americans in the subsequent regression analysis. Participation also varies by gender. The SSI participation rate for women age 45 to 64 is more than one percentage point higher than that for men in the same age range. Both groups show increasing participation over time, however.

As with other household surveys, program participation is underreported in the CPS. The national SSI participation rate for the adult population was 1.75 percent in 1992, compared to 1.27 percent in the CPS.[13] While participation rates also appear to be underreported at the state level, the gap between administrative data and the CPS varies. The participation rate in 1992 is underreported by 0.07 percentage points in Florida, by between 0.32 to 0.48 percentage points in Illinois, New York, and Texas, and by 0.95 percentage points in California. If the stigma of reporting SSI participation to a survey taker varies by state, then underreporting patterns would emerge.

The bottom half of Table 3 shows how Medicaid participation evolved over time. For the pooled sample and each demographic group, Medicaid participation exceeds SSI participation. As in the top half of the table, participation is highest among African Americans. The link between Medicaid and SSI is fairly tight for African Americans—the correlation coefficient between the year-to-year changes in SSI participation and changes in Medicaid participation is 0.56. For other groups, however, the changes are only weakly correlated. In fact, there is a fairly strong negative correlation in the year-to-year changes for older men.

[13]U.S. House of Representatives (1993).

Table 4 portrays trends in real SSI benefits and real Medicaid expenditure from 1987 to 1993.[14] The potential SSI benefit is computed from the CPS, based on the respondent's state of residence, year, and marital status. Clearly, two different patterns emerge here. Real SSI cash benefits remain unchanged. This is expected since the federal benefit level is indexed for inflation. Medicaid expenditure for the disabled increased more than $2,700 in real terms. This rise in Medicaid expenditure looks similar to the rise in overall SSI participation rates; thus, this is at least suggestive that a link between Medicaid expenditure and SSI may exist.

5. REGRESSION RESULTS

OLS Estimates Using Average Medicaid Expenditure of Disabled

For ease of interpretation, I present results from a linear probability model. The coefficients from the models below therefore may be interpreted as percentage point changes. The basic equation is given by:

$$\text{(2)} \quad SSI\_PART_i = \beta_0 + \beta_1 MEDICAID\_BEN_{ijt} + \beta_2 SSI\_BEN_{ijt} + \beta_3 X_i + \beta_4 ECONOMIC_{ijt} + \sum_j \delta_j S_{ij} + \sum_t \delta_t T_{it} + \varepsilon_i.$$

In this equation, i subscripts individuals, j subscripts states, and t subscripts time periods. The outcome, SSI participation (SSI_PART), is a binary variable equal to 1 if the respondent participated in the program in the previous year. Increases in two key policy variables, average real Medicaid expenditure (MEDICAID_BEN) and the average real SSI benefit (SSI_BEN), are expected to increase SSI participation.[15] The business cycle variables (ECONOMIC) include the annual unemployment rate

[14]These are deflated using the Consumer Price Index-Urban for the SSI benefit level and the medical services CPI for Medicaid.

[15]I include a third state-specific variable, whether or not the respondent lived in a section 209(b) state. Several states changed status, but in models with state-fixed effects, this effect is never reliably estimated. The coefficient should be negative—living in a state with extra application procedures for Medicaid increases transaction costs and thus lowers SSI participation.

TABLE 4

Trends in SSI Benefits and Medicaid Expenditure

	Annual SSI Benefit		Average Medicaid Expenditure	
1987	$7,211	(9.81)	$6,700	(22.84)
1988	$7,074	(9.45)	$6,482	(12.11)
1989	$7,163	(9.81)	$7,771	(13.06)
1990	$7,090	(9.67)	$8,308	(13.87)
1991	$7,112	(9.33)	$8,607	(13.14)
1992	$7,133	(9.20)	$9,730	(15.93)
1993	$7,218	(8.51)	$9,491	(17.04)

Source: Results from the March 1988–94 Current Population Survey.
Standard errors in parentheses.
All values are in 1990 dollars.

in the state and the number participating in the labor force.[16] The vector X_i contains several individual-level variables that may also influence SSI participation, including the respondent's age and its square, race, residence in a central city, education, marital status, number of children present, gender, and veteran status.[17] In addition, I amend this basic specification to allow for nationally uniform, time-varying shocks to SSI participation through the inclusion of five time dummies, as well as time-invariant, state-specific shocks to SSI participation through the inclusion of forty-nine state dummies.[18] The coefficients β_0, β_1, β_2, β_3, β_4, δ_j, and δ_t are to be estimated, and ε_i is an error term.

By including time fixed effects (T_{it}), the regression framework accounts for some of the other factors that may lead to an increase in SSI participation. I am able to control for the effects of the business cycle (at the national level) with the time dummies. Since Stapleton, Coleman, and Dietrich (1995) demonstrate that the business cycle influences DI participation, its influence on SSI participation may be expected. If changing economic conditions are correlated with Medicaid expenditure, the results will be biased by not accounting for this omitted variable.

Time dummies also control for two other sources of SSI growth. First, SSI spell lengths may have increased in duration because the Social Security Administration was performing fewer disability reviews. Rupp and Scott (1995) find that persons who were first awarded SSI disability benefits in 1993 are expected to stay on the rolls for almost 18 years. This is an increase from 11 years for those who

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[16]While employment outcomes may be determined simultaneously with SSI participation, I include these variables to make comparisons with Stapleton, Coleman, and Dietrich (1995). The effects of Medicaid, both in the OLS and TSLS models, get much stronger by excluding these variables.

[17]I include many of the same demographic variables that Winkler (1991) includes in her model on female heads. I have modified all specifications by replacing the respondent's age and its square with a full set of dummy variables for ages 18 to 64. In addition, I have restricted the sample to adults aged 22 to 64 since some rules that govern the SSI eligibility for a child who reaches the ages of 18 to 21 have changed. The results are similar to those reported here.

[18]In all the models presented, the joint significance of the time dummies and state dummies is overwhelming.

were initially awarded benefits in 1974. Second, some medical breakthroughs may have allowed disabled people to live longer than they otherwise would have.[19]

Several unmodeled or unobserved variables, which vary by state, could bias the results. Including state-fixed effects (S_{ij}) in the regression addresses these concerns. First, SSI reporting behavior in the CPS data varies. If admitting program participation represents permanent differences in attitudes that vary by state, then including state-fixed effects will account for these differences. Second, the availability of Medicaid coverage varies across states and this could affect SSI participation. For instance, a poor adult may be able to receive health insurance coverage through the Medically Needy (MN) or General Assistance (GA) programs. These programs may be correlated with Medicaid expenditure and also affect SSI participation. More liberal states may have these optional programs, which reduces SSI participation. They may also have more generous Medicaid services, which increases average Medicaid expenditure. This scenario would likely lead to the conclusion that increased Medicaid expenditure reduces SSI participation. If the MN and GA programs remain fixed within a state, then this heterogeneity would be accounted for with state-fixed effects.[20]

The result in the first column of Table 5 shows that the OLS estimate of β_1 is statistically insignificant and economically modest. The point estimate implies that increasing Medicaid by $1,000 leads to an increase in SSI participation of 0.009 percentage points. Since Table 4 illustrates that average Medicaid expenditure for the entire sample rose in real terms from $6,700 in 1987 to $9,491 in 1993, this coefficient estimate implies that increased Medicaid expenditure raised the probability of

[19]Ross (1995).

[20]The programs may have changed over time, however. Several states eliminated their GA program in the early 1990s. If GA cuts are correlated with Medicaid expenditure, then even the model that includes state- and time-fixed effects will be biased. In other specifications, I have also included a state-specific time trend in equation (2) to account for such changes. The basic conclusions remain unchanged, but the standard error on β_1 increases by 33 percent.

TABLE 5

Linear Probability Model from Full CPS Sample on SSI Participation

	(1) OLS		(2) IV	
Medicaid benefit/10^6	0.0961	(0.0897)	0.3896	(0.1380)
SSI benefit/10^6	0.3242	(0.2949)	0.3306	(0.2949)
Unemployment rate	4.6405	(2.0643)	3.5174	(2.1030)
State's labor force participation	-0.0033	(0.0011)	-0.0027	(0.0011)
Section 209(b) state	0.0039	(0.0027)	0.0041	(0.0027)
Respondent's age	0.0017	(0.0001)	0.0017	(0.0001)
$Age^2/100$	-0.0015	(0.0001)	-0.0015	(0.0001)
African American	0.0161	(0.0007)	0.0161	(0.0007)
Resides in central city	0.0018	(0.0004)	0.0019	(0.0004)
Education<9	0.0624	(0.0007)	0.0624	(0.0007)
9≤ Education<12	0.0206	(0.0006)	0.0206	(0.0006)
Education=12	0.0047	(0.0004)	0.0047	(0.0004)
Currently married	-0.0295	(0.0009)	-0.0295	(0.0009)
Number of own children under age 6	0.0041	(0.0003)	0.0041	(0.0003)
Number of own children aged 6 to 17	0.0001	(0.0002)	0.0001	(0.0002)
Male	-0.0045	(0.0005)	-0.0045	(0.0005)
Veteran	-0.0039	(0.0005)	-0.0039	(0.0005)
Observations	345,453		345,453	
Mean of dependent variable	0.0117		0.0117	
Adjusted R^2	0.0421		0.0420	

Source: Results from the March 1988–94 Current Population Survey.
Standard errors in parentheses.
All models also include state fixed effects (49), time fixed effects (5), and a constant term. Instruments in column (2) are average Medicaid expenditure for blind, elderly, AFDC children, and AFDC adults.

SSI participation by 0.026 percentage points. Since SSI participation for the whole sample increased from 0.98 to 1.35 percent (or 0.370 percentage points), the OLS estimate implies that rising health care costs can explain around 7 percent of the rise in SSI participation.

Increasing the SSI cash benefit increases SSI participation, but the coefficient is insignificant. Raising the benefit by $1,000 results in an increase in SSI participation of 0.032 percentage points. Although the coefficient suggests SSI cash benefits may explain the rise in participation, Table 4 shows little change in cash benefits over time. The CPS estimates indicate that, over the period, SSI benefits increased very slightly in real terms from $7,211 to $7,218. Even though cash benefits increase the probability of participation, they cannot explain the growth. The table also shows the effect of a third policy variable: residence in a Section 209(b) state—one that requires a second, separate application for Medicaid. Since very few states changed status, the effect of 209(b) status is essentially subsumed in the state fixed effect. The estimate in this column is not significantly different from zero.

The business cycle variables enter in the appropriate direction. In the OLS specification, increases in the unemployment rate raise SSI participation, but the significance is somewhat fragile in the other specifications. Increases in labor force participation have a more robust, negative effect on SSI participation.

Education and family structure play important roles in SSI participation. Relative to those with a college degree, individuals with less than nine years of education are 6.2 percentage points more likely to participate in SSI, while those with less than twelve years are 2.0 percentage points more likely to participate. In addition, those who only completed high school are significantly more likely to participate in the SSI-disabled program than those who entered college, but the economic impact is not as dramatic as for the other educational groups. Being married lowers SSI participation by 3 percentage points, while the presence of another young child increases the probability of participation.

The signs of the other demographic and location-specific characteristics enter into SSI participation largely as expected. SSI participation increases with age, but at a decreasing rate. The point estimates indicate that SSI participation increases all the way up until age 64. Since many physical disabilities may not occur until later ages, this finding is plausible. Relative to whites, being African American raises the probability of SSI participation by 1.61 percentage points. This is consistent with the continually higher levels of participation in Table 3. Living in a central city raises SSI participation. This may occur for two reasons. First, those in central cities may have more access to welfare and Social Security offices or health care facilities, which lowers the transaction costs of SSI participation and raises the value of Medicaid, respectively. Second, if living in a central city means that the individual has better information about the programs, he would be more likely to participate. Finally, being male or being a veteran significantly lowers SSI participation.

Instrumental Variables Estimates Using Average Medicaid Expenditure of Other Medicaid Groups as Instruments

The prior estimates used variation in disabled Medicaid expenditure. These may be biased if changes in the underlying health of the SSI population affected both Medicaid's value and SSI participation. If the eligibility criteria for disability become less strict, for example, so that people who were previously found to be ineligible are now deemed eligible for SSI, then the former estimates of β_1 would be too small. In the Supreme Court's *Sullivan v. Zebley* decision, such a reevaluation occurred for children, and this may have had spillovers into the adult population.[21] In addition, if states attempted to shift their General Assistance and Medically Needy beneficiaries onto the SSI rolls, and if these groups happened to be healthier, the OLS results would be biased. In this case, the marginal disabled

[21]The Supreme Court ruled that disability standards for children may not be narrower than those applied to adults. As a result, eligibility criteria for children are based on a child's developmental delay and limitations on the child's ability to engage in age-appropriate activities of daily living. This has increased the number of children classified as disabled. Prior to 1990, the same disability criteria that applied to adults were also applied to children.

SSI recipient will likely incur less health care expenditure than the average recipient, so that average expenditure falls while SSI participation increases. This would lead to a spurious negative correlation (which in turn biases the coefficient downward).

To correct for this simultaneity bias, I instrument for average Medicaid expenditure of the disabled in each state-year cell with the corresponding average expenditure of the elderly and blind SSI recipients, and adult and child AFDC recipients. These variables reflect different aspects of the state's Medicaid program that influence its value, such as variation in health care prices, access to care, and scope of services. Since the criteria to qualify for these other groups are more objective, then these instruments are unlikely to be correlated with changing definitions of disability.[22]

Appendix Table 1 presents the results from the first-stage regression. These instruments are highly correlated with the average Medicaid expenditure of the disabled: both the coefficients on expenditure for the blind and expenditure for the disabled have t-statistics over 200, while the coefficients on adult and child AFDC recipients have t-statistics over 70. The first stage F statistic is over 22,000 and the R^2 is 0.8317.[23] By instrumenting, the coefficient estimate in the second column of Table 5 increases dramatically, consistent with changing the budget constraint in Figure 1. Increasing Medicaid expenditure by $1,000 is now associated with an increase in the probability of SSI participation by 0.039 percentage points. Again, taking the rise in Medicaid expenditure from Table 4, this estimate implies that rising health care costs from 1987 to 1993 raised the probability of participation by 0.108 percentage points. Since the total increase in SSI participation was 0.370

[22]Persons 65 and over with limited income and resources can qualify under the aged SSI program, while blind individuals are defined as those with 20/200 vision or less with the use of a correcting lens in their better eye, or those with tunnel vision of 20 degrees or less. Adults and children qualify for AFDC by living in a single-parent family.

[23]Bound, Jaeger, and Baker (1995) explain that in finite samples, IV estimates are biased in the same direction as OLS estimates, and the magnitude of the bias of the IV estimates approaches that of OLS estimates as the R^2 between the instruments and the potentially endogenous explanatory variable approaches zero.

percentage points, then rising health care costs can explain nearly one-third of the rise in SSI participation. The point estimates on the other explanatory variables are similar to the OLS specification. By comparing the coefficient estimates on Medicaid expenditure and SSI benefit levels, a $1,000 increase in Medicaid's value leads to a similar rise in participation as a $1,178 increase in cash.

Recall that Table 3 showed dramatic differences in SSI participation rates across racial lines. This suggests that rising health care costs have different effects on the African-American and white populations. The two columns in Table 6 divide the sample into whites and African Americans, respectively. I use the same instruments as in Table 5.

The Medicaid coefficient estimates for the white population are roughly similar to the TSLS estimates from the second column of Table 5. The effect of Medicaid expenditure falls slightly, and the coefficient is more precisely estimated than in the full sample. Cash benefits appear to play a less important role for whites in SSI participation than for the pooled sample. In contrast, Medicaid appears to play little role in the SSI participation decision of African Americans, though the coefficient is imprecisely estimated. Although the policy variables explain little of the SSI participation decision for African Americans, the demographic variables on education, family structure, gender, and veteran status are all significant predictors of participation.[24]

Even though the proposed instruments are powerful in predicting Medicaid expenditure for the disabled, to be valid they must also be uncorrelated with the error term. In practical terms, this means that the health status of the other SSI and AFDC groups remains unchanged. Since the number of instruments exceeds the number of endogenous variables, it is possible to test the overidentifying restrictions on the excluded instruments. Table 7 shows this. I regress the predicted residuals from the

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[24]The finding of weak effects on policy variables for African Americans is quite common. For two recent examples, see Lundberg and Plotnick (1995) and Bronars and Grogger (1996).

TABLE 6

Differences in Medicaid's Impact Based on Race

	(1) IV, White		(2) IV, African American	
Medicaid benefit/10^6	0.3887	(0.1295)	0.5952	(1.1103)
SSI benefit/10^6	0.0816	(0.2807)	2.1081	(2.1185)
Unemployment rate	3.6056	(2.0072)	3.0846	(13.8530)
State's labor force participation	-0.0016	(0.0011)	-0.0203	(0.0081)
Section 209(b) state	0.0037	(0.0025)	0.0400	(0.0389)
Respondent's age	0.0018	(0.0001)	-0.0001	(0.0005)
$Age^2/100$	-0.0016	(0.0001)	0.0012	(0.0007)
Resides in central city	0.0014	(0.0004)	0.0043	(0.0025)
Education<9	0.0573	(0.0007)	0.1144	(0.0045)
9≤ Education<12	0.0187	(0.0006)	0.0386	(0.0032)
Education=12	0.0044	(0.0003)	0.0138	(0.0025)
Currently married	-0.0260	(0.0009)	-0.0558	(0.0061)
Number of own children under age 6	0.0031	(0.0003)	0.0144	(0.0024)
Number of own children aged 6 to 17	-0.0002	(0.0002)	0.0023	(0.0014)
Male	-0.0031	(0.0005)	-0.0252	(0.0032)
Veteran	-0.0034	(0.0004)	-0.0054	(0.0030)
Observations	319,049		26,404	
Mean of dependent variable	0.0099		0.0337	
Adjusted R^2	0.0362		0.0664	

Source: Author's results from the March 1988–94 Current Population Survey.
Standard errors in parentheses.
In addition to the coefficients shown, all models include state-fixed effects (49), time-fixed effects (5), and a constant term. Instruments in columns (1) and (2) are average Medicaid expenditure for blind, elderly, AFDC children, and AFDC adults.

TABLE 7

Overidentification Tests

(1) Model	(2) N	(3) Unadjusted R^2 from regression of predicted ε_2 on Z and X	(4) Test Statistic	(5) 95% χ^2 Critical Value
Full Sample (Table 5) 345,453	0.0000129	4.46	7.82	
Whites only (Table 6)	319,049	0.0000151	4.81	7.82
African American only (Table 6)	26,404	0.0000844	2.23	7.82
Private Coverage (Table 8) 204,091	0.0000631	12.88	7.82	
No Private Coverage (Table 8) 141,362	0.0000219	3.10	7.82	

Source: Results from March 1988–94 Current Population Survey Annual Demographic File.
In all models, instruments include average Medicaid expenditure for SSI blind, SSI elderly, AFDC adults, and AFDC children.

second stage on all the instruments and exogenous variables. The test statistic is calculated as $N*R^2$, the product of the number of observations and the uncentered R^2 from the regression of the predicted residuals on the exogenous variables and instruments. The test statistic is distributed as χ^2 with degrees of freedom equal to the number of overidentifying restrictions, in this case three. In four of the five models, this test statistic is smaller than the 95 percent critical value of 7.82, supporting the claim of exogeneity of the instruments.

Estimates Using Variation in Private Insurance Status

In this section, I see whether Medicaid's effect varies in sensible ways for two particular groups. As Blank (1989) notes, the size of the Medicaid notch depends upon the availability of private insurance. If an individual has private insurance, there is no Medicaid notch. I compare Medicaid's effect for those with and without private health insurance in their own name.[25] The portion who have private insurance serve as the "control" group. The other administrative barriers and benefits from SSI participation should be similar, but the value of Medicaid greatly differs. If Medicaid expenditure is simply capturing some omitted factor, we might expect to see similar effects across these groups.

Table 8 shows the TSLS results for these groups. The results are consistent with a "true" effect of Medicaid. The coefficient on Medicaid expenditure is extremely strong in column (1) for those lacking private coverage, while it is not significant for those with private coverage.

6. CONCLUSIONS AND EXTENSIONS

This paper finds that rising health insurance costs are an important reason for participation in the SSI-disabled program. By using a large, nationally representative household data set, I find that as

[25]This is in the spirit of one of Madrian's (1994) tests for job lock—comparisons of job mobility for those with and without employer provided health insurance.

TABLE 8

Effects of Medicaid on SSI Participation, Stratified by Private Health Insurance Status

	(1) IV, No private coverage		(2) IV, Private coverage	
Medicaid benefit/10^6	0.9228	(0.3250)	0.0261	(0.0478)
SSI benefit/10^6	1.6287	(0.6678)	0.0369	(0.1054)
Unemployment rate	4.2729	(4.9234)	1.0291	(0.7325)
State's labor force participation	-0.0047	(0.0026)	-0.0005	(0.0004)
Section 209(b) state	0.0061	(0.0064)	0.0007	(0.0009)
Respondent's age	0.0056	(0.0002)	-0.0001	(0.0001)
$Age^2/100$	-0.0047	(0.0002)	0.0001	(0.0001)
African American	0.0239	(0.0015)	0.0014	(0.0002)
Resides in central city	0.0029	(0.0011)	-0.0001	(0.0001)
Education<9	0.0826	(0.0015)	0.0041	(0.0003)
9$\leq$ Education<12	0.0290	(0.0013)	0.0011	(0.0002)
Education=12	0.0088	(0.0009)	0.0002	(0.0001)
Currently married	-0.0803	(0.0022)	-0.0015	(0.0003)
Number of own children under age 6	0.0122	(0.0009)	0.0003	(0.0001)
Number of own children aged 6 to 17	-0.0015	(0.0005)	0.0001	(0.0000)
Male	-0.0032	(0.0013)	-0.0005	(0.0002)
Veteran	-0.0090	(0.0013)	-0.0003	(0.0001)
Observations	141,362		204,091	
Mean of dependent variable	0.0275		0.0023	
Adjusted R^2	0.0744		0.0021	

Source: Author's results from the March 1988–94 Current Population Survey.
Standard errors in parentheses.
In addition to the coefficients shown, all models include state-fixed effects (49), time-fixed effects (5), and a constant term.

much as one-third of the rise in SSI participation may be due to increases in the value of Medicaid. The effects appear to be concentrated in the white population, not the African-American population.

I show that ordinary least squares produce a badly biased estimate since the health status of the disabled population is changing. The estimates using instrumental variables produce much stronger positive effects of Medicaid on SSI participation. Is it reasonable to assume that the health status of the disabled changed so dramatically while the health status of other SSI and AFDC recipients did not? Knowing the answer to this question is vital for assessing the validity of the instruments. Since the model is overidentified, I am able to look at the internal consistency of the model. The overidentification test statistics are well below their 95 percent critical values. Moreover, it is difficult to believe that the health status of the blind changed dramatically from 1987 to 1993, and the TSLS results do not change markedly by only using the Medicaid expenditure for the blind as an instrument. On the other hand, the health status of the elderly on SSI may have changed because the Qualified Medicaid Beneficiary (QMB) program in the 1980s and 1990s offered an incentive for them to leave SSI and still retain Medicaid. Around 1.4 million elderly were enrolled in this program in December 1992; however, it is not known whether the health status of former SSI recipients who left and enrolled in the QMB program was better or worse than the average SSI recipient. The same argument could be made for AFDC children and adults—other Medicaid program changes may have affected their health status.

Are the estimated effects too large? At this point, it is important to consider recent findings on other Medicaid populations. In other work, I found significant effects on AFDC participation for female heads and on SSI participation for elderly households (Yelowitz 1995a, 1995b). In those studies, the policy experiment was somewhat different from in this study, however. The policy changes for young children and the elderly who are offered Medicaid without the need to apply for AFDC or SSI. In an approach more similar to the current study, Moffitt and Wolfe (1992) value Medicaid and find strong effects on AFDC participation for female heads. It is plausible that health insurance plays a more important role in the economic decision-making of disabled adults than either female heads or elderly households, so the strong effects appear reasonable.

The findings have several policy implications for program design. If Medicaid is an important determinant of SSI participation, then offering health insurance without the need to participate in SSI may reduce total costs. This could occur because disabled adults may then forego the cash benefits from SSI. On the other hand, some disabled adults who were not previously participating in SSI because of the stigma they associate with the program may decide to participate in a Medicaid-only program, which could increase costs. To some extent, this might occur through the Medically Needy program, which many states offer. Since the MN program typically has lower income limits than SSI and fewer covered services under Medicaid than for categorically needy recipients, it may not offer enough of an incentive for the disabled to leave.

Perhaps the most useful extension of the current study would be to develop a model that includes a more broadened look at the effects of health on SSI participation, along the lines of Wolfe and Hill (1995). This would be important for two reasons. First, by examining a data set with better measures of health status (such as the SIPP), I could directly test the hypothesis that SSI recipients became healthier. Second, by incorporating health directly into the SSI participation equation, it may be possible to see which type of disabled person responds to different government policies concerning extension of Medicaid benefits.

Data Appendix: Sources

A. *Current Population Survey:* Table 1 shows the sample selection criteria.

B. *Medicaid expenditure data:*

Fiscal Year	Source
1987	U.S. House of Representatives, *Overview of Entitlement Programs*, 1989, pp. 1150–51.
1988	U.S. House of Representatives, *Overview of Entitlement Programs*, 1990, pp. 1302–1303.
1989	U.S. House of Representatives, *Overview of Entitlement Programs*, 1991, pp. 1435–36.
1990	U.S. House of Representatives, *Overview of Entitlement Programs*, 1992, pp. 1670–71.
1991	U.S. House of Representatives, *Overview of Entitlement Programs*, 1993, pp. 1664–65.
1992	U.S. House of Representatives, *Overview of Entitlement Programs*, 1994, pp. 811–12.
1993	U.S. Department of Health and Human Services, *Medicaid Statistics: Program and Financial Statistics 1993*, pp. 45–46, 64–65.

C. *SSI benefit data:*

U.S. House of Representatives, *Overview of Entitlement Programs,* 1993, pp. 824, 829–30.
U.S. House of Representatives, *Overview of Entitlement Programs,* 1991, pp. 741–42.

D. *Unemployment rate and labor force participation:*

Bureau of Labor Statistics. Local Area Unemployment Statistics—annual measures of the unemployment rate and the total civilian labor force. Available by World Wide Web at http://stats.bls.gov:80/lauhome.htm

E. *Price Indices for general inflation and medical prices:*

Council of Economic Advisers, *Economic Report of the President* , 1995. Table B-61, p. 344.

Year	1987	1988	1989	1990	1991	1992	1993
CPI-U	113.6	118.3	124.0	130.7	136.2	140.3	144.5
Medical CPI	130.0	138.3	148.9	162.7	177.1	190.5	202.9

F. *Medicaid 209(b) status:*

U.S. House of Representatives, *Overview of Entitlement Programs,* 1993, p. 1635.
U.S. House of Representatives, *Overview of Entitlement Programs,* 1992, p. 1642.
U.S. House of Representatives, *Overview of Entitlement Programs,* 1991, p. 1406.
U.S. House of Representatives, *Overview of Entitlement Programs,* 1990, p. 1278.
U.S. House of Representatives, *Overview of Entitlement Programs,* 1989, p. 1129.
U.S. House of Representatives, *Overview of Entitlement Programs,* 1988, p. 798.

APPENDIX TABLE 1

First-Stage Estimates for Table 3, Column (2)

	(1) Medicaid benefit for disabled	
Medicaid benefit for elderly	0.5734	(0.0022)
Medicaid benefit for blind	0.3316	(0.0012)
Medicaid benefit for AFDC children	1.0173	(0.0154)
Medicaid benefit for AFDC adults	0.7802	(0.0099)
Observations	345,453	
F(75,345377)	22,769	
Adjusted R^2	0.8317	

Source: Author's results from the March 1988–94 Current Population Survey.
Standard errors in parentheses.
In addition to the coefficients shown, model includes state-fixed effects (49), time-fixed effects (5), a constant term, the SSI benefit, unemployment rate, labor force participation, 209(b) status, age and its square, race, education, central city residence, marital status, number of children, gender, and veteran status.

References

Blank, Rebecca. 1989. “The Effect of Medical Need and Medicaid on AFDC Participation.” *Journal of Human Resources* 24:1 (Winter): 54–87.

Bound, John, David Jaeger, and Regina Baker. 1995. “Problems with Instrumental Variables Estimation When the Correlation between the Instruments and the Endogenous Explanatory Variable Is Weak.” *Journal of the American Statistical Association* 90:430 (June): 443–50.

Bronars, Stephen, and Jeffrey Grogger. 1996. “The Effect of Welfare Payments on the Marriage and Fertility Behavior of Unwed Mothers: Results from a Twins Experiment.” Mimeo, University of California, Santa Barbara.

Council of Economic Advisers. 1995. *Economic Report of the President*. Washington, D.C., Government Printing Office, February.

Currie, Janet, and Jonathan Gruber. 1994. “Saving Babies: The Efficacy and Cost of Recent Expansions in Medicaid Eligibility for Pregnant Women.” NBER Working Paper No. 4644.

Lundberg, Shelly, and Robert Plotnick. 1995. “Adolescent Premarital Childbearing: Do Economic Incentives Matter?” *Journal of Labor Economics* 13:2 (April): 177–200.

Madrian, Brigitte. 1994. “Employment-Based Health Insurance and Job Mobility: Is There Evidence of Job-Lock?” *Quarterly Journal of Economics* 109:1 (February): 27–54.

Moffitt, Robert. 1983. “An Economic Model of Welfare Stigma.” *American Economic Review* 73:5 (December): 1023–35.

Moffitt, Robert, and Barbara Wolfe. 1992. “The Effect of the Medicaid Program on Welfare Participation and Labor Supply.” *Review of Economics and Statistics* 74:4 (November): 615–26.

Ross, Jane. 1995. "Social Security: Federal Disability Programs Face Major Issues." GAO/T-HEHS-95-97. Statement of Jane Ross, Director, Income Security Issues, Health Education and Human Services Division.

Rupp, Kalman, and Charles Scott. 1995. "Determinants of Duration on the Disability Rolls and Program Trends." Mimeo, Social Security Administration, July.

Stapleton, David, Kevin Coleman, and Kimberly Dietrich. 1995. "Demographic and Economic Determinants of Recent Application and Award Growth for SSA's Disability Programs." Mimeo, Lewin-VHI, Inc. July.

U.S. Department of Health and Human Services. 1994. *Medicaid Statistics: Program and Financial Statistics, Fiscal Year 1993.* October. HCFA Pub. No. 10129.

U.S. House of Representatives. Various years. *The Green Book: Overview of Entitlement Programs: Background Materials and Data on Programs Within the Jurisdiction of the Committee on Ways and Means.* Washington, D.C.: Government Printing Office.

Winkler, Anne. 1991. "The Incentive Effects of Medicaid on Women's Labor Supply." *Journal of Human Resources* 26:2 (Spring): 308–37.

Wolfe, Barbara, and Steven Hill. 1995. "The Effect of Health on the Work Effort of Single Mothers." *Journal of Human Resources* 30:1 (Winter): 42–62.

Yelowitz, Aaron. 1995a. "The Medicaid Notch, Labor Supply, and Welfare Participation: Evidence from Eligibility Expansions." *Quarterly Journal of Economics* 110:4 (November): 909–39.

Yelowitz, Aaron. 1995b. "Using the Qualified Medicare Buy In Program to Estimate the Effect of Medicaid on the SSI Participation of the Elderly." Mimeo, University of California, Los Angeles.