DOES PUBLIC INSURANCE CROWD OUT PRIVATE INSURANCE?*
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The cost of expanding public sector health programs depends critically on the extent to which public eligibility will cover just the uninsured, or will crowd out existing private insurance coverage. We estimate the extent of crowd-out arising from the expansions of Medicaid to pregnant women and children over the 1987—1992 period. We estimate that approximately 50 percent of the increase in Medicaid coverage was associated with a reduction in private insurance coverage. This occurred largely because employees took up employer-based insurance less frequently. There is also some evidence that employers contributed less for insurance and that workers dropped coverage of dependents.

The recent debate over health care reform in the United States effectively ended public discussion of universal health insurance coverage. As a result, much attention is likely to be paid to partial alternatives, which offer free or subsidized coverage to particular groups. For example, a key feature of "compromise" reform plans is subsidized insurance coverage for the poor, and one option that remains politically popular is free or subsidized coverage for children [Boston Globe, September 12, 1994].

Partial coverage is not without its pitfalls, however. Perhaps the most important problem is the potential for public expansions to crowd out private insurance coverage. As the eligibility for public coverage expands, individuals may drop their private coverage and switch into public programs. This may substantially reduce the "bang for the buck" from public insurance coverage, as the costs of the program rise without commensurate increase in insurance coverage. This concern has been raised in discussions of coverage expansions [New York Times, August 21, 1994], but there is essentially no empirical evidence on its magnitude.\footnote{This is related to the general question of the crowd-out of private insurance mechanisms by social insurance programs: theoretical analyses of asset-tested social insurance programs typically suggest that they will crowd out private savings [Hubbard, Skinner, and Zeldes 1995], and substantial crowd-out has been estimated for the case of Social Security and private savings [Feldstein 1974], unemployment insurance and private savings [Engen and Gruber 1995], and welfare and family transfers [Schoeni 1994].}

*We are grateful to Janet Currie and Aaron Yelowitz for their assistance in developing the Medicaid Eligibility programs used here, to Linda Blumberg, Janet Currie, Paul Gertler, Lawrence Katz, Charles Nelson, Joseph Newhouse, and two anonymous referees for helpful comments, and to the National Institute of Aging for financial support.

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Our goal in this paper is to provide evidence on the effect of public health insurance eligibility on private health insurance coverage. We do so by taking advantage of the large increases in eligibility for Medicaid—the public insurance program for low-income persons—that occurred in the late 1980s and early 1990s. Historically, eligibility for Medicaid was tied to the receipt of cash welfare payments under the Aid to Families with Dependent Children program. Hence, eligibility was effectively limited to very low-income women and children in single-parent families. During the late 1980s and early 1990s, states were first permitted and then required to extend Medicaid coverage to pregnant women and children with higher incomes, and in families with two parents. As a result, by 1992, almost one-third of children were eligible for full public coverage, and over 40 percent of women of child-bearing age could receive public coverage for pregnancy-related expenses.

We consider the effect of these Medicaid expansions on coverage by Medicaid, private insurance coverage, and the rate of uninsurance. We do so in four steps. First, we calculate the extent of crowd-out for children and women of child-bearing age assuming that their insurance decisions depend only on their own Medicaid eligibility. Our estimates for children suggest that 31 percent of the increases in Medicaid coverage due to these expansions were offset by lower private insurance coverage. Our estimates for women are even higher but less precise.

Second, we investigate the mechanisms through which crowd-out might occur. We find no evidence that crowd-out is caused by employers choosing not to offer health insurance as eligibility rises, although there may have been some increase in the share of premiums employers require employees to pay. Rather, crowd-out appears to arise from employees not taking up coverage when it is offered. In addition, for those workers retaining their coverage, there is a move toward dropping coverage of dependents and covering oneself only.

We then use this information to estimate the amount of crowd-out allowing for the Medicaid eligibility of some members of the family to influence the insurance decisions of other members of the family. Accounting for these effects increases our estimated crowd-out, to 77 percent of the coverage increase.

This estimate is too high, however, because not everyone who enrolls in Medicaid does so on a continuous basis. Many people are "conditionally covered" for Medicaid: they remain uninsured until they become ill (children) or pregnant (women), and then
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sign up for coverage. These people are officially counted as uninsured but in fact have some conditional insurance through the public sector. Accounting for this conditional coverage reduces the amount of crowd-out to our final estimate of 49 percent.

The paper proceeds as follows. In Section I we discuss theoretically the effect of public insurance expansions on private coverage. Section II provides background on the Medicaid expansions. Section III presents a first look at the crowd-out question. Section IV considers how Medicaid eligibility has affected workers. Section V then calculates the net change in insurance coverage induced by Medicaid expansions. Section VI concludes.

I. SHOULD PUBLIC INSURANCE CROWD OUT PRIVATE INSURANCE?

The theoretical effects of public insurance expansions on private coverage can readily be analyzed with the framework introduced by Peltzman [1973] for the case of education. Consider a woman of child-bearing age or a child, deciding on their insurance choice. For simplicity, we assume that insurance is sold individually, and that policies differ only in the comprehensiveness of medical care that is covered. For example, more generous plans offer a greater range of providers or cover a wider set of medical services. People choose between more generous insurance and other goods as shown in Figure I. People valuing insurance highly (i.e., those demanding the highest quality providers) will choose a policy such as $D$, while those valuing insurance less highly will choose a point such as $E$.

Now the government introduces free public insurance with generosity $M$. On paper, Medicaid is a very valuable policy: almost everything is covered, and there is little or no cost sharing. For many reasons, however, the value of Medicaid is below that of private policies. Because of low Medicaid reimbursement rates, providers are often reluctant to treat Medicaid patients [Currie, Gruber, and Fischer 1995], thus reducing the value of coverage. In addition, individuals may not want to be enrolled in public programs, because of the stigma associated with public programs or the difficulty in enrolling. Finally, the value of Medicaid may

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2. We are grateful to an anonymous referee for suggesting this approach.

3. Takeup rates for AFDC, for example, are generally about two-thirds of eligibles [Blank and Ruggles 1991]. Potential Medicaid eligibles must complete lengthy and complex application forms, provide extensive documentation (such as birth certificates, pay stubs, and confirmation of child-care costs) that many eligibles do not have, and attend several interviews with caseworkers. About one-
be low because individuals may have difficulty shifting from Medicaid back into private coverage if they have preexisting medical conditions. We thus show the value of the Medicaid package as below the value of most private policies.

Individuals cannot purchase a supplement to Medicaid (for example, an option to see higher quality doctors by paying more on the margin). Thus, individuals who choose the public sector must consume insurance of exactly the amount $M$. If they want any higher quality insurance, they return to the original budget constraint. The budget constraint with Medicaid is therefore $ABMC$. In response to this public coverage, people with low values of private insurance (such as $E$) will choose to enroll in the public sector, while individuals with a high valuation of insurance (such as $D$) will choose to retain their private insurance.

The key empirical prediction of this model is readily apparent in Figure I. As the value of public coverage rises relative to the underlying demand for insurance quality, then individuals

third to one-half of all Medicaid applications are denied, and half of these denials are for procedural reasons; that is, because the applicant did not complete all of the necessary steps.
will be more likely to drop their private insurance and enroll in Medicaid. Testing this specific prediction is, unfortunately, impossible, since we do not know the desired insurance for any person, and (for reasons we discuss below) observed private policies may not be independent of Medicaid policy. We therefore test a weaker prediction of this model. On average, individuals made eligible for public insurance will reduce their private insurance coverage, relative to groups not eligible for public insurance. While not providing the underlying parameters of the insurance demand function, this responsiveness is an average elasticity among the group made eligible for Medicaid coverage.

Several characteristics of real-world insurance policies complicate, but do not change, this basic story. The first is the fact that health insurance is generally sold only for individuals or families, without gradations among types of dependents. Thus, a family that wants to cover both parents but not the children (because the children may qualify for Medicaid) may find it impossible to do so with only one policy. Similarly, there is often no savings from enrolling some dependents on a policy but not others.

This lack of distinction among dependents may increase or decrease the amount of crowd-out. To the extent that families value coverage of all members and some members cannot qualify for public coverage, crowd-out is likely to be smaller than an individual-by-individual calculation would suggest. On the other hand, if the Medicaid subsidy is large, families may drop coverage of all members, even those who do not qualify for public insurance directly.

These considerations suggest paying particular attention to the composition of coverage within the family in addition to whether a family has coverage at all. We shall return to this in Section IV. They also suggest paying particular attention to the measurement of family eligibility for Medicaid. Since insurance decisions are made at a family level, eligibility of the entire family should affect the private insurance coverage of each member of the family. Figure I suggests a natural way to account for this. For the family with average tastes for insurance, Medicaid generosity can be proxied as the share of expected spending that would be replaced by Medicaid (the Medicaid "replacement rate"). The average family eligibility is the analog to the individual eligibility measure noted above. We employ this measure in our examination of family insurance purchases below.
The second complication is that most private health insurance is provided through employment, rather than purchased individually. This has several implications for the analysis. First, individuals will generally not choose over an entire budget segment, but rather they will have the choice of one or several points along the budget line (corresponding to different health insurance packages offered by employers). Thus, knowing the generosity of an individual's current insurance policy need not indicate the generosity of their desired insurance policy.

Second, workers may not receive the savings from forgoing employer-provided coverage. While empirical evidence suggests that health insurance costs are passed back to workers [Gruber 1994; Sheiner 1994], this research has not established whether this pass-back occurs in response to individual or group choices of insurance. If individual workers do not receive the savings from choosing not to purchase insurance, they will perceive moving to Medicaid as a reduction in health insurance but not as an increase in other consumption. Fewer people will drop private insurance coverage in this case.

This issue is quantitatively important. In 1987 the typical employee paid directly $483 for health insurance, and had out-of-pocket costs for covered services of $779, for a total of $1262. The individual's employer paid $2078 for health insurance, however, or about 62 percent of the total amount. Thus, the extent to which individuals receive the benefits from reduced consumption of insurance is largely dependent on the extent to which employers can raise that individual's wages in response to dropping private coverage.

In the absence of complete wage shifting, employers may encourage workers to drop coverage in other ways. One way to do this is to reduce the generosity of the benefits offered (moving the constrained employer option down the budget constraint), or in the limit, simply to stop offering insurance to the workers (moving them to point C). In either case, these limitations on the private option will make the public option relatively more attractive. Alternatively, employers can reduce the share of the premium

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4. In this situation individuals' indifference curves need not be tangent to the budget line, since the packages offered by employers will reflect a workplace preference function that may not match the worker's underlying tastes. These workplace constraints are binding even given the availability of insurance purchase in the individual market, since insurance purchased on one's own is much more expensive due to both higher loading factors and the loss of the tax subsidy to employer-provided insurance purchase.
that they pay. When employees pay more of the premium, the link between Medicaid receipt and additional income may be more direct (since it does not operate through the veil of shifting to wages). In addition, because there is a tax subsidy for employer spending on insurance but not for individual spending, increasing the share of the premium that employees pay directly effectively raises the price of private insurance relative to Medicaid.

Because of IRS nondiscrimination rules, however, neither of these actions can be used selectively for those workers eligible for public insurance. If insurance is offered, it must be offered to all full-time workers [Cutler and Madrian 1995]. As a result, all of these actions increase the total cost of insurance for employees who do not qualify for public coverage, since they lose the tax subsidy for some insurance purchases, or (if employers drop coverage) they must purchase insurance in the more expensive individual market.5

On net, therefore, the link between health insurance and employment may increase or decrease the amount of crowd-out. If worker-specific shifting is not possible, then crowd-out may be reduced, as employees do not realize the savings from moving to the public sector. If employers increase cost sharing or reduce coverage for all workers, however, more workers may decide to drop coverage than are immediately eligible for Medicaid. The net effect could be in either direction. We explore the extent to which employers have taken actions to reduce the generosity of their benefits in Section IV.

II. MEDICAID ELIGIBILITY AND COVERAGE EXPANSIONS

Historically, Medicaid eligibility for pregnant women and children was tied to participation in the Aid for Families with Dependent Children (AFDC) program. This linkage with AFDC restricted access to the program in two ways. First, AFDC benefits are generally available only to single-parent households.6 Second, income cutoffs for AFDC are very low: generally below

5. This assumes that employee contributions are on a posttax basis, as most are. Note as well that the Medicaid expansions also create incentives for firms to increase outsourcing to temporary help agencies or other specialized low-wage firms (e.g., janitorial services) to get around IRS nondiscrimination rules. See Cutler and Madrian [1995] for more discussion.

6. Married participants in the AFDC-UP program do qualify for Medicaid, but this program is small. In 1990 only 5 percent of the AFDC caseload qualified under AFDC-UP [U. S. House of Representatives 1992].
the poverty line. In 1987 the average income eligibility level across the states was roughly 60 percent of the poverty line.

There were a number of other programs, offered at the discretion of the states, under which poor pregnant women and children could qualify for Medicaid, including the Medically Needy program (which allowed people to qualify for Medicaid if their incomes were somewhat above the cutoff but their medical expenses were particularly high); the Ribicoff option (which allowed states to cover children in two-parent families who met the AFDC income criteria); and other state options to cover pregnant women without children who met the AFDC income criteria. While these options relaxed the family structure restrictions for the program, eligibility was still restricted to only very poor persons.

To expand insurance coverage to the low-income population more broadly, in the late 1980s and early 1990s the linkage between AFDC coverage and eligibility for Medicaid was gradually weakened for pregnant women and children. Currie and Gruber [1994, 1995] describe the legislation accomplishing this in more detail. The expansions substantially increased (in most states) the income that a family could have and still qualify for Medicaid, and at the same time provided these higher eligibility levels to all family structures, not just to single-parent families. By 1992 states were required to cover all pregnant women and children under the age of 6 up to 133 percent of poverty (independent of family composition), and were allowed to expand coverage up to 185 percent of poverty. In addition, children born after September 30, 1983, were mandatorily covered up to 100 percent of poverty (once again independent of family composition).

The key point for our empirical work is that states initially had different qualification limits, and they took up the new options at different rates, so that there was a great deal of variation across states in both the size and timing of the expansions. There was also variation within states in the eligibility of children of different ages, due to different age thresholds in the laws. We use this cross-state and cross-age variation in the size and timing of coverage expansions to identify our crowd-out estimates.

We measure eligibility for Medicaid using data from the March Current Population Surveys (CPS) for 1988 through 1993. Since the CPS asks about insurance and income with a one-year

7. Several states expanded coverage above 185 percent of poverty, but they did not receive federal matching funds for those enrollees.
lag, the data are for 1987 through 1992. We restrict our analysis to the post-1987 period since the insurance questions asked in the March CPS have been consistent in that period. We impute eligibility to these women and to children in the CPS based on algorithms that are described in Currie and Gruber [1994, 1995].

The CPS is the largest nationally representative annual survey with the demographic data necessary to impute Medicaid eligibility and with information on insurance coverage. There are some drawbacks, however. First, the CPS does not indicate whether a woman is pregnant. We therefore impute eligibility for all women of child-bearing age (15–44). Second, the CPS does not have information about assets, which are used in determining Medicaid eligibility. Third, the CPS interviews individuals in March about their insurance coverage in the previous year, so that the questions about insurance coverage may be a mixture of information about the past year and the current situation, rather than a consistent answer to either. Finally, individual respondents in the CPS may not accurately distinguish Medicaid coverage of their family members from the coverage of themselves. Parents whose children are covered by Medicaid, for example, may respond that they too are covered, even if they are not. This might lead to an overstatement of the effects of eligibility expansions on Medicaid coverage, a point we return to below.

Table I shows trends in Medicaid eligibility and coverage for children and women of child-bearing age. For both groups there was a dramatic expansion in eligibility. Eligibility for children rose by over 50 percent, to 27 percent of all children by 1992. For women of child-bearing age, eligibility for Medicaid in the event

8. Since Medicaid coverage is determined on a monthly basis and our data are on an annual basis, some people will naturally be on Medicaid but not determined to be eligible. Given this problem, our Medicaid eligibility imputation is fairly accurate. Of the children on Medicaid in 1992, over three-quarters are imputed to be eligible for Medicaid. Similarly, 95 percent of women of child-bearing age who are on Medicaid were determined to be eligible for Medicaid.

9. Blank and Ruggles [1993] find that eligibility calculations for AFDC do not change much when asset tests are incorporated.

10. This problem is exemplified by the case of adult men who report receiving neither AFDC nor Supplemental Security Income (SSI). For this group there is no possibility of receiving Medicaid, but 1.4 percent of them report Medicaid receipt in 1987, and 2.4 percent report it in 1992. In 99 percent of these cases, someone else in the family also reported receiving Medicaid coverage. This problem might arise from the fact that respondents are first asked if anyone in the family is covered by Medicaid, and later on they are asked about the coverage of individual family members. At this later point, respondents may simply answer that all household members are covered. We are extremely grateful to Charles Nelson of the Census Bureau for discussing these issues with us.
TABLE I
TIME SERIES DATA ON "MEDICAID" ELIGIBILITY AND COVERAGE

<table>
<thead>
<tr>
<th>Year</th>
<th>Percent of children 0-18</th>
<th>Percent of women 15-44</th>
<th>Percent of HIU dollars</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Eligible</td>
<td>Covered</td>
<td>Eligible</td>
</tr>
<tr>
<td>1987</td>
<td>17.8%</td>
<td>14.8%</td>
<td>21.1%</td>
</tr>
<tr>
<td>1988</td>
<td>17.8%</td>
<td>15.1%</td>
<td>25.2%</td>
</tr>
<tr>
<td>1989</td>
<td>18.5%</td>
<td>15.3%</td>
<td>32.3%</td>
</tr>
<tr>
<td>1990</td>
<td>22.7%</td>
<td>18.0%</td>
<td>40.3%</td>
</tr>
<tr>
<td>1991</td>
<td>25.0%</td>
<td>19.9%</td>
<td>42.4%</td>
</tr>
<tr>
<td>1992</td>
<td>27.0%</td>
<td>21.0%</td>
<td>44.9%</td>
</tr>
</tbody>
</table>

Calculations based on CPS data for 1988 through 1993. The columns for women aged 15-44 combine eligibility for all expenditures with eligibility for pregnancy-related services only. Women aged 15-18 are included in the child sample if dependent and in the adult sample if not dependent. Men aged 15-18 are included in the child sample if dependent. Data are weighted to national totals.

of pregnancy doubled over this six-year period, and reached 45 percent of women by 1992. As the second and fourth columns indicate, coverage also increased substantially, although by less than eligibility. Coverage of children rose by six percentage points, and coverage of women of child-bearing age rose by three percentage points.

Most of the increase in eligibility in this period was due to the expansions beyond AFDC, and not changes in AFDC-related eligibility. For each child and woman of child-bearing age in 1987, we estimated eligibility under the rules in place in that year, and (after appropriately inflating their incomes) the rules that were in place in 1992. Among those made eligible for Medicaid between 1987 and 1992, only 3 percent of women and 10 percent of children were eligible because AFDC eligibility criteria expanded. The remainder were made eligible because of the expansion of coverage to nontraditional groups.

This overall trend masks substantial heterogeneity across the states in eligibility changes, as documented in Currie and Gruber [1994, 1995]. In the South AFDC eligibility has traditionally been very low, and thus eligibility increases were most pronounced. In the Northeast and Far West the eligibility increases were much smaller.

The potential for these expansions to crowd out private cov-

11. These figures come from calculating eligibility under AFDC rules only, and then under rules for AFDC and the other Medicaid options.
average is evidenced in Table II. As the first block of the table shows, about three-quarters of the nonelderly population have private insurance coverage, roughly 15 percent have public coverage, and about 15 percent are uninsured. As the second block shows, those traditionally eligible for Medicaid (based on 1987 rules) are much less likely to have private insurance (only 30 percent) and much more likely to have public coverage (about 50 percent) or be uninsured (about 25 percent). Strikingly, however, those made eligible for Medicaid between 1987 and 1992 are much more likely to have private coverage than the traditional Medicaid recipients. Almost two-thirds of this group had private insurance coverage in 1987, and only a quarter were uninsured. Indeed, the rate of uninsurance for the newly eligible group is the same as for the group already covered by Medicaid. Thus, as Medicaid eligibility limits increased, the population newly acquiring eligibility was increasingly populated by those with pri-
In fact, the rise in Medicaid coverage over this period was roughly commensurate with a decline in private insurance coverage. Figures IIa and IIb show the trends in Medicaid coverage, private insurance coverage, and uninsurance for children and women of child-bearing age. In both cases the percent of the population with Medicaid coverage increased dramatically, particularly after 1989, and private coverage fell substantially at exactly the same time. As a result, the fall in the percent uninsured was not very large for children. For women of child-bearing age the uninsurance rate actually rose.

While these figures are suggestive of a crowd-out, they are not definitive. There was a sizable recession in the early 1990s. Presumably, recessions reduce private coverage and increase the percent of the population with Medicaid coverage. In addition, there were a number of important changes in the structure of private insurance markets. Real health insurance premiums rose by roughly 70 percent from 1988 to 1993, real deductibles in fee-for-service plans increased by a third, and the share of health insurance premiums paid for by employers fell by up to 10 per-
cent for individual policies [Gabel et al. 1989, 1993]. These changes might have some independent effect on private insurance coverage. In the next section we use cross-sectional variation in Medicaid eligibility to estimate more convincingly the extent of crowd-out.

III. FIRST ESTIMATES OF CROWD-OUT

We begin by using data on children and women of childbearing age to estimate equations of the effect of own Medicaid eligibility on own insurance coverage. Our specification is

\[ \text{COV}_i = \beta_0 \text{ELIG}_i + X_i \beta + \sum \alpha_{\text{state}_i} + \sum \alpha_{\text{time}_i} + \epsilon_i. \]

COV, is an indicator for insurance coverage, either Medicaid, private insurance, or uninsured. ELIG, is a variable measuring the eligibility of individual i for Medicaid. As noted above, we use an indicator variable for Medicaid eligibility in the absence of better data about desired insurance purchases by individuals. We also defer to the next section the question of whether individual or family eligibility is the most appropriate independent variable. When Medicaid is the dependent variable, \( \beta_1 \) measures the mar-
TABLE III
SUMMARY STATISTICS FOR INDIVIDUAL DATA

| Variable              | Sample  
|----------------------|---------
|                      | Children | Women 15-44 | Other adults |
| Male                 | 0.51     | —           | 0.75         |
| White                | 0.80     | 0.82        | 0.86         |
| Married              | —        | 0.58        | 0.63         |
| Number of persons    | 4.12     | 2.77        | 2.29         |
| Workers              |          |             |              |
| 0 Workers            | 0.10     | 0.08        | 0.10         |
| 1 Worker             | 0.35     | 0.46        | 0.47         |
| 2 Workers            | 0.44     | 0.40        | 0.37         |
| 3+ Workers           | 0.10     | 0.06        | 0.06         |
| Health ins. unit type|          |             |              |
| Male/female          | 0.73     | 0.58        | 0.63         |
| Male head            | 0.04     | —           | 0.29         |
| Female head          | 0.24     | 0.42        | 0.08         |
| N                    | 266,421  | 194,139     | 355,333      |

People aged 15–18 are included in the child sample if dependent and in the appropriate adult sample if not dependent. Data are weighted to national totals.

original takeup rate for public insurance. For other forms of insurance coverage, \( \beta \), measures the crowd-out effect of increased eligibility.

\( X \) is a set of demographic controls, including the race of the child or woman; the sex of the child; the marital status of the woman; the number of persons and the number of workers in the household; the type of family (male and female head, male head only, female head only); and state and year dummy variables. Table III reports means of the demographic variables.

Several points about equation (1) deserve discussion. The first is pooling by age. We estimate equation (1) for all children jointly, and for women of child-bearing age. Without additional controls, however, the estimates for children may be biased. Infants (aged 0), for example, are generally more likely to be eligible for Medicaid than are older children, due to more generous expansions for this age group. It may also be the case that infants are less likely to have private insurance coverage than are older children—perhaps because one parent is at home with the children. Without controlling for age, our regressions would explain lower private insurance coverage for younger children by in-
creased Medicaid eligibility, when the relation is not causal. To control for this, we include age dummy variables for children, and age group dummy variables for women.\textsuperscript{12}

The second concern is the eligibility measure. Ideally, we would use the actual eligibility of each individual as the independent variable. Using actual eligibility is not appropriate, however, for three reasons. First, Medicaid eligibility is a complicated function of individual and family characteristics that may also be correlated with the demand for insurance, such as family structure. While we control for a number of the determinants of Medicaid eligibility, ELIG depends on these controls in a nonlinear, interactive manner that is difficult to fully capture in the regression. Second, ELIG may be endogenous to private insurance coverage, to the extent that the costs of private insurance coverage are reflected in lower wages and therefore lower income (and higher imputed eligibility). Finally, eligibility is likely to be measured with some error, for example, because actual eligibility is determined on a monthly basis and we have only annual income measures.

A solution to all three of these problems is to find an instrument that is correlated with individual eligibility for Medicaid, but not otherwise correlated with the demand for insurance. A natural strategy that meets these conditions is to create an instrument that varies only with the legislative environment in the state and year in which the individual lives. This will eliminate the bias from omitted individual factors and endogeneity, since the measure is not individual specific, and from error in accurately determining eligibility, to the extent that this error arises from imperfect information on individual characteristics.\textsuperscript{13} To create such an instrument, we first select a national random sample of 300 children of each age and 3000 women of child-bearing age in each year. We then assign that same sample to each state in that year and use our eligibility program to compute average state level eligibility measures for this sample of children and women of child-bearing age, which we denote SIMELIG. Since our sample is national, it is unaffected by state-level demographic differences or differences in local economic conditions. In essence,

\textsuperscript{12} Our children regression has dummy variables for ages 0 through 18. The regression for women has dummy variables for ages 15–18, 19–29, 30–39, and 40–44.

\textsuperscript{13} To the extent that measurement error arises from errors in our representation of the legislative regime, however, the error will remain.
our measure just weights the rules in each state by their effects if applied nationally. We instrument for individual eligibility with this simulated eligibility measure.14

The one potential problem with this measure is that it relies on the exogeneity of state legislation. If states expand Medicaid eligibility because of reductions in private coverage, it could lead to a spurious crowd-out estimate. This "endogenous legislation" scenario is unlikely to be very problematic in our context since much of the permanent variation in eligibility is coming from federal mandates on states of differing initial eligibility generosity, rather than state-specific expansions beyond the federal mandates. For example, 90 percent of the children made eligible between 1987 and 1992 qualified for Medicaid under Federally imposed minimum guidelines. Only 10 percent were eligible because of state coverage beyond the Federal mandate. For women of child-bearing age almost 70 percent of those newly eligible were eligible because of the Federal mandate.15 We have also experimented with including full sets of age-year and state-year interactions, to capture any omitted time trends in insurance coverage that vary by age group or location and are correlated with our eligibility measure. The results were very similar to those reported below for children, and for the estimates of Sections IV and V as well.16

Table IV reports instrumental variables estimates of equation (1). The first three columns report the effect of eligibility expansions on coverage of children. The latter three columns report the results for women of child-bearing age.17 Within each set of columns there are separate regressions for Medicaid coverage, private health insurance coverage, and uninsurance. All regres-

14. This approach follows Currie and Gruber [1994, 1995]. The first-stage coefficients from these regressions are 0.842 (0.016) for children, and 0.950 (0.028) for women. The fact that the coefficients are less than one reflects the measurement error from the size of the national sample.

15. States did have some discretion as to the timing of the eligibility expansions. But if we restrict our regressions to the first and last years of our sample, thus using only the permanent variation in eligibility, we get very similar results.

16. We cannot include state-year interactions in the regressions for the women of child-bearing age since our only variation in eligibility is by state and year.

17. Female 15–18 year-olds who are children are included in the child sample, and those who are household heads or spouses are included in the adult sample. Those who are children actually have two options for eligibility: as children, and (for the costs of pregnancy) as pregnant women. For the purposes of these regressions we consider only their eligibility as children. When we expand the measure to family dollars, we account for the pregnancy option as well.
sions are run as linear probability models for ease of computation with our sizable samples and for consistency of our instrumental variables procedure. We also weight the regressions to be representative of national totals. The standard errors are corrected for heteroskedasticity due to the dichotomous dependent variable.

In all cases, the demographic controls enter in the expected direction. Whites, households with a male head, households with more workers, and households with fewer people are more likely to have private coverage, and less likely to have Medicaid or be uninsured.

The first row shows the effect of Medicaid eligibility. There is a significant effect of eligibility on coverage for children, with an estimated take-up rate of 24 percent. This is substantially below the average take-up rates for other public programs estimated in the literature. These marginal and average estimates are not directly comparable, however, when the program is expanding to cover a population that generally already has private insurance coverage. Indeed, of the children who were made newly eligible between 1987 and 1992, only 27 percent were uninsured in 1987 (see Table II). Thus, if the increase in coverage were only among the uninsured population, the take-up rate would be almost 90 percent. Of course, given that some of the Medicaid increase reflects dropping of private coverage, this estimate is an upper bound on the true take-up rate.

For women there is no statistically significant increase in Medicaid coverage as eligibility expands. Roughly 11 percent of women are pregnant in any given year, so the estimated take-up rate is only about 7 percent. Thus, the expansions do not appear to be affecting coverage throughout the pregnancy.

18. We have reestimated some of our regressions using probit models in the second stage, and the results are quite similar.
19. All women who give birth in a year must have been pregnant at some time during that year. In addition, between two-thirds and three-quarters of women whose pregnancies begin in one year will give birth in the next year. With a birth rate of 6.5 percent, the percent pregnant in any year is at most \((1 + .75) \times 6.5\), or 11.4 percent. Since 24 percent of those made eligible by the 1987–1992 rule changes were uninsured, the take-up rate among the uninsured population is roughly 30 percent.
20. Note that this finding is consistent with the results of Currie and Gruber [1994], who find that Medicaid eligibility changes closely targeted to AFDC-like populations led to increased Medicaid coverage and improved birth outcomes, but that the type of broad expansions to moderate income groups on which we focus here had little impact on either coverage or outcomes. Note also that this weak result is not driven by our broad age range for women. The findings are quite similar if we restrict our analysis to 18–29 year-old women instead.
<table>
<thead>
<tr>
<th>Independent variable</th>
<th>Children</th>
<th>Women</th>
</tr>
</thead>
<tbody>
<tr>
<td>Medicaid</td>
<td>0.235</td>
<td>0.008</td>
</tr>
<tr>
<td>Private</td>
<td>-0.074</td>
<td>-0.119</td>
</tr>
<tr>
<td>Uninsured</td>
<td>-0.735</td>
<td>-0.018</td>
</tr>
<tr>
<td>Eligible for Medicaid</td>
<td>0.001</td>
<td>0.001</td>
</tr>
<tr>
<td>Medicaid</td>
<td>0.002</td>
<td>0.002</td>
</tr>
<tr>
<td>Private</td>
<td>-0.003</td>
<td>-0.002</td>
</tr>
<tr>
<td>Uninsured</td>
<td>-0.001</td>
<td>-0.001</td>
</tr>
<tr>
<td>Demographics</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Male</td>
<td>0.089</td>
<td>0.081</td>
</tr>
<tr>
<td>White</td>
<td>-0.069</td>
<td>-0.019</td>
</tr>
<tr>
<td>Married</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Health ins. unit</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Number of people</td>
<td>0.025</td>
<td>0.010</td>
</tr>
<tr>
<td>Male/female</td>
<td>-0.151</td>
<td>0.176</td>
</tr>
</tbody>
</table>

Note: Significant at the 0.01 level.
Male head  | -0.124 | 0.030 | 0.063 | — | — | —  
| —         | (0.005) | (0.006) | (0.006) | — | — | —  

0 Workers | 0.392  | -0.536 | 0.113 | 0.513 | -0.651 | 0.084  
|          | (0.009) | (0.011) | (0.010) | (0.012) | (0.017) | (0.016) |

1 Worker  | 0.044  | -0.156 | 0.109 | 0.064 | -0.159 | 0.076  
|          | (0.003) | (0.004) | (0.003) | (0.004) | (0.006) | (0.006) |

2 Workers | 0.004  | -0.051 | 0.041 | 0.045 | -0.041 | 0.005  
|          | (0.002) | (0.003) | (0.003) | (0.002) | (0.004) | (0.003) |

Summary statistics  

| R²     | 0.355  | 0.270  | 0.042  | 0.312  | 0.233  | 0.048  
|        | (0.005) | (0.006) | (0.006) | (0.006) | (0.006) | (0.006) |

| Dep. var mean | 0.174  | 0.719  | 0.131  | 0.097  | 0.736  | 0.157  
|               | (266,421) | (266,421) | (266,421) | (266,421) | (266,421) | (266,421) |

| N      | 194,139 | 194,139 | 194,139 | 194,139 | 194,139 | 194,139 |
|        | (266,421) | (266,421) | (266,421) | (266,421) | (266,421) | (266,421) |

All regressions are estimated by instrumental variables, where eligibility is instrumented by the simulated percent eligible in the state-year-age group. Regressions also include age dummy variables (by year for children and in groups for women), state, and year dummy variables. Data are weighted to national totals. Standard errors are corrected for heteroskedasticity.
The remaining columns estimate the effect of Medicaid eligibility on private insurance coverage and the rate of uninsurance. The increase in the Medicaid population plus the fall in private insurance need not sum to the effect on the uninsured because of coverage under other government programs and because individuals may have both Medicaid and private coverage in a given year.21

As Medicaid eligibility expands, there is a statistically significant offset in private insurance coverage for children and an insignificant reduction in coverage for women of child-bearing age. For children, each ten-percentage-point increase in Medicaid eligibility (roughly the range in Table I) results in a 0.74 percentage point decline in private insurance coverage. On net, the rate of uninsurance for children falls, but by only one-half as much as the increase in Medicaid coverage. For women, each ten-percentage-point increase in Medicaid eligibility (about half the range in Table I) results in a 0.45 percentage point decline in private insurance coverage, and a 0.46 percentage point increase in uninsurance. Recall that these findings are independent of general time trends in private insurance coverage and uninsurance, since the model includes a full set of year dummy variables.

What do these findings imply for the extent of crowd-out? One estimate of crowd-out is the reduction in private insurance coverage accompanying the increase in Medicaid coverage. For children this estimate is 0.074/0.235, or 31 percent. If the goal of the expansions is to reduce the rate of uninsurance, an alternative calculation is the percent of the increase in Medicaid that was not associated with a reduction in the uninsured population. For children this estimate is 1-0.119/0.235, or 49 percent.

For women both of these estimates suggest crowd-out of over 100 percent, although the estimated crowd-out effect is less precise than for children. In addition, the issue of crowd-out is more complicated for women, because the coverage that they receive

21. Two other government programs may cover these groups: CHAMPUS (for dependents of active duty military) and Medicare (for dependents of elderly or the disabled). As expected, for children Medicaid eligibility has only a small and insignificant effect on coverage through these sources, with a coefficient of −.004 (.005). More importantly, as Medicaid expands, the number of children with both Medicaid and private health insurance during the year increases. The coefficient on having dual coverage is .038 (.010). The net effect of Medicaid on insurance coverage is therefore the increase in either Medicaid or health insurance coverage (.235 − .074) less the increase in dual coverage (.038) and less the reduction in CHAMPUS and Medicare (−.004), for a net of .119. This is the reduction in the share uninsured, noted in the third column.
through the public sector is only conditional coverage: health insurance when they have one particular medical need (pregnancy) but not at all times. Thus, the right notion of coverage should add in to the direct Medicaid coverage increase the additional implicit coverage that all women have for pregnancy. We return to this point below.

IV. CHANGES IN EMPLOYER-BASED COVERAGE

Having documented that there is a reduction in private health insurance as Medicaid eligibility increases, we now consider the mechanisms through which this crowd-out occurs. Most (nearly 90 percent) private health insurance in the United States is provided through employment. Thus, to understand the reduction in private health insurance, we focus on employer-provided insurance. Following the discussion of Part I, there are at least three mechanisms through which employer-provided coverage could fall as Medicaid eligibility increases. First, employers may reduce the generosity of their insurance as Medicaid expands: through less service coverage, increased cost sharing, or dropping insurance entirely. Second, workers may choose to decline coverage for themselves and their family, even with no change in employer behavior. Finally, workers who are not eligible for Medicaid but who have eligible dependents may decide to maintain coverage for themselves but to decline coverage for their dependents. We investigate each of these responses.

IVA. Empirical Framework

To estimate these responses, we need to incorporate the effect of Medicaid expansions on the family as a whole. Fathers, for example, will be affected by Medicaid expansions for their children, even if they are not affected on their own. Similarly, the private insurance coverage of mothers might be affected not only by their own eligibility for pregnancy, but by the eligibility of their children as well. Returning to the analytic framework described in Section I, we need some measure of the family-specific value ($M$) of the insurance provided by Medicaid.

Ideally, our measure would incorporate not only each family's valuation of Medicaid, but also the generosity of the insurance option that they have at their workplace (i.e., the constrained point on their budget set). We have no data on desired insurance coverage or the benefits offered on the current policy, however.
We therefore proxy for the value of the public option by assuming that all families value insurance equally and have the option of equivalent private coverage. In this case, the relative value of Medicaid can be measured by the share of family medical spending that is eligible for Medicaid coverage, the Medicaid replacement rate.\footnote{This family measure may be too high or too low. Since Medicaid is likely to be less generous than private insurance, the share of dollars of those eligible is likely to be too high. On the other hand, to the extent that those who drop private insurance and are not covered by Medicaid can get care through provider "uncompensated care," then we are understating the benefits of dropping private insurance for the family and moving onto Medicaid.}

To implement this measure, we first sort individuals into "health insurance units" (HIUs): the group for whom health insurance is traditionally sold. A health insurance unit consists of a family head, spouse, and any children under 19 (or under 23 who are full-time students). Other relatives, while they may be in the same Census family, are typically not covered by the same health insurance policy.\footnote{The regressions in Table IV included demographic information about health insurance units, rather than families.} Our measure of Medicaid generosity is therefore

\[
\text{\%HIU Dollars} = \frac{\sum_k \text{SPEND}_k \times \text{ELIG}_k}{\sum_k \text{SPEND}_k \times \text{NUM}_k},
\]

where \(k\) indexes single-year age groups of children, and broader age groups for adults.\footnote{We divide adults into those age 19–29, 30–39, 40–49, 50–59, and 60–64. We further divide women into ages 40–44 and 45–49 because pregnancy is assumed to occur only in the first group.} \(\text{SPEND}_k\) is the expected health spending in a year for that age group based on data from the 1987 National Medical Expenditure Survey (NMES), the most recent national survey of individual medical spending. The Appendix shows average spending by age and sex.

\(\text{ELIG}_k\) is the eligibility for Medicaid of family members in age group \(k\), and \(\text{NUM}_k\) is the number of family members in that age group.\footnote{Note that all members of a given age group within a family will have the same eligibility status.} By definition, eligibility is zero for men 19 and older and women 45 and older. For children below age 15 \(\text{ELIG}\) is simply a 0/1 variable. For women of child-bearing age, the eligibility measure is more complex, since in some circumstances (i.e., AFDC eligibility) all spending will be covered while in other circumstances only pregnancy-related spending will be covered. To address this, we divide the eligibility measure into two parts: full-
year eligibility (as either a child or an adult), and incremental eligibility for pregnancy-related coverage only. We weight the probability of full-year eligibility by total annual spending, and the probability of coverage for pregnancy only by the expected cost of pregnancy.\footnote{The expected cost of pregnancy is the product of the age-specific fertility rate and total annual spending for women who had a child during the survey year. This measure may overstate or understate the true cost of pregnancy. Because it includes spending on all services, it is likely to overstate pregnancy costs. Because the survey is only for one year, however, it will understate costs for women who incurred some pregnancy-related costs in previous years. Our resulting estimate of pregnancy costs is $3996 in 1987, which is close to other estimates in the literature.}

Trends in family eligibility for Medicaid are summarized in the final column of Table I. The percent of family dollars covered by Medicaid rises by over 50 percent over our sample period, from 8 percent in 1987 to over 12 percent in 1992.

Since we are estimating models for worker coverage, we augment our regression equation (1) in several ways. First, we add controls for education, one-digit industry and occupation, firm size, full-time/part-time and full-year/part-year work, and the state/year unemployment rate. A first pass estimating equation is therefore

\[
(1') \text{COV}_i = \beta_1 \%\text{HIU Dollars}_i + X_i \beta + \sum_\text{state} \alpha_i + \sum_\text{time} \alpha_i + \epsilon_i.
\]

There are two factors to consider in estimating equation (1'). The first issue is controlling for the demographic structure of the HIU. As noted above, we do not want the impact of Medicaid to be identified simply by the age structure of the children in the HIU, since families with older or younger children may have different levels of private insurance coverage for reasons independent of Medicaid eligibility. Indeed, the problem of controlling for the effect of family structure on insurance demand is more complex than just the age distinction suggests. Consider two families with the same income, but where the adults are of different ages. Since spending increases with age, families with older adults will have a smaller fraction of spending covered by Medicaid than families with younger adults, holding constant the eligibility of the children. Families with older adults may be more likely to have private insurance coverage for other reasons, however. Without controlling for these age differences, the estimates of crowd-out will be biased.
To control for these factors, we note that in a given family, a child or woman in age group $k$ accounts for $\frac{\text{SPEND}_k}{(\sum \text{SPEND}_k \times \text{NUM}_k)}$ share of expected family spending. Therefore, the importance of each age group in total family spending—with constant Medicaid policy—can be represented as

$$\text{DEMOG}_k = \frac{\text{SPEND}_k \times \text{NUM}_k}{(\sum \text{SPEND}_k \times \text{NUM}_k)}.$$  

There are 22 DEMOG terms: one for each age child, and three for women of child-bearing age. Once these family structure variables are included in the regression, the variation remaining in our Medicaid eligibility measure is due entirely to differing eligibility rates by state, year, and age, not to the demographic composition of the family. Our final estimating equation is therefore

$$(1') \quad \text{COV}_i = \beta, \%\text{HIU Dollars}_i + X_i \beta + \text{DEMOG}_{i,k} \gamma_k + \sum \alpha_{\text{state}_i} + \sum \alpha_{\text{time}_i} + \epsilon_i.$$  

In estimating equation (1'), we are once again concerned about omitted variables bias, endogeneity, and measurement error arising from using actual family Medicaid eligibility. Thus, we create an instrument for HIU dollars. To form this instrument, we use the simulated eligibility measures noted above. Since simulated eligibility reflects only the state's legislative environment, a measure of eligibility for each family reflecting only the state's Medicaid rules is given by

$$(2') \quad \%\text{SIM Dollars}_i = \frac{\sum \text{SPEND}_k \times \text{SIMELIG}_k \times \text{NUM}_k}{\sum \text{SPEND}_k \times \text{NUM}_k}.$$  

Our first-stage regression explains eligibility for each family with these simulated eligibility differences.

We restrict the sample to workers aged 24 to 64. We exclude young workers because the distinction between working and school enrollment is critical in determining an individual's health insurance unit, and we do not want to be biased by movements in estimating equation (1').

27. That is, we can rewrite (2) as $\%\text{HIU Dollars}_i = (\sum \text{DEMOG}_k \times \text{ELIG}_i).$ Thus, when we control for the full set of DEMOG controls, we are only identifying the model from differences in eligibility regimes, interacted with individual family structure, and not from family structure itself.

28. Focusing only on workers potentially creates a bias if Medicaid policy affects labor supply decisions. The direction of this bias is not obvious: for single women the Medicaid expansions induced labor supply by allowing them to leave welfare and retain insurance for themselves and their children [Yelowitz 1995]; but if some married women are working solely to provide insurance for their family, then the availability of public insurance may induce them to stop working. In any case, we find strong insurance effects on males as well as females, suggesting that our findings are not driven by labor supply responses.
into and out of schooling. We exclude older workers because they are automatically entitled to public insurance under the Medicare program.

**IV.B. The Distinction between Offering and Take-up**

We begin with the effects of Medicaid eligibility on the decision of employers to offer health insurance and the decision of employees to take up this coverage. While we would like to examine the effect of Medicaid expansions on the generosity of benefits offered, no individual-level survey spanning this period includes data on the type of benefits offered. We thus restrict ourselves to the question of whether coverage is offered and taken up.

The March CPS data we have been using do not contain information on whether insurance was offered: only on whether the worker received coverage from the employer. In order to separate these two effects, we turn to data from the May 1988 and April 1993 Employee Benefits Supplements to the CPS. These supplements collected information about a variety of workplace benefits for a subsample of workers, including whether their employer offers insurance and whether the employee chose to be covered. Because the supplements are small and cover only two years, there are fewer than one-tenth as many workers in these data as are in the full set of March surveys.

In analyzing the Benefits data, we are hampered by the fact that no survey of workers accurately captures information about the firm as a whole. Presumably, firms should decide to stop offering insurance if a significant share of total employees value Medicaid highly. Unfortunately, we know only about the worker who is surveyed, not about his or her coworkers. As an approximation, we simply use the worker-level information when we estimate the probability that the firm offers insurance. To the extent that the workers we observe are not representative of workers at the firm as a whole, our estimates of the propensity to offer insurance may be misstated.

The probability that a worker is offered insurance has fallen over time, from 85 percent in 1988 to 77 percent in 1993. The take-up rate has also fallen, but not by as much (from 85 percent in 1988 to 83 percent in 1993). General time series trends in offering and coverage, however, will be captured in the year dummies included in the regression.

Table V presents instrumental variables estimates of the effect of Medicaid eligibility on offering and take-up of health insurance. The first three columns report the results for the 1988 and
The table presents the employer and employee responses to increases in Medicaid eligibility. The independent variable is the percentage of HIU dollars, and the dependent variables include coverage from employer, employer offers, and take-up if offered. The table also includes summary statistics such as R-squared, dependent variable mean, and sample size.

The sample includes workers aged 24-64. All regressions are estimated using instrumental variables, where the percent of HIU dollars eligible is instrumented by the simulated percent of dollars eligible, as described in the text. The regressions include controls for the demographic characteristics of the worker (sex, race, marital status, age, education) and the HIU (number of persons, male/female head, male head only, number of workers); characteristics of the job (full-time/part-time dummy, one-digit industry and occupation dummies, firm size dummies); the state/year unemployment rate; state and year; and the number of children of each age and women of child-bearing age times their share of HIU costs. The March data subdivide the full-time/part-time variable into full-year/part-year status as well. Data are weighted to national totals. Standard errors are corrected for heteroskedasticity.
1993 Supplement Data. In the first column we report the result for coverage through an employment-based policy. The next two columns then separate coverage into whether health insurance is offered and whether coverage is chosen. For simplicity, we do not report the coefficients on the demographic or work variables. The coefficients on these variables are generally in accord with intuition: males, whites, and full-time workers are more likely to be offered coverage and to accept this coverage.

Increasing the share of dollars made eligible for Medicaid reduces the probability that a worker has health insurance coverage. Each percentage point increase in the share of HIU costs made eligible for Medicaid (roughly one-quarter of the increase over this period) lowers the probability of employer coverage by 0.18 percentage points. This result is large and significant at the 10 percent level. Since the sample size in the Supplement data is small and it covers only two years, we use the annual March data to confirm this result. These results are in the fourth column of the table. The coefficient on Medicaid eligibility is somewhat smaller in the March data, but of the same order of magnitude. The standard error of the estimate with the March data is, of course, much smaller than that with the Supplement data.

The next two columns divide the coverage result into the probability that employers offer insurance, and the probability that workers take up this insurance. There is no effect of Medicaid eligibility on the decision to offer insurance. While the standard error is once again large, the point estimate is approximately zero. At the same time there is a large and statistically significant effect of Medicaid eligibility on the take-up rate, conditional on being offered insurance. Thus, the estimates suggest that all of the reduction in worker coverage is coming through lower take-up rates, not reduced employer offering of insurance.

If employees are choosing to decline coverage more as Medicaid eligibility increases, it may be because employers are encouraging them to do so by reducing the generosity of workplace insurance or by increasing the share of health insurance premiums paid by workers. In 1991 the typical employer paid about 86 percent of a single premium, but only about 60 percent of the additional costs for dependents, and the share of premiums paid

29. The average employer paid about 70 percent of a family premium. A typical family policy costs 2.5 times a single policy, yielding the 60 percent marginal payment estimate.
for by employers was falling over this period. While neither CPS survey reports the share of the premium paid for by employers, the March survey does ask whether employers pay for all, part, or none of the cost of the insurance policy. Among workers with insurance, 36 percent report that their employer pays all of the premium, 59 percent report that the employer pays some of the premium, and 5 percent report that the employer pays none of the premium.

In an earlier version of this paper [Cutler and Gruber 1995], we considered the effect of Medicaid eligibility on the probability that an employer pays for all, part, or none of the cost of health insurance. We found a sizable and significant drop in the probability that employers pay all of the cost of insurance. For every percentage point increase in the share of family expenditures made eligible for Medicaid, there is a 0.07 percentage point reduction in the likelihood that employers will pay all of the costs of insurance. Unfortunately, the fact that the number of workers with coverage is changing makes it impossible for us to definitively decompose this response into changes in employer payment practices and changes in employee take-up behavior. However, since it is unlikely that workers in firms where employers were paying all of the cost of insurance were the ones to drop their coverage, our findings do suggest that employers were reducing the generosity of their insurance contributions in the face of increased Medicaid eligibility for their workers. This mechanism may explain part of the reduction in take-up rates associated with eligibility increases. Longitudinal data on workers in the same firm over time will be required to answer this question definitely, however.

IV.C. Shifts in Who Is Covered

Since most of the people made eligible by the expansions are dependents, a natural reaction for workers may be to drop coverage of eligible dependents only, while maintaining insurance for themselves. If this happens, then along with the overall fall in coverage there should be a net shift from family to individual coverage among workers. We can examine whether there is such a response using data from the March CPS on who is covered by

---

30. In the longer run one might expect health insurance policies to develop that “carve out” people eligible for Medicaid from the rest of the family. It would be interesting to examine whether this was happening, but it is not an issue we can readily examine with our data.
TABLE VI
THE EFFECT OF MEDICAID ELIGIBILITY ON OWN AND DEPENDENT COVERAGE FOR
MEN WITH DEPENDENTS POTENTIALLY ELIGIBLE FOR MEDICAID

<table>
<thead>
<tr>
<th>Independent variable</th>
<th>Any employer coverage</th>
<th>Who covered in family</th>
</tr>
</thead>
<tbody>
<tr>
<td>% of HIU dollars</td>
<td>-0.172</td>
<td>0.066</td>
</tr>
<tr>
<td></td>
<td>(0.065)</td>
<td>(0.043)</td>
</tr>
<tr>
<td></td>
<td>(0.069)</td>
<td></td>
</tr>
<tr>
<td>Summary statistics</td>
<td></td>
<td></td>
</tr>
<tr>
<td>R²</td>
<td>0.299</td>
<td>0.074</td>
</tr>
<tr>
<td>Dependent variable mean</td>
<td>0.689</td>
<td>0.081</td>
</tr>
<tr>
<td>N</td>
<td>119,000</td>
<td>119,000</td>
</tr>
</tbody>
</table>

The sample is male workers aged 24-64, who have a woman of child-bearing age or child in their HIU. All regressions are estimated as instrumental variables, where the percent of HIU dollars eligible is instrumented by the simulated percent of dollars eligible, as described in the text. All regressions have the same set of controls as in Table V. Data are weighted to national totals. Standard errors are corrected for heteroskedasticity.

We divide policies into individual coverage only, and those that also cover some other dependents. In each case, we count workers without employer-provided insurance as not covering any dependents. This once again avoids any bias from sampling only workers with coverage.

We focus on married male workers aged 24–64, with either a 15–44 year-old spouse or a child or both. The male workers are not themselves eligible, but they have dependents who are. Therefore, if the decision is to drop dependent coverage and keep individual coverage, it should be concentrated in this population. Among this group only 8 percent of those with coverage from their employer have individual coverage.

Table VI reports regressions explaining who is covered by the worker’s policy. All of the independent variables from Table V are included in the regression, but are not reported. The first column replicates the regression from column 4 of Table V for this subpopulation. There is in fact a somewhat stronger effect of Medicaid policy on this group: a one-percentage-point rise in the fraction of family medical costs eligible for Medicaid leads to a 0.17-percentage-point reduction in the probability of being cov-

31. The survey actually asks more detailed questions about whether the policy covers one’s spouse, children in the household, children outside of the household, and others. We focus on dependent coverage in general, since most policies do not offer workers the ability to differentiate between dependents. Indeed, when we estimate the regressions below separately for spousal and child coverage, the results for both groups are very similar to the results for any dependent.
ered by insurance. This stronger effect may arise from the fact that the firms in which these workers (who have a high share of HIU dollars eligible) are employed are the ones who are moving aggressively to get their employees to drop employer-provided coverage.

As the next two columns show, there is a sizable increase in individual coverage for workers, which is matched by a drop in dependent coverage that is larger than the overall coverage reduction in column 1. Given that we do not know the type of coverage specifically for those workers who drop coverage, we cannot infer for certain the net shift from family to individual coverage. However, since individual coverage in this population is relatively small, and since a male with individual coverage would have no incentive to drop coverage in response to Medicaid expansions, most of the overall decrease in coverage is likely to come from those with family policies. In that case, the estimates imply that, in addition to the 0.17 percent of workers who drop their employer-provided insurance when there is a one-percentage point rise in HIU costs covered, another 0.07 percent of workers are dropping their dependents and shifting to individual policies.

Thus, increases in the eligibility of family medical spending for Medicaid coverage significantly lower the probability of worker insurance coverage. This reduction in coverage appears to arise solely from worker take-up decisions, although this may be a response to increased employer cost sharing. In addition, workers are reacting by both dropping insurance coverage entirely and by eliminating coverage of their dependents only.

V. TOTAL CROWD-OUT ESTIMATES

Having examined the sources of crowd-out, we now return to estimate its total magnitude. As the previous section showed, the most appropriate specification of Medicaid effects is in terms of family eligibility. We thus reestimate equation (1) as

\[ COV_i = \beta_1 \%OWN\ Dollar_i + \beta_2 \%OTHER\ Dollar_i + X_i \beta + DEMOG_i \gamma_k + \sum \alpha_i state_i + \sum \alpha_i time_i + e_i. \]

COV, is again an indicator for Medicaid or private insurance coverage. %OWN Dollar and %OTHER Dollar split %HIU Dollars into the share attributable to the individual child or woman, and
the share attributable to all other family members. We allow insurance coverage to depend differently on these two factors. These two variables are somewhat collinear, making distinct estimation of the effects difficult, but we get almost precisely the same results if we combine the two measures into one total measure of family dollars. $X$ is the set of demographic controls noted in Table V.\textsuperscript{32} As in Table V we include the full set of DEMOG controls. Once again, the regressions are estimated using instrumental variables, where the instruments are calculated from the simulated eligibility measures, split into the $\%$OWN and $\%$OTHER components.

Table VII presents estimates of (1''). Results are presented for children (the first set of columns), women of child-bearing age (the second set of columns), and other adults (the third set of columns). For simplicity, we report only the coefficients on the share of HIU dollars accounted for by oneself and by others.

For children the estimates imply that both own and other eligibility significantly affect insurance coverage. A one-percentage-point increase in family dollars attributable to own coverage raises Medicaid coverage by 0.3 percentage points and lowers private insurance by nearly 0.2 percentage points. Eligibility of other family members has an even stronger effect on Medicaid coverage than the share of dollars attributable to oneself, but a weaker effect on private coverage.

For women of child-bearing age there is a very small effect of the woman's own potential coverage on receipt of Medicaid, which is consistent with Table IV. Medicaid coverage does respond to eligibility of other family members, however. Once again, own dollars have a larger effect on private coverage than do family dollars. Finally, for other adults increased family Medicaid eligibility leads to reductions in private insurance coverage.

For both children and women of child-bearing age, Medicaid coverage depends strongly on eligibility of other members of the family. There are two potential explanations for this finding. First, there may be substantial fixed costs of enrolling in Medicaid, such as stigma, so that only when there is significant coverage of total family dollars does the family enroll. Similarly, individuals may not sign up for Medicaid until someone in the

\textsuperscript{32} If the individual is not a worker, we impute job information from the primary worker. In the case of multiple workers, we use job controls for the head of the family. There is a dummy for having no workers in the HIU. In that case, all job controls are set to zero as well.
### TABLE VII
**REGRESSIONS EXPLAINING NET CHANGE IN MEDICAID AND PRIVATE INSURANCE COVERAGE**

<table>
<thead>
<tr>
<th>Independent variable</th>
<th>Children</th>
<th></th>
<th>Women 15–44</th>
<th></th>
<th>Other adults</th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Medicaid</td>
<td>Private insurance</td>
<td>Medicaid</td>
<td>Private insurance</td>
<td>Medicaid</td>
<td>Private insurance</td>
</tr>
<tr>
<td>% Own dollars</td>
<td>0.282</td>
<td>0.064</td>
<td>0.064</td>
<td>0.320</td>
<td>—</td>
<td>—</td>
</tr>
<tr>
<td></td>
<td>(0.028)</td>
<td>(0.033)</td>
<td>(0.033)</td>
<td>(0.053)</td>
<td>—</td>
<td>—</td>
</tr>
<tr>
<td>% Other dollars</td>
<td>0.481</td>
<td>0.524</td>
<td>—</td>
<td>—</td>
<td>0.171</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(0.027)</td>
<td>(0.040)</td>
<td>—</td>
<td>—</td>
<td>(0.037)</td>
<td></td>
</tr>
</tbody>
</table>

**Summary statistics**

- \( R^2 \)
- Dependent variable mean
- \( N \)

<table>
<thead>
<tr>
<th></th>
<th>Children</th>
<th>Women 15–44</th>
<th>Other adults</th>
</tr>
</thead>
<tbody>
<tr>
<td>( R^2 )</td>
<td>0.419</td>
<td>0.382</td>
<td>—</td>
</tr>
<tr>
<td>Dependent variable mean</td>
<td>0.174</td>
<td>0.719</td>
<td>0.097</td>
</tr>
<tr>
<td>( N )</td>
<td>266,421</td>
<td>266,421</td>
<td>355,333</td>
</tr>
</tbody>
</table>

All regressions are estimated as instrumental variables, where the percent of own dollars and other dollars eligible for coverage are instrumented by the simulated percent of dollars eligible, as described in the text. The table includes the same set of controls as in Table V. Information about work conditions is imputed for nonworkers, based on the primary worker in the family. Where there are no workers, the job controls are set to zero (and there is an indicator for no workers included in the regression). Data are weighted to national totals. Standard errors are corrected for heteroskedasticity.
family is sick. When more people are eligible, the likelihood that one eligible person is sick increases. Alternatively, this finding may be due to the reporting bias in the CPS discussed above. When more children are covered by Medicaid, mothers may report themselves covered by Medicaid as well. Or, parents may report all children as covered by Medicaid even if only some children are. In this case, the Medicaid increase calculated using the family dollars figure could be an overstatement of the actual Medicaid increase.\textsuperscript{33}

We can convert these results into population figures to estimate the total amount of crowd-out due to eligibility increases. Analytically, the change in insurance is the weighted sum of the changes in OWN and OTHER dollar eligibility, with the weights being the coefficients in the coverage regressions:

\begin{equation}
\Delta \text{Coverage} = (\Delta \% \text{OWN Dollar}_{87-92} \times \beta_{\text{OWN}} \\
+ \Delta \% \text{OTHER Dollar}_{87-92} \times \beta_{\text{OTHER}}) \times \text{POP}_{87}.
\end{equation}

We estimate the change in \%OWN Dollar coverage and \%OTHER Dollar coverage from the simulated expansion data in Table II (that is, by applying the 1992 eligibility rules to the 1987 population).\textsuperscript{34} Among children, for example, the increase in \%OWN dollar eligibility was 1.7 percent of expected family medical spending, and the increase in \%OTHER dollar eligibility was 3.7 percent of expected spending. The increase in Medicaid coverage for children was therefore 2.3 percent of the total number of children (66.6 million), or 1.5 million people.

The table below shows the results of these calculations for all demographic groups. As the first row reports, the Medicaid expansions increased coverage of children by 1.5 million persons. The decline in private coverage is about 0.6 million, however. The estimated crowd-out for children, therefore, is 40 percent of the increase in coverage. This is very similar to the estimate in Table

\textsuperscript{33} In addition, this same type of effect may lead us to understate the reduction in private coverage: in families where dependents have been dropped due to Medicaid eligibility increases, the family may still report that all members are covered. This would lead to a further understatement of the total crowding-out effect of Medicaid expansions.

\textsuperscript{34} We use the simulated increases in coverage instead of the actual increases in coverage because the actual increases include variation in economic conditions as well as legislative changes. Medicaid increases resulting from a weak economy may have a different effect on crowd-out than increases resulting from legislative change. In practice, however, our crowd-out estimates are nearly identical using the actual increases in eligibility.
IV. For women of child-bearing age there is a Medicaid increase of 0.7 million, but a reduction in private coverage of essentially the same amount. The crowd-out for this group is 100 percent. The estimate of crowd-out for women of child-bearing age is more dependent on the treatment of other family members than is the estimate of crowd-out for children. Finally, coverage of other adults falls by 0.3 million as eligibility for their children or spouse expands. Thus, our first estimate of the total crowd-out from the Medicaid expansion is 77 percent.

<table>
<thead>
<tr>
<th>Coverage of:</th>
<th>Change in coverage (millions)</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Medicaid</td>
</tr>
<tr>
<td>Children</td>
<td>1.5</td>
</tr>
<tr>
<td>Women 15–44</td>
<td>0.7</td>
</tr>
<tr>
<td>Other adults</td>
<td>-</td>
</tr>
<tr>
<td>Total</td>
<td>2.2</td>
</tr>
<tr>
<td></td>
<td>(-77%)</td>
</tr>
<tr>
<td>Conditional coverage of women</td>
<td>0.9</td>
</tr>
<tr>
<td>Conditional coverage of children</td>
<td>0.4</td>
</tr>
<tr>
<td>Total</td>
<td>3.5</td>
</tr>
<tr>
<td></td>
<td>(-49%)</td>
</tr>
</tbody>
</table>

Based on estimates in Table VII.

As noted in Section III, however, this coverage increase is an underestimate of the true increase. The Medicaid expansions explicitly did not give continuous coverage to women. Rather, they created a form of "conditional coverage": women are covered, but only for some expenses. As a result, women who are eligible for Medicaid in the event of pregnancy but who report themselves to be uninsured actually have some partial (conditional) insurance coverage.

Accounting for this conditional coverage is difficult because it requires estimating both the value of pregnancy coverage only and the likelihood that uninsured women would avail themselves of this coverage if they became pregnant. Evidence from case studies of the Medicaid expansions suggests that very few women who were made eligible took advantage of this eligibility to fund prenatal care services. Rather, due perhaps to poor information about eligibility, most women were enrolled in Medicaid by hospi-
tals at the time of their delivery. In fact, many hospitals have set up special departments, or contracted with outside parties, explicitly to sign up Medicaid eligibles for the program.\footnote{See U. S. General Accounting Office [1994], Currie and Gruber [1994], and Piper et al. [1990] for further discussion. These hospital efforts are profitable because uninsured individuals pay only a small share of their bills, so that if the hospital can sign them up for Medicaid it will increase hospital reimbursement.}

As a rough proxy for the value of conditional coverage, therefore, we use the dollars of expected hospital expenditures for pregnancy that an uninsured woman would receive. This will overstate conditional coverage to the extent that some uninsured women do not get signed up for Medicaid at the point of delivery, but it will understate conditional coverage if these women are receiving some prenatal care when they sign up for Medicaid.

In 1987 the hospital share of pregnancy costs was approximately $2680. Since approximately 11 percent of 15–44 year old women are pregnant in any given year, this is about $295 of expected coverage annually, or 25 percent of annual health spending for women of child-bearing age. Between 1987 and 1992 an additional 3.7 million uninsured women were made eligible for Medicaid coverage for pregnancy-related services. Therefore, the amount of conditional coverage is the equivalent of 0.9 million people (3.7 million × 25 percent).

In the same vein, Medicaid also provides a form of conditional coverage for uninsured children. The fact that these children are not continuously covered by Medicaid suggests that they are not availing themselves of the insurance for the purpose of their primary medical care. Once again, however, when these children appear in the hospital, they may be signed up for Medicaid, so that they have conditional coverage for their hospital spending. On average, 44 percent of spending for children is for hospitals. Between 1987 and 1992 the number of uninsured children eligible for Medicaid coverage rose by 0.9 million, so that the conditional coverage for this group amounts to 0.4 million children (0.9 million × 44 percent).

Adding the resulting 1.3 million (statistical) persons who are conditionally covered by Medicaid to the 2.2 million direct coverage increase leads to an effective coverage increase of 3.5 million people. With this estimate, the reduction in private coverage as a percentage of the rise in effective Medicaid coverage is 49 per-
Our final estimate of crowd-out is therefore 49 percent of the coverage increase.

While our estimates suggest substantial crowd-out, they do not suggest that most of the reduction in private insurance coverage over this period is because of Medicaid crowd-out. Relative to the percent of the population with private health insurance in 1987, there was a 9.9 million person decline in private coverage by 1992. Of this reduction about 17 percent was due to crowding out. The remainder is due to macroeconomic factors, changes in the demographic mix of the population, or changes in employer or worker behavior unrelated to Medicaid generosity.

VI. CONCLUSIONS

The past decade has witnessed substantial increases in public sector eligibility for health insurance. The percent of children who can receive Medicaid rose by 50 percent between 1987 and 1992, and the percent of women eligible for Medicaid if they are pregnant more than doubled. These expansions substantially increased Medicaid coverage—by about 2.2 million people. The expansions did come at a price, however. As Medicaid eligibility increased, there was a direct offset in private insurance coverage. This reduction in private coverage largely came from workers deciding to drop their own insurance coverage, and particularly coverage of dependents, rather than through firm decisions to stop offering insurance. For children our results consistently suggest crowding out of 30 to 40 percent. For adults the estimate of crowd-out is more sensitive to these interfamily eligibility issues. Accounting for these effects and allowing for conditional eligibility for Medicaid, we estimate a net crowd-out from the Medicaid expansions of one-half of the coverage increase.

It is important to note, however, that we have considered in this paper only the costs of reducing the uninsurance rate, not the health effects of the expansions. The Medicaid expansions led to more generous insurance coverage for the previously uninsured,36 which may have had benefits in terms of improved health outcomes.

36. They also lead to more generous insurance for those who receive a more generous policy under Medicaid than they did with private insurance (those whose private insurance choice lay below point M in Figure 1). We count such people as “crowded out” of private insurance, but they in fact wind up with more coverage.
Drawing the link between insurance and health is complicated, for two reasons. First, there is an alternative to private or public insurance coverage: receiving uncompensated care. People who are not insured typically receive emergency care, which is financed by shifting costs to people who are insured. As the quality of uncompensated care rises, the additional care received by those with insurance coverage declines. Generally, studies conclude that the insured receive better care than the uninsured. For example, Hadley et al. [1991] find that the insured are 30 to 300 percent more likely to receive expensive medical procedures than are the uninsured, and Wenneker et al. [1990] find that insured patients are 30 to 80 percent more likely to receive intensive surgery after a heart attack than are the uninsured.

How this additional care translates into health outcomes, however, is a subject of much debate. Hadley et al. [1991] do find that the uninsured are more likely to die in the hospital than the insured, but it is difficult to assess whether this is because they received less intensive treatment or because they were sicker when they were admitted. Currie and Gruber [1995] find that Medicaid eligibility increases for children were associated with increases in medical care utilization and health improvements. On the other hand, Piper et al. [1990] find no health benefits from Medicaid expansions in Tennessee, and Newhouse et al. [1993] conclude that there are no sizable health differences resulting from more or less generous insurance. Clearly, evaluation of the health effects of private insurance, Medicaid, and no insurance, is of primary importance for assessing the net benefits of public insurance expansions.

Beyond the effects of Medicaid expansions on the care received by Medicaid recipients is the issue of transfers to hospitals that care for the poor. Even if Medicaid expansions had no effects on the care received by those on the program, they had large effects on the income available to hospitals that see Medicaid patients, since these hospitals would otherwise absorb the costs of uncompensated care. To the extent that hospitals use this additional income to increase the level of care to everyone, Medicaid expansions may have large effects on the availability of care to the entire poor population. Indeed, Gruber [1994] finds that in-

37. Hospitals are not allowed to turn away a woman who is in labor, and obstetric costs are the single highest component of hospital uncompensated care [Saywell et al. 1989].
come pressure on hospitals in California led to reductions in overall delivery of uncompensated care during the 1980s.

More generally, our results may provide some guidance in the design of public insurance policy. The Medicaid expansions had an exclusionary structure, whereby individuals were either entirely eligible for public insurance or not eligible at all. An alternative would be to subsidize the private insurance purchase of all low-income people, with a sliding scale that offers high subsidies for very low-income persons but that declines as income rises. This approach would provide subsidies to everyone in a given income range, but would induce less distortion across types of insurance.\textsuperscript{38}

Alternatively, a waiting period could be imposed between when an individual loses private coverage and when they become eligible for Medicaid, or individuals could be excluded from Medicaid coverage entirely if they are offered but decline private coverage [Nichols and Zeckhauser 1982]. Each of these might limit the ability to move easily from private to public insurance. Exploring these types of policy designs is a clear priority for future research.

Finally, our results raise the fundamental question of how the public sector should design programs to improve the health of the poor. Expanding public insurance is one common suggestion, but this may be quite expensive, due partially to private crowd-out. Another alternative to subsidized insurance is direct provision of care to the poor, in the form of better public hospitals and outpatient care in poor areas. Direct provision of care may also cause private insurance crowd-out, however, raising the costs of this policy direction as well. The extent of crowd-out from public provision (as opposed to public insurance) is an unanswered question. And some have suggested putting aside coverage expansions and focusing on adverse social behavior, such as smoking, drinking, firearms, and bullets [Cutler 1995]. Evaluating the costs and benefits of policies from this broad menu of public options is an important step for future research.

\textsuperscript{38} At the same time, such a sliding scale will affect labor supply incentives for those near recipiency. The net effect on work is unclear.
### APPENDIX: EXPECTED HEALTH CARE COSTS BY AGE AND SEX

<table>
<thead>
<tr>
<th>Age</th>
<th>Male</th>
<th>Female</th>
</tr>
</thead>
<tbody>
<tr>
<td>0</td>
<td>$2486</td>
<td>$2486</td>
</tr>
<tr>
<td>1</td>
<td>1266</td>
<td>1266</td>
</tr>
<tr>
<td>2-5</td>
<td>450</td>
<td>450</td>
</tr>
<tr>
<td>6-9</td>
<td>399</td>
<td>399</td>
</tr>
<tr>
<td>10-14</td>
<td>537</td>
<td>537</td>
</tr>
<tr>
<td>15-18</td>
<td>930</td>
<td>930</td>
</tr>
<tr>
<td>19-29</td>
<td>625</td>
<td>1228</td>
</tr>
<tr>
<td>30-39</td>
<td>724</td>
<td>1350</td>
</tr>
<tr>
<td>40-49</td>
<td>1326</td>
<td>1518</td>
</tr>
<tr>
<td>50-59</td>
<td>1979</td>
<td>2115</td>
</tr>
<tr>
<td>60-64</td>
<td>2865</td>
<td>2277</td>
</tr>
<tr>
<td>65-74</td>
<td>4142</td>
<td>3662</td>
</tr>
<tr>
<td>75-84</td>
<td>4504</td>
<td>4812</td>
</tr>
<tr>
<td>85+</td>
<td>5470</td>
<td>5736</td>
</tr>
</tbody>
</table>

Source: National Medical Expenditure Survey. Data are weighted to national totals. Pregnancy cost is $3996.

REFERENCE


