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PUBLIC POLICY, HEALTH INSURANCE AND THE TRANSITION TO ADULTHOOD

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Public Policy, Health Insurance and the Transition to Adulthood
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ABSTRACT

This paper assesses the impact of two recent policies designed to increase insurance coverage for older teens and young adults. The introduction of SCHIP in 1997 enabled low and moderate income teens up to age 19 to gain access to public health insurance. More recent policies adopted by a number of states have enabled young adults between the ages of 19 and (typically) 24 to remain covered under their parents' health insurance. We take advantage of the discrete break in coverage at age 19 to evaluate the impact of SCHIP. We also use quasi-experimental variation across states and years along with the targeted nature of eligibility to evaluate the impact of these "extended parental coverage" laws. Our results suggest that both types of policies were effective at increasing health insurance coverage, especially among their respective target populations. Overall, SCHIP increases insurance coverage by 3 percentage points; those with incomes under 150 percent of poverty are found to experience a 7 percentage point increase. We find little evidence of crowd-out associated with the introduction of SCHIP. Extended parental coverage laws have minimal aggregate effects on coverage, but they increase coverage by up to 5 percentage points for select groups. These laws may generate reverse crowd-out, as individuals leave public insurance coverage to take advantage of the private coverage now available to them.

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I. INTRODUCTION

For the past 25 years, public policy regarding health insurance has focused on increasing the rate of coverage among children and youth in the United States. As shown in Figure 1, these efforts have been quite successful, with insurance coverage rates among children under the age of 19 rising by roughly 15 percentage points between 1982 and 2007. In contrast, coverage rates among young adults (aged 19 to 29) remain as low as, and in some cases much lower than, they were 25 years ago. Less than two-thirds of 23-year-olds are currently covered by health insurance. Because of these relatively low rates of insurance coverage, young adults comprise a large fraction of all those without health insurance. Indeed, one-third of all the uninsured in the U.S. fall into this young adult age group.

As a result, public policy has increasingly focused on raising insurance coverage among young adults. One option is to expand existing public insurance programs that are targeted at children and teens, such as Medicaid and the State Children's Health Insurance Program (SCHIP), by increasing the maximum age to qualify for coverage. Children generally lose their Medicaid or SCHIP coverage on their 19th birthday, but recent proposals have sought to increase this age limit to as high as 25 years old.[1] A second option is to expand private insurance coverage among young adults by increasing the age of dependency for the purposes of health insurance coverage. Due to incentives of the tax code, children generally lose access to private insurance coverage through their parents at age 19 (or age 24 if they are full-time students). However, in recent years, several states have expanded eligibility to young adults through (typically) age 24 in certain population subgroups to be covered under their parents' employer-provided plan.

[1] See, for example, the Children's Health First Act (H.R.1535 and S.895), introduced in March 2007 by Rep. John Dingell and Sen. Hillary Clinton.

This paper provides evidence on the efficacy of these two policy options to increase health insurance coverage among young adults. While the existing literature has examined the take-up and crowd-out rates associated with SCHIP, it is unclear that these estimates are applicable to young adults. We provide new estimates of take-up and crowd-out, which are more relevant to the young adult population, by taking advantage of the discontinuity in SCHIP eligibility upon turning age 19 that was generated when the program was instituted in the late 1990s. "Extended parental coverage" laws are recent enough that no past research of which we are aware has examined their effect on take-up and crowd-out. We take advantage of variation across states and years in the introduction of these laws, along with other provisions that target certain population subgroups, to identify the effects of these extended parental coverage laws.

Our estimates suggest that both policies have been successful in increasing insurance coverage for older teens and young adults, although the aggregate impact is larger for SCHIP. We find that SCHIP increased insurance coverage by three percentage points for older teens as a whole. Those under 150 and 300 percent of the poverty line experienced a seven and four percentage point increase, respectively. We also find weaker evidence of crowd-out than in the existing SCHIP literature.

We also find that extended parental coverage laws are effective at increasing rates of insurance coverage for those groups targeted by the policy. These groups include those whose parents have *group* health insurance coverage and who work in firms that are unlikely to be exempt from state insurance regulation due to ERISA. For young adults in this subpopulation, coverage rates are found to rise by five percentage points, which corresponds to a 25 percent decline in uninsurance rates for this subpopulation. In addition, we find that effects are concentrated among non-students, which is consistent with the fact that many students are eligible for parental coverage even in the absence of these state laws. These laws are also found to have their largest effects in the income distribution on those between 150 and 300 percent of the poverty line. These young adults experience a 2.7 percentage point increase in coverage in response to extended parental coverage laws. Because the targeted groups are a subset of the full population of young adults, we are unable to find an impact of extended parental coverage laws on aggregate rates of insurance coverage. Our results also provide evidence of reverse crowd out; increases in private insurance coverage may be at least partially offset by reductions in public insurance coverage.

II. BACKGROUND

A. Legislative Background

While Medicaid policy in the 1980s and early 1990s dramatically expanded coverage for pregnant women, infants, and young children, coverage expanded relatively little for teenagers during this time. Teenagers who received Aid to Families with Dependent Children (AFDC) were always covered by Medicaid, but the income eligibility limits for AFDC – and therefore Medicaid – coverage were extremely low in most states. For example, income eligibility limits for 18-year-olds were 15 percent of the federal poverty line in Alabama, 28 percent in Florida, 41 percent in New Jersey, 33 percent in Ohio, and 17 percent in Texas (Rosenbach, et al., 2003).

The only substantial Medicaid expansion that affected teenagers was the provision of OBRA 1990 requiring that that all states provide Medicaid to children with incomes below the federal poverty line and birth dates after September 30, 1983 (U.S. Congress, 1994). This change led to a gradual expansion of Medicaid eligibility to older children and teenagers. For older teenagers, however, Medicaid eligibility remained extremely restrictive through 1997, when those born after September 30th, 1983 were no more than age 14.

This changed in 1997 upon the introduction of SCHIP, which was created by the Balanced Budget Act of 1997 (BBA 97) and enacted in August of that year. This legislation provided federal matching funds for states to expand coverage up to age 19 and to children in families with income beyond previous Medicaid eligibility levels. States could either expand their existing Medicaid programs or create new programs. States launched their SCHIP programs at different times, but all did so within a fairly narrow window. Most began enrolling children in 1998 and by the end of 1999 all SCHIP programs were up and running (Wooldridge, et al., 2003). States were allowed to cover children in families up to 200 percent of the poverty line, but they could extend those income limits even further, depending upon the income thresholds in their existing Medicaid programs. This provision generates some variability in income eligibility limits across states, although the range is very narrow for most states.

Table 1 details the variation in the implementation and provisions of Medicaid expansions, including SCHIP programs, across states around the time that SCHIP was introduced. In most states, children lost coverage at age 14 in 1997 and at age 19 in 1999. There is some variation in the income eligibility limits for SCHIP across states, with limits ranging from 100 to 300 percent of the federal poverty line; however, the majority of states ultimately used an eligibility limit of 200 percent of the poverty line.

BBA 97 intentionally allowed states to incorporate features intended to minimize crowd-out into their SCHIP programs. For example, states are permitted to impose waiting periods before enrollment in SCHIP and most states have chosen to do so (cf. Gruber and Simon, 2008). In addition, states are permitted to charge premiums and copayments for services, so long as total out-of-pocket costs (including premium) do not exceed 5 percent of family income. These features may be expected to reduce both take-up and crowd-out relative to rates that were observed in prior Medicaid expansions.

Like Medicaid coverage, private coverage through a parent's policy has traditionally ended on a child 19th birthday (or, for full-time students, the child's 24th birthday). Indeed, Anderson and Dobkin (2009) document a substantial discontinuity in private insurance coverage at exact age 19. These age limits for parent insurance coverage reflect the tax code, which specifies that the tax exclusion for health insurance includes coverage of a dependent child until the 19th (or 24th) birthday (Internal Revenue Service, 2008). While, in principle, an employer *could* have extended coverage to older children, the potentially complicated tax implications would have provided a very strong disincentive. Extended parental coverage laws have mandated that, in spite of the tax implications, certain firms must extend coverage to young adult children.

Virtually all extended parental coverage laws have been implemented by individual states over the past five years. While there have been proposals to extend parental coverage at the federal level, all of the laws to date have been implemented by states. The first such law was implemented in Utah in 1994 where young Mormon adults were losing health insurance while carrying out their traditional two years of missionary work (Associated Press, 2007). Members of the Church of Latter-Day Saints applied pressure to the state legislature and a policy allowing parents to cover these young adults was implemented. Kriss, et al. (2008) provides an extensive description of the laws introduced through 2007, most of which were implemented in the few years preceding that date.

Table 2 summarizes these extended parental coverage laws. The table includes the years when extended parental coverage laws were implemented, the limiting age of dependency status (coverage expires at that birthday), as well as whether the young adult must be a student, must be unmarried, or cannot have their own dependents. The limiting age of dependency status, or age up to which young adults are covered, is 25 in the majority of states that passed such laws, with a few states opting for age 24, 26 or 30. Nearly all states require young adults to be unmarried to qualify. While most states allow non-students to qualify under extended parental coverage laws, a minority require the young adult to be enrolled in school.[2] A few states require the young adult not to have any dependents of their own. As of the end of 2008, 25 states havd implemented laws.[3]

Eligibility for extended parental coverage depends not only on the characteristics of the adult child, but also on the characteristics of the parents' insurance coverage. Naturally, parents who do not have private insurance coverage for themselves will not be able to extend coverage to their adult children. In addition, roughly 55 percent of those parents who have employer-provided health insurance are covered by firms that self-insure (Pierron and Fronstin, 2008). Under the Employee Retirement Income Security Acts of 1974 (ERISA), self-insured firms are exempt from state-level health insurance regulation. The implication for extended parental coverage laws is that a substantial fraction of parents who have employer-provided health insurance do not have the option to extend it to their adult children under existing laws. We use this variation in some of our empirical specifications below.

B. Prior Literature

Although the decision to remain uninsured may be rational if young adults are risk neutral, there are reasons to believe that some form of market failure may also be at play here. Levy (2007) analyzes several reasons underlying the low and declining insurance rates among young adults. First, increasingly unstable employment patterns in this population, including

[2] Providing health insurance to students up until their 24th birthday is a nontaxable benefit even without these laws, but employers do not need to offer it. States that require school enrollment for extended parental coverall are expanding coverage by mandating the provision of this benefit as well as extending the age at which it is offered. In all of our empirical results presented below, we include these states as having an extended parental coverage law despite its potentially limited effect. We have also experimented with dropping these states and our results were virtually unchanged.

[3] See the notes to Table 2 for sources.

frequent job changes and employment in part-time and temporary positions, diminish the likelihood of employer-sponsored health insurance coverage. Levy (2007) reports empirical evidence that suggests that employment patterns among young adults contribute substantially to their low and declining insurance rates. Second, adverse selection, which has been widely documented in health insurance markets (cf. Cutler & Reber, 1998), implies actuarially unfair premiums for young adults in both group and non-group insurance markets and therefore lower insurance rates. Increasing health care costs have likely exacerbated this actuarial unfairness for healthy, young adults. Both of these potential explanations for falling insurance rates imply a potential role for the government. Even if the decision were rational for individual young adults, one could still argue that the government should intervene in this market if young adults' failure to purchase insurance generates negative externalities on others. For instance, Gross (2008) finds that young adults overuse expensive emergency room care because they lack health insurance and those costs may be transferred to others.

Prior public interventions into insurance markets through Medicaid expansions have been studied extensively. However, that literature provides many conflicting estimates of take-up and crowd-out. For example, Cutler and Gruber (1996) identify the effects of Medicaid expansions, using a measure of "simulated eligibility" that quantifies the legislative variation across states and years. They find crowd-out rates of 30 to 50 percent. Ham and Shore-Sheppard (2005), however, suggest that these results are sensitive to specification and choice of data set. Other papers, using longitudinal data sets, have found lower, but non-zero, crowd-out rates (Yazici and Kaestner, 2000; Thorpe and Florence, 1999; Blumberg Dubay, and Norton, 2000). Moreover, Card and Shore-Sheppard (2004) use a regression discontinuity design to analyze expansions of Medicaid that affects those under 100 percent and 133 percent of the poverty line and find that the coverage effect is low because take-up is low rather than because of large crowd out. On balance, the evidence on crowd-out in Medicaid expansions is quite mixed.

The more recent literature on SCHIP has also relied primarily on variation in program generosity across states and years to identify the effect on insurance coverage. Using this variation, LoSasso and Buchmueller (2004), for example, conclude that 9 percent of eligible children gained public insurance coverage, but that crowd-out was nearly 50 percent. Similarly, Gruber & Simon (2008) use the SIPP and find take-up rates ranging from 5 to 15 percent depending on the specification; their central crowd-out estimate is 60 percent, although the estimates that are most closely related to the results presented in the current paper are in the 20 to 40 percent range.

Our contribution, relative to this body of previous literature, is an entirely new identification strategy, which relies on the discontinuity in eligibility at age 19, to identify the impact of public insurance on coverage rates and crowd out.[4] Because our identification strategy is so simple, we are able to show the basis for many of our results in graphs. In addition, our evidence focuses on the marginal older teenager (aged 16-18) who is induced to participate in public insurance due to the introduction of SCHIP rather than the marginal child (aged 0-18) who is induced to participate, as in the prior literature. Therefore, our findings are more relevant for thinking about the likely effects of expanding SCHIP coverage to older ages.

The literature on parental extended coverage laws is much less developed because the laws have been introduced so recently. To our knowledge, this paper is the first to address the impact of these laws on the level of insurance coverage.

[4] Gross (2008) and Anderson and Dobkin (2009) also use the discontinuity of SCHIP eligibility at exact age 19 to identify the impact of insurance coverage on the use of emergency rooms, outpatient and inpatient services.

III. DATA

For our analysis, we use several years of data from the March Current Population Survey (CPS). The CPS is a monthly survey of about 50,000 households (although the sample size varies somewhat over time) conducted by the Bureau of Labor Statistics. Each month, information about employment status and demographic characteristics is obtained from individuals aged 16 and older. The March supplement to the CPS asks additional questions including whether individuals had private health insurance, public health insurance, or were uninsured (in more recent years) in the preceding calendar year.

Our analysis of SCHIP uses these data between the years 1992 and 2008, whereas our analysis of extended parental coverage laws uses the data between 2000 and 2008. These datasets provide reports of insurance coverage in 1991 to 2007 and 1999 to 2007, respectively. We use different sample windows for these exercises to focus more closely on the periods in which the laws changed.

Several potential concerns must be addressed when using the CPS health insurance variables. First, the CPS was re-designed in 1995, which led to some changes in the reporting of health insurance. This change should not contaminate our analysis, however, as long as the reporting changes did not vary discretely with age, which seems unlikely.[5] Second, there is a long-standing concern that CPS respondents may actually provide their current insurance status when asked about their insurance status in the prior year; thus, health insurance variables are potentially measured with error (Swartz, 1986). The quasi-experimental nature of our identification strategy, however, is unlikely to be seriously affected by this problem. All pre- and post-policy change years are grouped together so these errors are likely to generate only a relatively small number of miscodings. In addition, the recent literature on SCHIP has raised the

[5] We have verified that the results of our analysis are not sensitive to excluding the data collected prior to 1995.

concern that, as states outsource their public insurance plans to HMOs, some individuals who are actually on state Medicaid and SCHIP programs may incorrectly report that they are on non-group private health insurance. Such incorrect reporting could bias downwards estimates of both take-up and crowd-out. Indeed, Lo Sasso and Buchmueller (2004) find that their results are sensitive to the treatment of reported non-group private insurance coverage. We will also examine the sensitivity of our results to this issue.

Finally, the literature on Medicaid and SCHIP has raised questions about how to code individuals who report both public and private health insurance coverage in the prior year. If these individuals were on private health insurance at the beginning of the year and dropped it in order to take up public insurance, then coding them as being on both types of insurance coverage would bias downwards estimates of crowd-out. Gruber and Simon (2008) explored this issue using panel data in the SIPP and found that 40 percent of the children who reported both public and private insurance in the middle wave also reported overlapping public and private coverage before and after the middle wave. This finding suggests that it is, in many cases, appropriate to code these individuals as holding both private and public insurance. Nonetheless, we explore the sensitivity of our estimates to this concern below.

To capture the nature of the data available to us, we present Figure 2, which displays the age pattern of health insurance status over the 2000 to 2007 period. This figure is comparable to figures reported by Levy (2007). It shows that the majority of teens are covered by a parent's health insurance plan through age 18 and then parental coverage drops off sharply through the early 20s. Coverage from a public plan through age 18 captures a substantial share of those without parental coverage, but that drops off as well soon after that. Own private coverage, which includes spousal coverage, makes up a lot of the lost coverage through the late teens and early 20s, but not enough to prevent a sizeable drop-off in insurance coverage. Rates of coverage by any insurance rebound in the middle 20s, but even by age 35 do not reach the level of childhood coverage. Insurance coverage status measures for each individual in the relevant CPS surveys provide the dependent variable in our subsequent econometric analyses.

IV. THE IMPACT OF SCHIP

A. Identification Strategy

Our empirical strategy for identifying the effect of SCHIP on older teenagers relies on the age discontinuity in eligibility. As discussed earlier and shown in Table 1, the introduction of SCHIP created a discrete discontinuity in eligibility at age 19 that did not exist prior to BBA 97. There are, of course, other sources of variation in eligibility for SCHIP that have been exploited by other authors, such as differences in income eligibility limits by state and differences in the timing of SCHIP implementation by state. However, the variation across states was relatively small, with the vast majority of states choosing income eligibility cut-offs that were precisely, or very close to, 200 percent of the poverty line and most states implementing their SCHIP programs by the end of 1998. In comparison, the variation in SCHIP eligibility across different young adult age groups was quite dramatic. Therefore, our analysis asks whether there is any discontinuous change in insurance coverage for individuals under the age of 19 after BBA 97.

B. Descriptive Analysis

An advantage of this identification strategy is that any effects should be apparent in a simple, graphical analysis. Indeed, Figure 3A provides suggestive evidence that an age discontinuity in insurance coverage emerged after BBA 97. The figure shows the age pattern of insurance coverage in three different time periods: 1990-1995 (before SCHIP), 1999-2003 (immediately after SCHIP), and 2004-2007 (several years after SCHIP). At most ages, the profile of insurance coverage falls relatively uniformly over time. The exception to this pattern is an *increase* in insurance coverage for individuals under 19 years of age in the post-SCHIP periods relative to the pre-SCHIP time period. This suggests that the introduction of SCHIP generated a substantive increase in insurance coverage for older teenagers.

Public insurance coverage displays a similar pattern, as shown in Figure 3B. The age profile of coverage shows that the percentage of individuals covered by public insurance fell over time at most ages with the notable exception of 16-18-year-olds, who experience a relative increase in public insurance coverage in the post-SCHIP periods. Figure 3C provides the same graphical analysis for private insurance coverage. If crowd-out were a substantial issue for this age group, we would expect to see a relative decline in private insurance coverage for 16-18-year-olds after SCHIP was introduced. Interestingly, this figure suggests that there was, if anything, an increase in private insurance coverage. This figure provides suggestive evidence that the magnitude of crowd-out in this age group is unlikely to be large.

C. Econometric Specification

Our regression analysis formalizes this graphical analysis, using the following framework:

$$(1) \qquad Insurance_{iast} = \alpha + \beta Below19_a \cdot Post1998_t + \delta_a + \tau_t + \pi_s + \rho Unemp_{st} + X_{iast}\theta + \varepsilon_{iast}$$

where $Insurance_{iast}$ is a measure of insurance status for individual *i* in age group *a*, state *s* and year *t*. Our independent variable of interest is an indicator for being under the age of 19 after 1998.[6] The regression includes a full set of age, year, and state fixed effects. In addition, we control for differences in economic conditions across states, using the state- and year-specific

[6] In our analysis, we assume that the age reported in the March CPS corresponds to the individual's age in the prior year, the year for which health insurance variables are reported. However, we have confirmed that our results are not sensitive to excluding 19-year-olds—who might plausibly have been below 19 and therefore in the treatment group, rather than the control group, during the prior year.

unemployment rate.[7] We also control for individual-level covariates, including gender, marital status, student status, household income as a share of the poverty line, and the square of household income as a share of the poverty line. We use OLS to estimate the regressions and cluster our standard errors at the state level. We limit the sample to ages 16 to 22, so we are including only three years above and below the age cut-off. The implicit identifying assumption of our regression is that, in the absence of SCHIP, there would not have been differential trends in insurance coverage at ages around the age 19 eligibility limit.

D. Results

Table 3 shows our main results for SCHIP. Column 1 in this table displays the percentage of individuals in the relevant population subgroup that is covered by any form of insurance. This is a useful metric for gauging the magnitude of the estimated effects reported in the remainder of the Table. Each cell in Columns 2 through 4 reports the results from a different regression; the reported coefficient reflects the estimated value of β from equation (1), which is the coefficient on an indicator for being under the age of 19 after 1998. Each column shows coefficients and standard errors for a different dependent variable, and the rows show results for different subsamples.

The first row of Table 3 shows results for the full sample of 16 to 22 year olds. Our results suggest a statistically significant 3.0 percentage point increase in insurance coverage for older teenagers associated with the introduction of SCHIP (Column 2). The third column shows a larger, statistically significant 4.2 percentage point increase in public insurance and the fourth column shows a statistically insignificant 0.4 percentage point decrease in private insurance

[7] Results are similar when we also include state-specific linear trends in our econometric specification.

coverage.[8] The implied level of crowd-out is 10 percent (.004/.042), although this estimate is not significantly different from zero.

The lower rows of Table 3 explore heterogeneity in the effect of SCHIP by income category. Because we prefer a measure of household income that does not reflect the temporarily low incomes for otherwise middle- and high-income young adults who are transitioning into the labor market, we restrict this analysis to young adults who live with their parents. Therefore, the second row replicates our main results for the full sample of 16 to 22 year olds who live with their parents; the increase in insurance coverage is somewhat smaller than the results in the full sample, but the difference is small and not statistically significant.

For the lowest income group, with household incomes below 150 percent of the poverty line, we see a substantial, 6.7 percentage point increase in insurance coverage. Prior to SCHIP, 64 percent of this group was covered by insurance, so 36 percent remained uninsured. This estimated impact of SCHIP suggests that it reduced uninsurance rates by 19 percent. Estimates of crowd-out in this lowest income category are small. The 8.5 percentage point increase in public insurance is accompanied by a statistically insignificant 0.5 percentage point decline in private insurance coverage, suggesting a 6 percent crowd-out rate.

For the middle income group, with household incomes between 150 percent and 300 percent of the poverty line, take-up rates are lower, which is not surprising, and the overall increase in insurance coverage is smaller (a 4.2 percentage point increase, relative to a pre-policy uninsurance rate of 19 percent). Crowd-out for this income group is wrong-signed and statistically insignificant. For the highest income group, with household incomes above 300

[8] The coefficients from the public and private insurance models would sum to that for the any insurance model, if the two forms of insurance were mutually exclusive. The nature of the questions asked in the CPS, however, makes it possible for a respondent to report both types of coverage in the past year. We discuss this issue in more detail below.

percent of the poverty line, we expect little impact of the introduction of SCHIP, because very few states extended eligibility to teenagers in this part of the income distribution. Indeed, there are no statistically significant changes in insurance coverage in this income group, and all coefficients are very small.

In Table 4, we examine the time pattern of the change in insurance coverage patterns for individuals under the age of 19, to ensure that it corresponds to the timing of SCHIP implementation. For this section of the analysis, we use six separate independent variables of interest, which interact the dummy variable for being under the age of 19 with dummies for different time periods. The omitted time period is 1997, the year before changes in insurance coverage should have begun, so all coefficients are measured relative to 1997. Each column of the table reports the six key coefficients from a different regression.

The coefficients in the first two rows show that the difference in insurance coverage for those under the age of 19, compared to those over the age of 19, were not statistically different in the years leading up to the SCHIP legislation than it was in 1997; these coefficients provide evidence that our estimates are unlikely to be merely capturing pre-existing differential trends in insurance coverage for those under the age of 19. The third coefficient shows the change in insurance coverage for those under the age of 19 during the transition years, 1998 and 1999, when most SCHIP plans were being implemented. During this time period, we see no statistically significant difference in overall insurance coverage, although the second column does provide evidence of a small, statistically significant increase in public insurance coverage. Beginning in 2000, the regression shows a substantial and statistically significant increase in both public insurance coverage and overall insurance coverage. This rise in insurance coverage increases somewhat over time, which is consistent with overall SCHIP enrollment patterns, which nearly doubled between 2000 and 2007 (Smith et al., 2008). Our estimates for private insurance coverage, reported in the third column, provide only statistically insignificant evidence of crowd-out, with point estimates that never exceed a 15 percent crowd-out rate.

E. Alternative Measures of Crowd-Out

The prior literature on crowd-out suggests two alternative ways to measures insurance status, which could potentially increase our estimate of crowd-out. One alternative is to consider any increase in non-group private insurance to be an increase in public insurance coverage. This alternative measure of public insurance addresses the concern that CPS respondents who are on SCHIP programs might erroneously report that they had private insurance coverage. Crowd-out, in this case, should be measured as the reduction in group insurance coverage, as a percentage of the increase in public and non-group insurance coverage. When we follow this approach, our estimates of crowd-out rise. Specifically, we find a 5.3 percentage point increase in public plus non-group insurance, and an insignificant 1.5 percentage point decline in group insurance coverage, for an overall crowd-out rate of 28 percent. However, crowd-out among the lowest income group – those with household incomes below 150 percent of the poverty line – remain low at only 10 percent.

Another alternative approach is to measure crowd-out as the reduction in reporting *only* private insurance coverage, as a percentage of the increase in public insurance coverage. This approach addresses the concern that individuals who are in the process of switching from private to public health insurance coverage may report both types of coverage. This approach is the most conservative possible solution to this problem, because it assumes that all individuals who report overlapping public and private coverage are being crowded out. As mentioned above, Gruber & Simon's (2008) investigation of this issue in the SIPP suggests that this approach is likely too conservative, since roughly half of the individuals who report having both public and private insurance continue to do so in later interviews. Nonetheless, we used this approach and found a statistically significant decline in private insurance coverage, resulting in an overall crowd-out rate of 29 percent. Among the lowest income groups, the crowd-out rate is 21 percent and remains statistically insignificant. We consider these estimates to be upper bounds on crowd-out among older teenagers due to SCHIP expansions.

While our estimates of crowd-out are lower than those in the prior literature, we do not believe that they are necessarily inconsistent with the prior literature. Our identification strategy relies on a marginal SCHIP participant who is a 16- to 18-year-old who was made eligible for public insurance by SCHIP's expansion of the age eligibility limit. In contrast, the identification strategy in Gruber and Simon (2008) and Lo Sasso and Buchmueller (2004) relies on differences in SCHIP eligibility across all age groups and states. Their strategy, therefore, relies on a marginal SCHIP participant who is likely to be much younger and with higher income. For example, prior to the introduction of SCHIP, the minimum income eligibility limit for 1-5-year-olds was 133 percent of the federal poverty line, whereas the typical 18-year-old had to meet the dramatically lower AFDC eligibility rules. So, if higher income individuals have lower take-up rates and higher crowd-out rates, it may not be surprising that our analysis estimates lower crowd-out. However, for the purposes of predicting crowd-out rates for further expansions of public insurance to older age groups – including recent proposals to expand SCHIP eligibility to adults in their early 20s – our lower estimates of crowd-out in SCHIP are likely more relevant than estimates from the prior literature.

V. THE IMPACT OF EXTENDED PARENTAL COVERAGE LAWS

A. Identification Strategy and Econometric Specification

The second portion of our analysis focuses on the impact of extended parental coverage laws. As shown in Table 2, there is substantial variation across states in the implementation of these laws. Our basic identification relies on standard quasi-experimental variation comparing changes in outcomes in states that adopt laws before and after their implementation relative to changes in outcomes over time in states that did not introduce these laws. In addition, since the provisions of these laws in states that adopted them made certain population subgroups eligible and others ineligible, we also adopt triple-difference estimation strategies that provide even stronger methods of drawing causal conclusions.

Our basic regression framework is as follows:

$$(2) \qquad Insurance_{iast} = \alpha + \beta Law_s \cdot Post_t + \delta_a + \tau_t + \pi_s + \rho Unemp_{st} + X_{iast}\theta + \varepsilon_{iast}$$

where $Insurance_{iast}$ is a measure of insurance status for individual i in age group a, state s and year t. Our independent variable of interest is an indicator for living in a state that has adopted an extended parental coverage law in a year after the law was implemented. As in the SCHIP analysis, the regressions include a full set of age, year, and state fixed effects. We again control for differences in state-specific economic shocks using the state- and year-specific unemployment rate, and we control for individual-level covariates, including gender, marital status, student status, household income as a share of the poverty line, and the square of household income as a share of the poverty line. We continue to use OLS to estimate the regressions and cluster our standard errors at the state level.

To begin our analysis, we use a sample of all young adults between the ages of 19 and 24, the age group that would be affected by laws, to estimate the aggregate impact of the law using standard difference-in-difference methods.[9] We then separate that sample into population subgroups that we would expect to be more and less affected by its introduction. Initially, we restrict the sample to include just those who meet the eligibility criteria in their state. For states

[9] We exclude all data from Massachusetts in this analysis, because of the broader health insurance reform that was implemented at the same time as the extended parental coverage law.

that have not passed extended parental coverage laws and therefore have not established eligibility criteria, we assume that the only eligibility criterion is that the young adult must be unmarried. As an additional test, we run a triple-difference specification that uses the ineligible young adults in the states as the third difference. This approach uses the ineligibles within the states as an additional control group that can adjust for other within-state differences that may be affecting levels of health insurance coverage in the state. We continue our analysis by further separating the sample into narrower population subgroups, who we might expect to be more and less likely to be affected by extended parental coverage laws.

B. Results

Tables 5 and 6 present the results of our analysis. As in Table 3, the first column of these tables displays the percentage of individuals in the relevant population subgroups who are covered by any form of insurance. In Columns 2 through 4 of both tables, every cell represents the results from a different regression. Different dependent variables are shown in different columns, and different samples and specifications are shown in different rows.

The first row of Table 5 shows the difference-in-difference results for the full sample. We do not find that extended parental coverage laws are associated with a statistically significant increase in insurance coverage overall or private insurance coverage in this model. In the second row, we focus more narrowly on those young adults who would be eligible for this form of coverage. Perhaps not surprisingly, we see larger effects for this group; private insurance coverage is estimated to increase by a statistically significant 1.9 percentage points. The magnitude of the effect on private insurance coverage is larger in the triple-difference specification in the third row, at 3.8 percentage points.

Interestingly, the net effect on overall insurance coverage in both specifications is muted by an offsetting decline in public insurance coverage, which is statistically significant in the difference-in-difference specification. This finding suggests the possibility of "reverse crowd-out" resulting from extended parental coverage laws. When young adults are able to move onto their parents' private insurance plans, some of them appear to drop public insurance to do so. This reverse crowd-out might be expected to occur if the private insurance plan provides more generous coverage or better access to providers.[10]

In the second panel of Table 5, we restrict our analysis to individuals who are not students. Since the tax incentive for employer-provided health insurance has traditionally extended to students up to the age of 24, the impact of extended parental coverage laws should be much smaller for students than for non-students. In the bottom of Table 5 we show results for non-students who should be more strongly affected.[11] These results show that, indeed, increases in private insurance coverage associated with extended parental coverage laws are concentrated among non-students. The triple difference estimate for the sample of non-students indicates that these laws increased private insurance coverage by 5.6 percent, compared to 3.8 percent for the full sample. However, Panel B also shows that the decreases in public insurance coverage are not statistically significant. Although the additional evidence that is presented below is uniformly consistent with reverse crowd-out, the findings in Panel B of Table 5 lead us to consider our overall evidence for reverse crowd-out to be suggestive, rather than conclusive.

Table 6 reports the results of additional difference-in-difference models by more narrowly defined population subgroups. To identify these subgroups, we need identifying information in the data that is only available on the parents' record, so we are required to restrict

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[10] We conducted one additional test to validate the notion of reverse crowd-out. Among those in their early 20s, public insurance coverage largely results from welfare receipt and most welfare recipients are female. As an additional test, we estimate these models separately by gender. The results indicate that reverse crowd-out is largely attributable to women, as we might suspect.

[11] We do not show results for students, because there are so very few ineligible students that our triple-difference empirical strategy is far less compelling for this subsample

the population to those respondents living with their parents. Because the main reason a young adult would be ineligible for extended parental coverage is marriage and few married young adults live with their parents, the triple-difference estimating strategy does not make sense in this context. In this part of the analysis, we restrict our attention to difference-in-difference models relying just on those eligible for extended parental coverage, as described earlier. The first row of this table reports the results of estimates comparable to the second row of Table 5 (eligible sample) except that it focuses on just those young adults living with their parents. As described earlier, we impose this sample restriction so that we can observe relevant household characteristics that enable us to better identify who may be affected by these laws. The results from this specification are similar to those in Table 5.

The next rows of the table show effects by income category. These laws clearly affect a different income group than SCHIP introduction does. Increases in private insurance are statistically insignificant for young adults in households with incomes below 150 percent of the poverty line. This finding is not surprising, since the parents in these households are unlikely to have private health insurance to extend to their adult children. The impact of the extended parental coverage – in terms of an increase in private insurance coverage and a decrease in public insurance coverage – is concentrated almost entirely among the group with incomes between 150 and 300 percent of the poverty line. For this group, extended parental coverage laws are associated with a 6.4 percentage point increase in private insurance coverage, a 2.4 percentage point decrease in public insurance coverage, and a 2.7 percentage point increase in overall insurance coverage. This 2.7 percentage point increase in insurance coverage is relative to a pre-policy uninsurance rate of 33 percent, so it represents an 8 percent decline in the prevalence of uninsurance in this income group. Interestingly, there is not a statistically significant increase in insurance coverage for those with incomes over 300 percent of the poverty line. In the pre-policy period, though, 85 percent of these young adults had private insurance, so there was less scope for expanding insurance coverage in this income group than in the lower income groups.

In the bottom rows of Table 6, we examine heterogeneity in the effects by parents' insurance coverage. We expect the impact of extended parental coverage laws to be strongly concentrated among the children of parents who are covered by group health insurance plans. Although some states included non-group insurance policies in their extended parental coverage laws, most focused on group health insurance policies. Moreover, since group health insurance policies are typically much less expensive than non-group health insurance policies, young adults who become eligible for extended parental coverage have a much stronger incentive to take advantage of it if their parents have group insurance. In fact, we see that the effects of the extended parental coverage laws – including increases in private insurance and decreases in public insurance – are concentrated among those whose parents have group insurance coverage.

In addition, among children whose parents have group health insurance, we expect the effects of extended parental coverage laws to be strongest for those whose parents work for relatively small firms. This prediction arises from the ERISA legislation of 1974, which exempts the insurance policies of self-insured firms from state regulation. As a result, roughly half of those who have employer-provided health insurance have insurance policies that are not affected by these legal changes.[12] In the absence of ERISA's effects, this prediction would likely be reversed, since small employers face higher expected premiums and therefore are less likely to offer health insurance (cf. Ellis and Ma, 2009). While we cannot observe in the CPS which parents work in self-insured firms, we do know that the probability of self-insurance increases dramatically with firm size: In firms with less than 200 employees, only 12 percent of

[12] According to the 2008 Kaiser/Health Research and Education Trust (HRET) Employer Heath Benefits Survey, 55% of workers who receive health insurance through their employer are in self-funded insurance plans.

employees were in self-funded plans; in contrast, 89 percent of employees of firms that have 5,000 or more employees are in self-funded plans (Henry J. Kaiser Foundation and HRET, 2008).

We take advantage of this fact and use firm size as a proxy for self-insurance status. We create two further sub-samples among the children of parents with group insurance coverage: those whose parents work for firms with less than 100 employees and those whose parents work for firms with more than 100 employees.[13] We expect to find larger effects in the smaller firms, because these smaller firms are substantially more likely to be required to follow the state insurance law.

Indeed, we find our largest increase in private insurance coverage in the group of children whose parents have group insurance coverage through a firm with fewer than 100 employees. For this group, we find that private insurance coverage increases by a statistically significant 6.7 percentage points, compared to a 1.5 percentage point increase in private insurance coverage among those whose parents work for larger firms.[14] The 4.9 percentage point increase in overall insurance coverage for this group is substantial, particularly when compared to the pre-policy uninsurance rate of 19.5 percent for this particular population. We also continue to find evidence of reverse crowd-out. The decline in public insurance coverage is considerably larger for children of parents who work in small firms, as we would expect if it was caused by the extended parental coverage laws.

This set of findings has two important implications: First, it provides some validation of our empirical strategy, by showing that the overall effects that we estimate are driven by the sub-

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[13] We also experimented with alternative firm size cut-offs (like 500 employees) and obtained qualitatively similar results.

[14] We also find that, for those whose parents are included in a group plan, there is a decrease in nongroup coverage and an increase in group coverage. Individuals who used to pay for their own individual plans appear to drop them and join their parents' plan. This transition likely saves money for these individuals and potentially provides them with superior insurance.

sample that should be most strongly affected by extended parental coverage laws. Second, it provides some explanation for the relatively small overall effects of extended parental coverage laws. Even if a young adult is legally eligible for a state-level extended parental coverage law, he or she can only take advantage of that eligibility if his or her parent has group insurance coverage through a firm that does not self-insure. Among those who are both legally eligible and have parents who are able to extend coverage, the take-up rates – and the corresponding reductions in uninsurance – are quite high.

VI. DISCUSSION

The results of our analysis suggest that both the introduction of SCHIP and extended parental coverage laws were effective in increasing rates of insurance coverage for older teens and young adults, at least for certain population subgroups. SCHIP was particularly effective at increasing insurance coverage for older teens who were under 150 percent of the poverty line without evidence of substantial crowd-out. Extended parental coverage laws had a sizeable impact on the insurance coverage of non-students, young adults between 150 and 300 percent of the poverty line and those whose parents worked for relatively small firms that had group insurance plans. We also found suggestive evidence that the increase in insurance coverage that this policy generated may have been accompanied by a reduction in public insurance coverage (reverse crowd-out). The magnitude of the coverage effects is somewhat larger in response to the introduction of SCHIP than the introduction of extended parental coverage laws, but both led to meaningful increases in coverage for their target populations.

Determining whether policymakers should endorse extending either or both of these programs requires a weighing of their advantages and disadvantages as implemented in the past and as they would be extended going forward. In the case of SCHIP, for instance, our policy experiment addressed the discontinuity of coverage at age 19 created when SCHIP was introduced. An expansion of this policy could increase the maximum age of coverage to, say 22 or 25. It is not clear whether or not the impact of such an extension would be the same as that of the earlier policy.

One issue to consider would be differences in crowd-out. If anything, our prior would be that crowd-out would be smaller for this group than it is for younger children and, in fact, our estimates of crowd-out based on the introduction of SCHIP were smaller than those from earlier Medicaid expansions. We hypothesize that the reason for this is that the marginal "child" gaining coverage in SCHIP is a teen who is lower in the income distribution than the younger children who gained coverage in earlier Medicaid expansions. Uninsured young adults would similarly be more likely to be drawn from the lower parts of the income distribution and, if our hypothesis is correct, crowd-out may be less important for this group.

Another factor that may suggest even larger effects for young adults is that they have fewer options for coverage. Young children have the ability to be included in their parents' plan if their parents are covered; parental coverage is quite high for younger children as shown in Figure 2. Young adults would have to rely on their own private coverage as their alternative to a public plan, yet private coverage rates are relatively low for young adults (also shown in Figure 2).

To provide a ballpark estimate of the number of additional young adults who may be covered if SCHIP was extended to cover individuals up to their 22nd or 25th birthday, we ignore these caveats for the moment. We assume that the impact on the incremental groups covered is the same as we have estimated for those at age 19 with the same extent of crowdout. Based on our results in Table 3, we apply the aggregate estimate that insurance coverage of any type would rise by 3 percentage points to simulate this statistic. We combine this value with population estimates from the 2005-2007 American Community Surveys indicating that there are about 13.3 million individuals between the ages of 19 and 21 and another 12.2 million between 22 and 24. If 3 percentage points more members of these groups received health insurance in response to an SCHIP expansion, this would increase coverage by 399,000 and 366,000 individuals, respectively.

We also hypothesize that extended parental coverage laws would have larger effects than those we have estimated if they are implemented at the federal level. A key constraint on the effectiveness of these laws as currently enacted at the state level is that ERISA significantly restricts the target population. As discussed earlier, 55 percent of those parents who have employer-provided health insurance are covered by firms that self-insure and ERISA excludes those plans from state regulation. If implemented at a national level, that restriction would not exist. Moreover, state level plans require children to be residing in the same state as the parents and that would not need to be true if the policy were implemented at the federal level.

To simulate the effect of a national extended parental coverage law for those ages 19 to 24, we apply the estimate obtained from parents receiving group insurance from the smaller firms that do not self-insure to approximate the impact of a law that was not constrained by ERISA. For that sample, insurance rates rose 5 percentage points. Without the constraint of ERISA, we assume this impact would extend to all of those whose parents held group coverage, although this group is only around three-quarters of the population. Therefore, we apply the 5 percentage point estimate along with the three-quarters group coverage estimate to the 25.5 million individuals between the ages of 19 and 24 to get a simulated coverage increase of 956,000 young adults. This would include those who have already received coverage as a result of the state laws.

The question of who pays for the insurance coverage is also a relevant issue. We find some support for the notion that extended parental coverage laws generate reverse crowd out in that some individuals may drop public insurance coverage for what may be superior private insurance coverage. This shifts the cost of the coverage from the taxpayer to the individual's family and the insurance company. To the extent that the family pays, such an outcome may be optimal. To the extent that insurance companies are unable to sufficiently assess and spread risk in the presence of adverse selection, it may be suboptimal.

In the end, policymakers will be required to weigh these trade-offs in determining whether to introduce either or both of these policies. Nevertheless, our analysis has provided evidence indicating that both types of policies as implemented have increased insurance coverage among the targeted groups.

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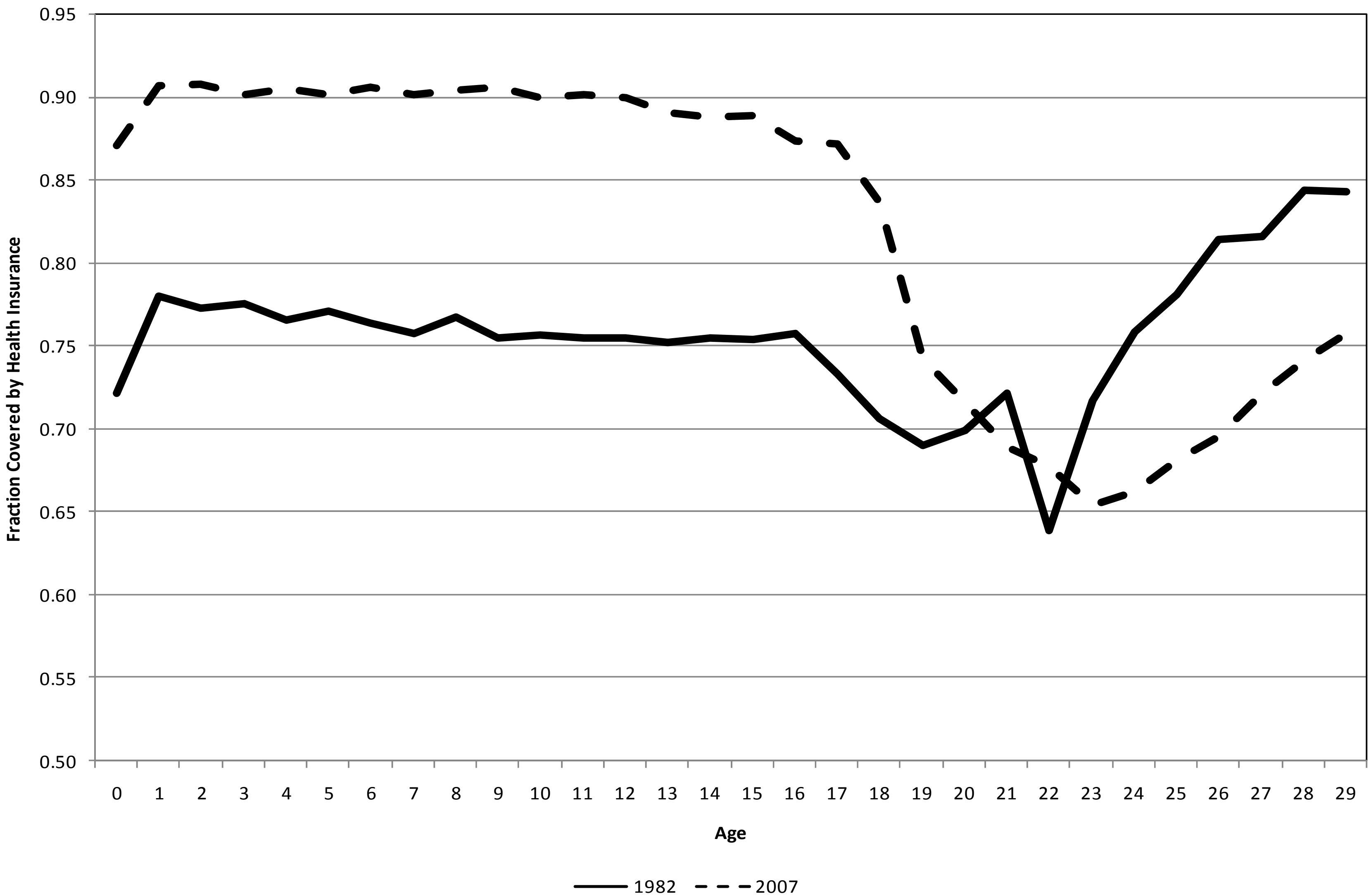
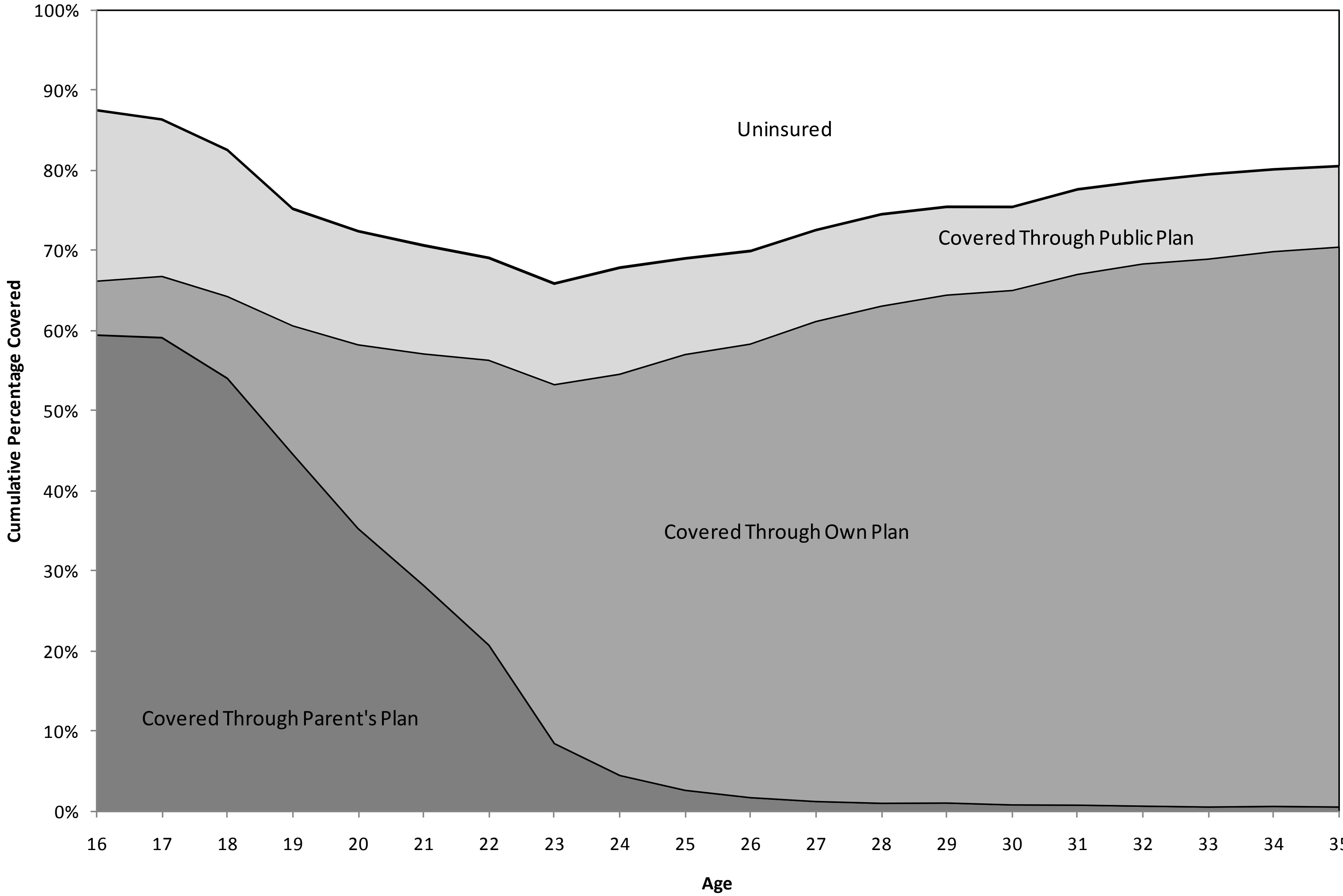
Figure 1: Rate of Health Insurance Coverage by Age, 1982 and 2007



Source: Authors' calculations from the Current Population Survey.

Figure 2: Health Insurance Coverage Status: 2000-2007, by Age



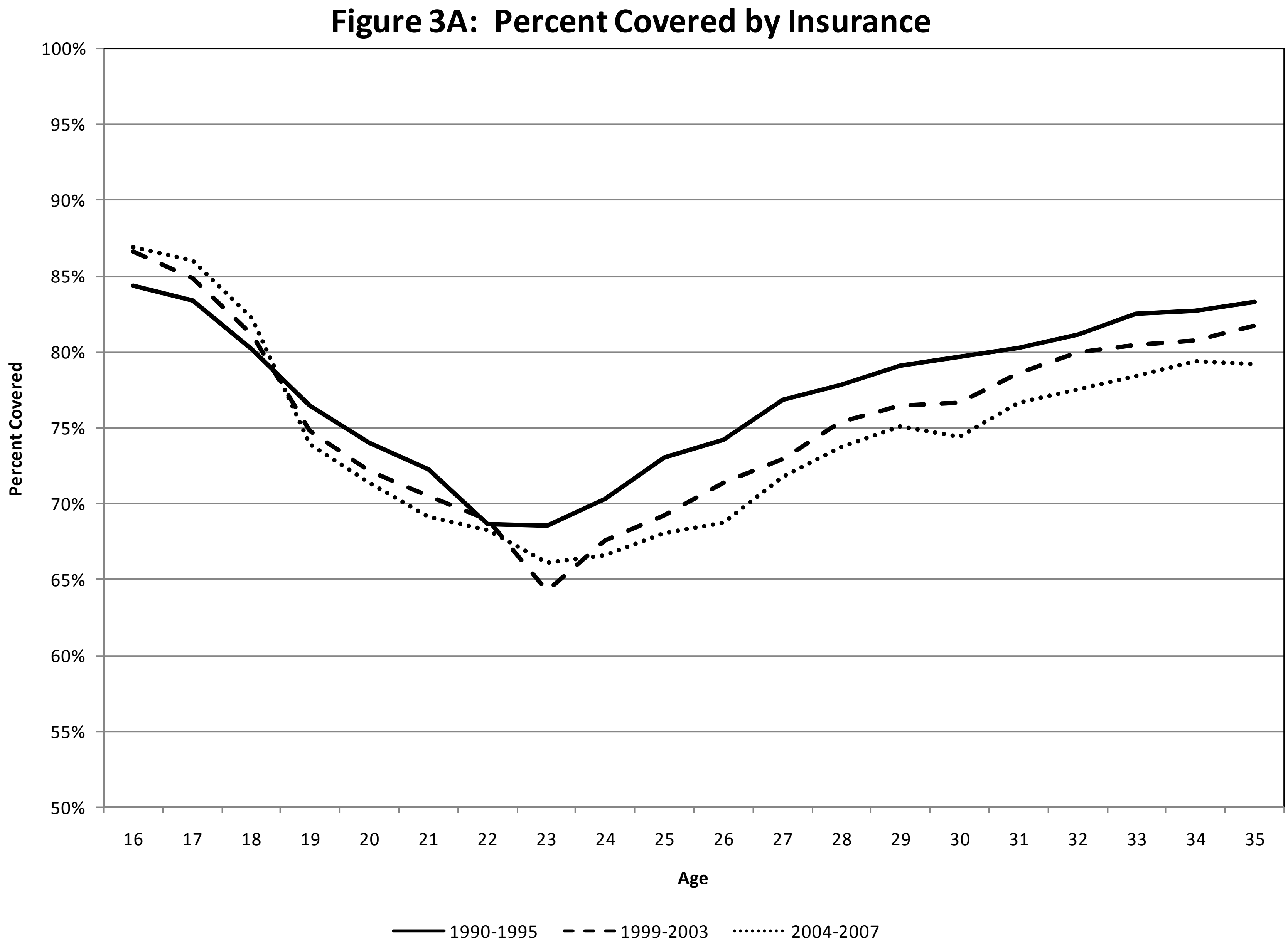
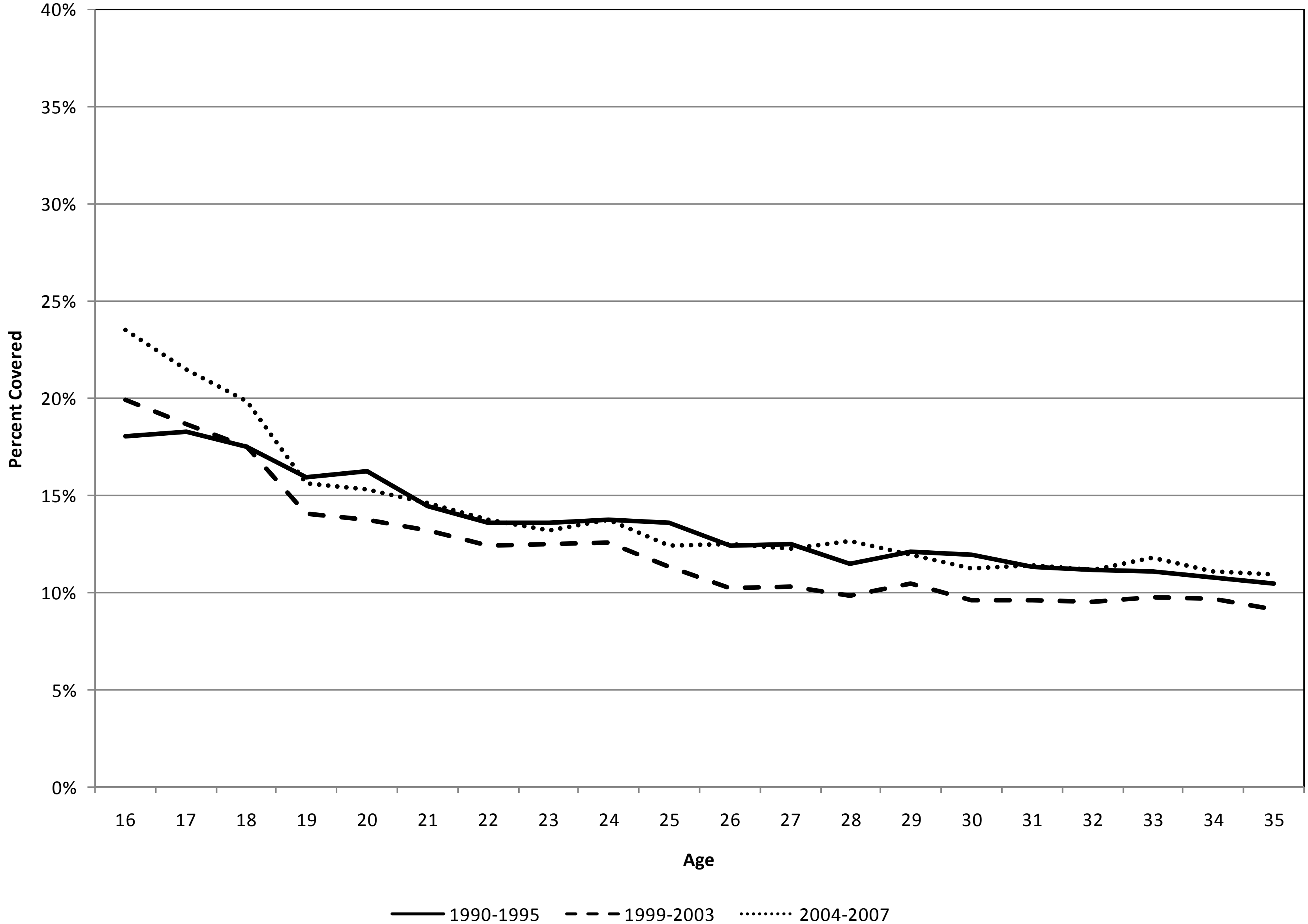


Figure 3A: Percent Covered by Insurance
100%
95%
90%
85%
80%
75%
70%
65%
60%
55%
50%
Percent Covered
16
17
18
19
20
21
22
23
24
25
26
27
28
29
30
31
32
33
34
35
Age
1990-1995
1999-2003
2004-2007

Figure 3B: Percent Covered by Public Health Insurance



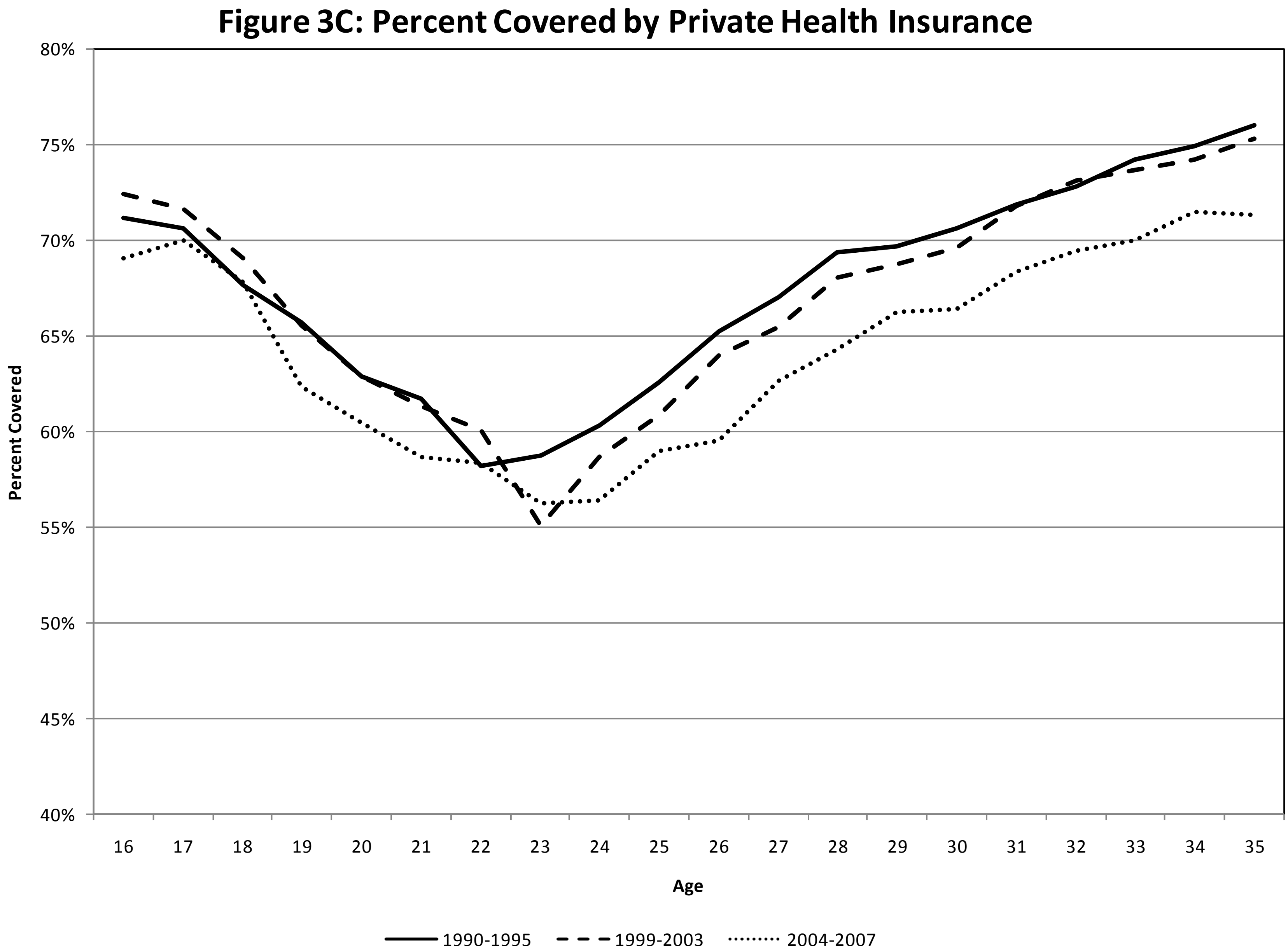


Figure 3C: Percent Covered by Private Health Insurance
80%
75%
70%
65%
60%
55%
50%
45%
40%
Percent Covered
16
17
18
19
20
21
22
23
24
25
26
27
28
29
30
31
32
33
34
35
Age
1990-1995
1999-2003
2004-2007

Table 1: Changes in Income and Age Eligibility Limits Associated with Medicaid Expansions and the Introduction of SCHIP

	Percent of Federal Poverty Line				Upper Age Limit		
State	1995	1997	1999	2001	1995	1997	1999
Alabama	100	133	200	200	12	14	19
Alaska	100	100	200	200	12	14	19
Arizona	100	100	150	200	16	18	19
Arkansas	100	100	200	200	12	14	19
California	100	100	200	250	12	14	19
Colorado	100	100	185	185	12	14	19
Connecticut	185	185	300	300	12	14	19
Delaware	100	100	200	200	19	19	19
District of Columbia	100	100	200	200	12	14	19
Florida	100	100	200	200	12	19	19
Georgia	100	100	200	200	19	19	19
Hawaii	300	300	300	200	19	19	19
Idaho	100	100	150	150	12	14	19
Illinois	100	100	185	185	12	14	19
Indiana	100	100	150	200	12	19	19
Iowa	100	100	185	200	12	14	19
Kansas	100	100	200	200	12	16	19
Kentucky	100	100	100	200	12	19	19
Louisiana	100	100	133	200	12	14	19
Maine	125	125	185	200	19	19	19
Maryland	185	185	200	200	12	14	19
Massachusetts	200	100	200	200	12	14	19
Michigan	150	150	200	200	16	18	19
Minnesota	275	275	275	275	18	18	19
Mississippi	100	100	100	200	12	14	19
Missouri	100	100	300	300	18	18	19
Montana	100	100	100	150	12	14	18
Nebraska	100	100	185	185	12	14	19
Nevada	100	100	200	200	12	14	19
New Hampshire	185	185	300	300	19	19	19
New Jersey	100	100	200	350	12	14	19
New Mexico	100	185	235	235	12	19	19
New York	160	100	230	250	15	14	19
North Carolina	100	100	200	200	18	18	19
North Dakota	100	100	100	140	12	18	19
Ohio	100	100	150	200	12	14	19
Oklahoma	100	100	185	185	12	14	18
Oregon	100	100	170	170	12	19	19
Pennsylvania	185	185	235	235	12	14	19
Rhode Island	100	250	250	250	12	19	19
South Carolina	100	100	150	150	12	14	19
South Dakota	100	100	140	200	12	19	19
Tennessee	100	400	200	200	12	18	19

Texas	100	100	100	200	12	14	19
Utah	100	100	200	200	18	18	19
Vermont	225	225	300	300	18	18	18
Virginia	100	100	185	200	19	19	19
Washington	200	200	200	250	19	19	19
West Virginia	150	150	150	200	19	19	19
Wisconsin	100	100	100	185	12	14	19
Wyoming	100	100	100	133	12	14	19

Sources: Shore-Sheppard (2003), Centers for Medicare and Medicaid Services (2004), and National Governors Association (various years).

Table 2: Extended Parental Coverage Law Details

State Name	Year Law was Implemented	Limiting Age of Dependency	Must be a student	Must be unmarried	Cannot have dependents
Colorado	2006	25		X	
Delaware	2007	24		X	X
Florida	2007	25		X	
Idaho	2007	25	X	X	
Indiana	2007	24			
Maine	2007	25		X	X
Massachusetts	2006	25			
New Hampshire	2007	26		X	
New Jersey	2006	30		X	X
New Mexico	2005	25		X	
Rhode Island	2007	25	X	X	
South Dakota	2005	24	X		
Texas	2003	25		X	
Utah	1994	26		X	
Virginia	2007	25	X		
Washington	2007	25		X	
West Virginia	2007	25		X	

Notes: Entries in table show whether law was in effect as of December of that year. Age of dependency is birthday at which dependency expires.

Sources: Kronstadt, et al. (2007), Kriss, et al. (2008) National Conference of State Legislatures (2008), and our own reading of state laws.

Table 3: The Effect of SCHIP on Insurance Coverage

	Percent Covered by Any Insurance	Estimation Results for Dependent Variable: Any Insurance Coverage	Public Insurance Coverage	Private Insurance Coverage
	(1)	(2)	(3)	(4)
Full Sample	76.1	0.030**	0.042**	-0.004
N=270,636		(0.004)	(0.005)	(0.004)
Sample Living with Parents	81.8	0.026**	0.037**	-0.003
N=202,783		(0.005)	(0.005)	(0.004)
By Income:				
<150% of poverty line	63.9	0.067**	0.085**	-0.005
N=38,066		(0.010)	(0.014)	(0.010)
150-300% of poverty line	80.6	0.042**	0.054**	0.004
N=71,818		(0.008)	(0.006)	(0.007)
>300% of poverty line	91.9	0.001	0.005*	-0.003
N=92,899		(0.004)	(0.003)	(0.004)

Notes: Each cell in Columns 2-4 presents the results from separate difference-in-difference regressions models. Control variables include state, year, and age fixed effects, the state- and year-specific unemployment rate, and individual-level covariates. Standard errors are clustered on state. The sample includes all observations between the ages of 16 and 22, using the 1992-2008 CPS.

** denotes statistical significance at the 5% level.

* denotes statistical significance at the 10% level.

Table 4: The Effect of SCHIP on Insurance Coverage
Dynamic Specification

	Estimation Results for Dependent Variable:		
	Any Insurance Coverage (1)	Public Insurance Coverage (2)	Private Insurance Coverage (3)
Under Age 19			
* Years 1991-1993	-0.004	-0.002	-0.002
	(0.009)	(0.007)	(0.010)
* Years 1994-1996	-0.006	-0.006	-0.001
	(0.007)	(0.006)	(0.008)
* Years 1998-1999	0.007	0.013**	-0.002
	(0.009)	(0.006)	(0.008)
* Years 2000-2002	0.019**	0.033**	-0.005
	(0.008)	(0.007)	(0.009)
* Years 2003-2005	0.032**	0.045**	-0.004
	(0.009)	(0.008)	(0.009)
* Years 2006-2007	0.037**	0.053**	-0.005
	(0.009)	(0.008)	(0.010)

Notes: Each column presents the results from separate difference-in-difference regressions models. The omitted period is 1997. Control variables include state, year, and age fixed effects, the state- and year-specific unemployment rate, and individual-level covariates. Standard errors are clustered on state. The sample includes all observations between the ages of 16 and 22, using the 1992-2008 CPS. N=270,636.

** denotes statistical significance at the 5% level.

* denotes statistical significance at the 10% level.

Table 5: The Effect of Extended Parental Coverage Laws on Insurance Coverage

	Percent Covered by Any Insurance (1)	Estimation Results for Dependent Variable: Any Insurance Coverage (2)	Public Insurance Coverage (3)	Private Insurance Coverage (4)
Panel A: Full Sample				
Difference-in-Difference	69.0	-0.004	-0.012*	0.008
Full Sample		(0.008)	(0.007)	(0.010)
N=112,625				
Eligible Sample	69.2	0.002	-0.016**	0.019**
N=88,514		(0.008)	(0.007)	(0.009)
Triple Difference				
Full Sample	69.0	0.015	-0.010	0.038**
N=112,625		(0.019)	(0.015)	(0.015)
Panel B: Non-students				
Difference-in-Difference				
Full Sample	62.2	-0.003	-0.010	0.008
N=69,288		(0.010)	(0.011)	(0.013)
Eligible Sample	59.9	0.013	-0.013	0.031**
N=48,225		(0.010)	(0.010)	(0.015)
Triple Difference				
Full Sample	62.2	0.037	-0.004	0.056**
N=69,288		(0.023)	(0.011)	(0.021)

Notes: Each cell in Columns 2-4 presents the results from separate difference-in-difference regressions models. Control variables include state, year, and age fixed effects, the state- and year-specific unemployment rate, and individual-level covariates. Standard errors are clustered on state. The sample includes all observations between the ages of 19 and 24, using the 2000-2008 CPS.

** denotes statistical significance at the 5% level.

* denotes statistical significance at the 10% level.

Table 6: The Effect of Extended Parental Coverage Laws on Insurance Coverage among Targeted Populations

	Percent Covered by Any Insurance	Estimation Results for Dependent Variable:		
		Any Insurance Coverage	Public Insurance Coverage	Private Insurance Coverage
	(1)	(2)	(3)	(4)
All Living with Parents	74.7	0.011	-0.015**	0.029**
N=53,292		(0.009)	(0.007)	(0.008)
By Income:				
<150% of poverty line	53.3	-0.006	-0.029	0.006
N=7,480		(0.031)	(0.026)	(0.030)
150-300% of poverty line	66.9	0.027*	-0.024*	0.064**
N=*17,359*		(0.014)	(0.012)	(0.016)
>300% of poverty line	85.1	0.003	-0.006	0.009
N=28,453		(0.011)	(0.006)	(0.013)
By Parents' Insurance				
Group Coverage	84.1	0.009	-0.020**	0.028**
N=39,073		(0.008)	(0.006)	(0.008)
And firm size<100	80.5	0.049**	-0.038**	0.067**
N=9,064		(0.020)	(0.011)	(0.018)
And firm size>100	85.1	-0.003	-0.014**	0.015*
N=30,009		(0.009)	(0.006)	(0.009)
No Group Coverage	49.0	0.006	-0.003	0.021
N=15,582		(0.022)	(0.017)	(0.015)

Notes: Each cell in Columns 2-4 presents the results from separate difference-in-difference regressions models. Control variables include state, year, and age fixed effects, the state- and year-specific unemployment rate, and individual-level covariates. Standard errors are clustered on state. The sample includes observations on all targeted groups between the ages of 19 and 24, using the 2000-2008 CPS. All targeted groups are restricted to those who live with their parents and are eligible for coverage in treatment states and are unmarried in control states.

** denotes statistical significance at the 5% level.

* denotes statistical significance at the 10% level.