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USING STATE-LEVEL EVIDENCE TO INFORM NATIONAL POLICY: RESEARCH FROM THE STATE HEALTH ACCESS REFORM EVALUATION (SHARE) PROGRAM

# How Have State Policies to Expand Dependent Coverage Affected the Health Insurance Status of Young Adults?

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**Research Objective.** Nearly one in three adults of ages 19–29 lack health insurance, representing the highest uninsured rate of any age group. To help address this gap, 38 states have enacted laws requiring insurers to permit young adults to enroll as dependents on their parents' plans. This paper evaluates their impact on coverage for young adults.
**Study Design/Methods/Data.** This study uses data for individuals ages 19–29 from the Current Population Survey's Annual Demographic Supplement for calendar years 2000–2008. Linear probability models are used to obtain difference-in-differences estimates of the impact of dependent coverage expansions in 19 early-adopting states on young adults' insurance status. The models also address possible policy endogeneity due to the nonrandom enactment of expansion policies across states.
**Principal Findings.** State young adult dependent coverage policies yielded small increases in dependent coverage ranging from 1.52 percentage points for all young adults to 3.84 percentage points for those ages 19–25 residing with parents. These increases were largely offset by declines in employer-sponsored insurance (ESI) in the young adults' own name. No significant impact on young adult uninsured rates was observed.
**Conclusions and Implications.** Adult dependent coverage expansions have had a relatively small impact on enrollment as an ESI dependent and appear to have the unintended consequence of reducing ESI policyholder coverage. This policy did not achieve a reduction in uninsured rates as policy makers had intended. Federal reform efforts to expand dependent coverage are likely to be more successful because reform will be accompanied by subsidies and enrollment mandates.

**Key Words.** Health care financing, insurance, premiums, health policy, politics, law, regulation, health economics

Among the U.S. population, young adults are more likely to lack health insurance than any other age group, with potentially serious consequences for their health and financial well-being. In 2008, 28.6 percent of persons ages

18–24 and 26.5 percent of those between 25 and 34 lacked coverage. By comparison, less than a fifth of persons ages 35–64 and less than a tenth of children under age 18 were uninsured (DeNavas-Walt, Proctor, and Smith 2009).

The relatively high uninsured rate of young adults has important implications for their access to health care, protection against the financial consequences of illness, and may indirectly impact their future health and health care needs. For example, uninsured young adults are three to four times more likely than their insured peers to delay or forgo medical care due to costs and two to four times less likely to see a medical provider, have a usual source of care, or fill a prescription due to cost (Callahan and Cooper 2005; Nicholson et al. 2009). Additionally, uninsured young adults are twice as likely as those with coverage to have trouble paying medical bills and to have medical debt (Nicholson et al. 2009). Lack of coverage may also compromise young adults' ability to address their frequently observed obesity and alcohol and tobacco use that lead to health and economic problems in adulthood (Merluzzi and Nairn 1999).

Several reasons help to explain why young adults lack coverage. Completion of high school or college frequently results in loss of eligibility as a dependent on a parent's health plan. Young adults who fail to obtain post-secondary education or high-skilled vocational training may lack the human capital necessary for jobs that provide health insurance. Such transitions lead to sharp increases in young adult uninsured rates: 38 percent of high school graduates who did not go to college were without coverage for some subsequent period, and after turning age 19, uninsured rates increase to nearly 29 percent for young adults ages 19–29, up from 11 percent for children 18 and under (Nicholson et al. 2009).

State efforts to improve access to affordable coverage through regulation of small group and individual health insurance markets may have unintended consequences for young adults. In states that tightly constrain premium variation based on individual health status, young adults may face premiums that fail to reflect their actuarial risk and together with their relatively low incomes,

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may make coverage economically unattractive. Apart from affordability, some young adults may have a low demand for coverage due to their relatively good health, attitudes toward risk taking, and lack of information regarding the health and financial consequences of going without coverage.

In response to the significant disparity in coverage for young adults relative to other age groups, by 2008, 21 states had implemented legislation requiring private insurers to expand dependent coverage. As shown in Table 1, Utah was the first state to do so in 1995, permitting adults through age 25 to enroll in a parent's plan.[1] Between 2003 and 2006, six other states followed suit, including New Jersey, which implemented the most expansive dependent eligibility (through age 29 with an expansion through age 30 beginning January 2009). In 2007 through January 2008, 14 additional states implemented expansion policies. In addition, as of mid-August 2009, 8 other states had enacted adult dependent coverage expansions that were implemented after January 2008 or not yet implemented (National Conference of State Legislatures 2009).

Among enacting states, requirements for eligibility vary on the basis of age limits, marital status, residence with parents, transitions from prior insurance, among other factors. In all cases, state laws do not apply to self-funded employer benefit programs due to their exemption from state regulation under a provision of the 1974 Employee Retirement Income Security Act (ERISA). Such an exemption will likely limit the reach of the expansion legislation because many large employers offer self-funded health benefits.[2] Of the 21 states implementing expansions through January 2008 (Table 1), 19 increased eligibility of nonstudents an average of 5.3 additional years, and 14 increased eligibility for full-time students by an average of 3.5 years.

In this paper, we address the question of whether state implementation of expanded dependent coverage has been effective in increasing coverage among young adults. We do so through an econometric analysis of the relationship between the implementation of this policy and its impact on young adults' health insurance status. We address the issue of policy endogeneity, consider the expansion's effect on different groups of young adults, examine the timing of the legislation's impact, and consider how the distribution of young adult coverage might change were all states to implement the expansions. For our most expansive sample, young adults ages 19–29, we find that state expansion legislation had a small impact on their insurance status, increasing coverage as a dependent on employer-sponsored insurance (ESI) by 1.52 percentage points (an 8.5 percent increase in such coverage over the 17.9 percent of targeted young adults in this group with dependent coverage in the preintervention period). For young adults ages 19–25 who live with their

Table 1: States Implementing Adult Dependent Coverage Expansions by January 2008

| | | *Postreform Maximum Age of Dependent Coverage (No. Years of Eligibility Added)* | |
|---|---|---|---|
| *State* | *Implementation Date* | *Students** | *Nonstudents* |
| Colorado | January 2006 | 24 (1) | 24 (6) |
| Delaware | June 2007 | 23 (0) | 23 (2) |
| Florida | July 2007 | 24 (0) | 24 (6) |
| Idaho | July 2007 | 24 (2) | 20 (2) |
| Illinois | July 2004 | 26 (2) | 21 (2) |
| Indiana | July 2007 | 23 (0) | 23 (4) |
| Maine | September 2007 | 24 (2) | 24 (6) |
| Maryland | January 2008 | 24 (2) | 24 (6) |
| Massachusetts† | January 2007 | 25 (2) | 20 (2) |
| Minnesota | January 2008 | 24 (0) | 24 (6) |
| Missouri | January 2008 | 24 (6) | 24 (6) |
| Montana | January 2008 | 24 (2) | 24 (6) |
| New Hampshire | September 2007 | 25 (7) | 25 (7) |
| New Jersey‡ | January 2006 | 29 (7) | 29 (11) |
| New Mexico | July 2003 | 24 (0) | 24 (6) |
| Rhode Island | January 2007 | 24 (0) | 18 (0) |
| South Dakota§ | July 2007 | 29 (7) | 18 (0) |
| Texas | January 2004 | Unlimited | 24 (6) |
| Utah† | January 1995 | 25 (5) | 25 (5) |
| Virginia | July 2007 | 24 (0) | 24 (6) |
| West Virginia | July 2007 | 24 (1) | 24 (6) |

*Limited to full-time students in most cases.

†Excluded from analyses of change in coverage (see text).

‡Maximum age was increased to 30 in January 2009.

§Maximum age for students was increased from 22 to 23 in 2005 and to 29 in July 2007. Tabulations do not include states that extended or considered extending dependent coverage to children with mental or physical disabilities, illnesses, or injuries that prevent them from attending school full time or for children who serve in the military and then return to school full time.

*Source*: Authors' review of public records available through offices of state insurance commissioners and National Conference of State Legislatures (2007, 2009).

parents, ESI dependent coverage increased by 3.84 percentage points (an 11.9 percent increase over the 32.4 percent of targeted young adults with dependent coverage in the preimplementation period). In all cases, we also find that the increase in dependent coverage was largely offset by a reduction of coverage as an ESI policyholder. We find no impact on young adult uninsured rates, suggesting that the expansion legislation may have had an unintended consequence of reducing young adult coverage as an ESI policyholder.

## Data and Methods

The data in this study are obtained from the Current Population Survey's (CPS) Annual Demographic Supplement for 2001–2009, corresponding to calendar years 2000–2008. Using young adults as the units of observation, we fit linear probability models and obtain difference-in-differences (DD) estimates of the impact of coverage expansions on their health insurance status.[3] In doing so, we consider the full array of insurance possibilities, including coverage as a dependent on an ESI plan; coverage as an ESI policyholder; coverage through nongroup insurance; public coverage; and being uninsured.

Our analysis also recognizes that public policy adoption by states is unlikely to be random. Failure to account for omitted and unobserved factors that may be correlated with the implementation of coverage expansions and outcomes of interest can yield biased impact estimates. To address such policy endogeneity, we use the state as the unit of observation in a separate equation and model the likelihood that a state adopted coverage expansion as of 2007. We focus on a 2007 cross-section of states because enactment of such legislation (with the exception of Utah) is tightly clustered among the years just before and including 2007 (Table 1). Following work by Stream (1999) and Besley and Case (2000), we apply logistic regression to predict the likelihood of policy adoption, using attributes of the state's political, economic, fiscal, and regulatory environment as explanatory variables. The details of this analysis are available from the authors upon request. As in Besley and Case (2000), we include the statistically significant correlates of enactment as control variables in our models of young adult coverage status to reduce possible bias in our impact estimates.

We estimate the following linear probability model to assess the impact of state expansion provisions on young adult coverage as

$$\begin{aligned}\mathrm{COV}_{ist} = {} & a_1 + a_2\mathrm{STATE}_s + a_3\mathrm{YEAR}_t + a_4\mathrm{TREND}_t \\ & + a_5(\mathrm{STATE}_s \times \mathrm{TREND}_t) + a_6\mathrm{TARGET}_{is} + a_7\mathrm{POLICY}_{ist} \\ & + a_8(\mathrm{TARGET}_i \times \mathrm{POLICY}_{ist}) + a_9 X_{ist} + a_{10}\mathrm{ADOPT}_{st} + e_{ist}\end{aligned}$$

In this model, $\mathrm{COV}_{ist}$ represents the coverage status of the $i$th young adult in state $s$ at time $t$. The coefficients for STATE and YEAR represent state and year-specific fixed effects to account for time-invariant differences across states that may contribute to coverage differences and for year-specific differences in coverage outcomes, respectively. TREND is a linear time trend to account for secular changes in coverage apart from policy implementation and state effects, and the interaction term STATE $\times$ TREND accounts for any time-varying state-specific changes in coverage. TARGET is set to 1 for young adults eligible for the expansions based on age, student, and marital status requirements in their state (regardless of implementation year) and is 0 for all others (ineligible young adults in expansion states as well as those in nonexpansion states). This variable accounts for time-invariant differences across these two groups that may be associated with their coverage status. POLICY is equal to 1 for all years that a state's coverage expansion was in effect (year of implementation and beyond) and is 0 otherwise. This variable controls for secular changes in coverage among states that adopted reform in their post-implementation periods. We also consider an alternative to POLICY defined for specific periods (for the first year of implementation, the second year, and subsequent years) to capture any time-dependent impact of the expansions on coverage status.

Coefficient $a_8$ on TARGET $\times$ POLICY is the DD estimate of the impact of reform. This coefficient captures the change in coverage status for the targeted group of young adults after policy implementation, relative to a control group of young adults in reform states not targeted by the legislation and those in states that never adopted the expansions. The DD estimator also nets out the effect of any unobserved and time-invariant differences between states that implemented expansions and those that did not, and together with our other controls for endogeneity, lends a causal interpretation to our findings. The POLICY and TARGET variables were constructed using information from departments of insurance in states implementing dependent coverage expansions on or before January 1, 2008. Three states are excluded from the analysis. Utah is excluded because its policy was implemented well before the study period and Massachusetts and Hawaii are excluded because young adult coverage is likely to be influenced by their respective individual and employer coverage mandates.

To complete our specification, the vector $X$ includes young adult characteristics that are likely to affect coverage status (described in notes to Table 2). In samples of young adults living with their parents, we include variables to characterize the parent policyholder's firm size. We do so to account for the fact that expansion legislation is not applicable to self-insured firms that are typically large, and because the price of ESI is likely to decline as firm size increases. Finally, $\text{ADOPT}_s$ is a vector of time-varying state-level variables obtained from the enactment analysis noted above to address the bias from possible policy endogeneity. These variables include the political affiliation of the state's governor, the dominant political party in the state legislature (for both legislative chambers), whether the state has a budget

Table 2: Difference-in-Difference Estimates of the Impact of State Health Insurance Expansions on Young Adult Insurance Status, 2000–2008

| | *Percentage Point Change in Insurance Status (Standard Errors in Parentheses)* | | | | |
|---|---|---|---|---|---|
| *Sample* | *ESI Dependent* | *ESI Policyholder* | *Private, Nongroup* | *Public* | *Uninsured* |
| A. Ages 19–29 ($N$ = 228,465) | 1.52** | − 0.97 | − 0.64 | 0.09 | 0.57 |
| | (0.69) | (0.84) | (0.46) | (0.63) | (0.94) |
| | $R^2 = 0.2679$ | $R^2 = 0.2205$ | $R^2 = 0.0177$ | $R^2 = 0.1085$ | $R^2 = 0.1593$ |
| B. Ages 19–29, living with parents ($N$ = 70,890) | 2.58*** | − 2.11* | − 1.11 | 0.97 | 0.46 |
| | (1.19) | (1.14) | (0.77) | (0.93) | (1.34) |
| | $R^2 = 0.4395$ | $R^2 = 0.1808$ | $R^2 = 0.0360$ | $R^2 = 0.1154$ | $R^2 = 0.1705$ |
| C. Ages 19–25 ($N$ = 144,206) | 2.77*** | − 2.36*** | − 0.26 | 0.43 | 0.50 |
| | (0.81) | (0.90) | (0.53) | (0.71) | (1.04) |
| | $R^2 = 0.3259$ | $R^2 = 0.1957$ | $R^2 = 0.0210$ | $R^2 = 0.1027$ | $R^2 = 0.1479$ |
| D. Ages 19–25, living with parents ($N$ = 61,172) | 3.84*** | − 3.33*** | − 1.03 | 0.58 | 0.56 |
| | (1.29) | (1.14) | (0.85) | (0.96) | (1.37) |
| | $R^2 = 0.4513$ | $R^2 = 0.1496$ | $R^2 = 0.0397$ | $R^2 = 0.1190$ | $R^2 = 0.1809$ |

*Notes.* Coefficients for DD estimator POLICY × TARGET are obtained from linear probability models of young adult health insurance status. All regression models have been weighted using CPS survey population weights and standard errors have been adjusted for the CPS complex sampling design following Davern et al. (2007). Models include state and year fixed effects, a linear time trend, the interaction of state and the linear time trend, a variable indicating whether the young adult is targeted by the expansion legislation, and a variable indicating the postpolicy implementation years. Other controls include age, gender, race/ethnicity, marital status, educational attainment, status as a full-time or part-time student, whether young adult is in fair/poor health, family income as percent of federal poverty line, percent of state population that are college graduates, and variables to control for the state's political, regulatory, and economic environment including whether the state has democratic governor and legislature; a budget surplus; the number of state insurance benefit mandates; unemployment rate; share of young adults in state population. Samples of young adults living with their parents include variables characterizing the firm size of the parent-policyholder: workplace size $<10$ employees; 10–24 employees; 229–99 employees; 100–499 employees; 500–999 employees (1,000 or more employees is the reference group).

*$p<.10$; **$p<.05$; ***$p<.01$ for two-tailed test.

CPS, Current Population Survey; DD, difference-in-differences; ESI, employer-sponsored insurance.

surplus, the number of mandated insurance benefits, the state unemployment rate, and the percentages of young adults and college graduates in the state. Finally, $e_{ist}$ is a stochastic error term.

We initiate our empirical analysis by focusing on young adults ages 19–29 and examine the period from 2000 to 2008, considering the impact of coverage expansions among states that had at least 1 year of postimplementation experience. We also examine the sensitivity of our findings to different sample definitions. We consider subsamples of young adults residing with their parents. We do so because for such households CPS data allow us to link a young adult's dependent coverage directly to a parent policyholder, thereby providing a more precise measure of dependent coverage. Additionally, in such households we are also able to identify the firm size of the parent ESI policyholder. Because large firms are more likely to be self-insured and thus exempt from state regulation under ERISA, we are able to control for this factor in this subsample. We also consider findings for a subsample of young adults ages 19–25 because 25 represents the upper age limit targeted by all but two states. Finally, we apply population weights to all our coverage regressions and adjust standard errors for the CPS' complex sample design following Davern et al. (2007).[4] We also test the sensitivity of our findings to an alternative standard error adjustment and to use of population weights.

## RESULTS

In Table 2, we present DD estimates of the impact of state expansion policies on young adult health insurance status. Table rows are defined for the different samples of young adults described above. In (D), we constrain our sample to consist of young adults of ages 19–25 who reside with their parents. We do so because this group may be more likely to take advantage of the expansions: they are likely to be financially dependent on their parents and have fewer alternatives to coverage than those older than 25 who have access to their own ESI or that of a spouse. Full regression results for samples A and D are available in an Appendix SA2.

We find largely consistent results across all these samples: implementation of expansion coverage yields a relatively small increase in the likelihood of ESI dependent coverage for young adults, but it is offset by a decline in their likelihood of being an ESI policyholder. The smallest increment in dependent coverage is obtained for our full sample (a 1.52 percentage point increase representing a 8.5 percent increase over the 17.9 percent of targeted young adults with ESI dependent coverage in the preimplementation period), and the largest increment obtained for sample D, young adults ages 19–25 residing with their parents (an increase of 3.84 percentage points or an 11.9 percent increase over the 32.4 percent of targeted young adults with dependent coverage preimplementation). We expected a somewhat larger impact of the expansions on the young adults of sample D compared with those of sample A, because the former may be more financially reliant on their parents, have fewer opportunities for comparable coverage, and thus be more predisposed toward taking advantage of the expansions. Additionally, most states require that young adults be financially dependent on their parents and that nonstudents must reside in the same state as their parents. However, the difference between these percentage point increments is statistically significant only at $p<.10$ (for a one-tail $t$-test).

In data not shown, we also considered the impact of the expansions on young adults not living with their parents. For those ages 19–29, we find a 2.37 percentage point increase in dependent coverage and a 2.94 percentage point increase for those ages 19–25. Although these groups experience a decline in ESI policyholder coverage (by 1.33 and 2.18 percentage points, respectively), the reductions are not statistically significant.

More generally, for each of our findings for ESI dependent status, we find that that the offsetting decline in ESI policyholder is statistically equivalent (i.e., $t$-tests indicate that the difference between the absolute values of the DD coefficients for dependent and policyholder status are not statistically significant). We find no evidence that the increase in dependent coverage was accompanied by a decline in the likelihood that a young adult would be uninsured or enrolled in nongroup or public insurance.

### *Timing of Expansion Effects*

The increments in dependent coverage displayed in Table 2 represent an average effect over the postimplementation period for states that adopted reform. However, the length of time for the impact of such a policy to be observed is of relevance to policy makers who face raised expectations regarding the speed with which interventions obtain discernable results. To address this issue, we specify DD estimators defined as the product of TARGET and each of three variables representing the year of policy implementation; year 2 of implementation; and the third and succeeding years of implementation.

Results presented in Table 3 are revealing. For our full sample A of young adults, we find that the impact of the expansion on dependent coverage occurs in the third and subsequent years of implementation (a 2.47 percentage point increase in the likelihood of such coverage). We find some evidence of such timing for all young adults ages 19–29 (sample B), but this finding is not statistically significant. For all young adults ages 19–25 (sample C), we find that the impact takes 3 or more years to be observed (a 3.98 percentage point increase, $p < .01$) and some weak evidence of an impact in the implementation year (a 2.14 percentage point increase, $p < .10$). For young adults ages 19–25 living with parents (sample D) deemed most likely to be most responsive to the expansions, we find strong evidence of an immediate response within the first year of implementation (a 4.27 percentage point increase) as well as a response 3 or more years postimplementation (a 4.11 percentage point increase). Finally, for young adults not living with parents we also find some evidence of a response 3 or more years postimplementation: a 4.06 percentage point increase for those ages 19–29, and a 4.52 percentage point increase for those ages 19–25 (data not shown). Thus, for most of our sample, the effect of implementation does not appear immediately; instead, it takes 3 or more years to have an impact and approach a steady state.[5]

*Implementing Expansion Coverage for All Young Adults*

To evaluate how the distribution of young adult coverage might change were all states to implement the expansions, we present the results of a simulation exercise (Table 4). We begin by predicting the distribution of dependent coverage from our regression model for a standard population consisting of all young adults ages 19–29, first by assuming that no young adult resides in a state that implemented reform (the values of our DD estimator POLICY $\times$ TARGET is set to 0), and next by assuming all young adults are in states that have implemented expansions (POLICY $\times$ TARGET is set to unity). In both cases, we permit each observation to retain its value for variables POLICY and TARGET (because the impact estimates are based on the DD coefficients, we want the independent effects of POLICY and TARGET to cancel out). We predict the likelihood of each type of coverage using our most expansive sample A under each scenario and then compute the weighted proportions of young adults with each type of coverage. We repeat this exercise for sample D, which yielded our strongest findings.

Our results for sample A reveal very little change in the coverage distribution were the expansion policy applied to all young adults. However, for

Table 3: Timing of Impact of Dependent Coverage Expansions on Young Adult Dependent Coverage

| | *Percentage Point Change in Insurance Status (Standard Errors in Parentheses)* | | | |
|---|---|---|---|---|
| *Year since Enactment* | *A. Ages 19–29* *N = 228,465* | *B. Ages 19–29, Living with Parents* *N = 70,890* | *C. Ages 19–25* *N = 144,206* | *D. Ages 19–25, Living with Parents* *N = 61,172* |
| First year of implementation | 0.96 | 3.12* | 2.14* | 4.27** |
| | (1.01) | (1.75) | (1.17) | (1.89) |
| Second year | 0.83 | 2.02 | 1.83 | 2.90 |
| | (1.17) | (1.97) | (1.36) | (2.12) |
| Third year and beyond | 2.47*** | 2.46 | 3.98*** | 4.11** |
| | (0.96) | (1.75) | (1.15) | (1.92) |

*Notes.* Coefficients are for DD estimators TARGET × first year of implantation, TARGET × second year of implementation, and TARGET × third year of implementation and beyond. Estimates are obtained from linear probability models of the likelihood that a young adult will be covered as an ESI dependent.

*$p<.10$; **$p<.05$; ***$p<.01$ for two-tailed test.

DD, difference-in-differences; ESI, employer-sponsored insurance.

Table 4: Simulation: Predicted Impact of Expansion Policies on the Distribution of Young Adult Insurance Coverage

| | *Sample A: All Young Adults, Ages 19–29* | | *Sample D: Young Adults, Ages 19–25, Living with Parents* | |
|---|---|---|---|---|
| *Type of Coverage* | *No Expansion* | *Expansion* | *No Expansion* | *Expansion* |
| | Percent Distribution | | | |
| ESI dependent | 21.2 | 22.7 | 40.6 | 44.4 |
| ESI policyholder | 31.1 | 30.2 | 14.5 | 11.2 |
| Private, nongroup | 5.9 | 5.3 | 7.1 | 6.1 |
| Public | 12.6 | 12.7 | 11.6 | 12.2 |
| Uninsured | 30.0 | 30.6 | 25.4 | 26.0 |

*Note.* Column totals may exceed 100% since individuals can be covered by more than one source of health insurance in CPS data.

CPS, Current Population Survey; ESI, employer-sponsored insurance.

young adults in sample D, we observe a more substantive 3.8 percentage point increase in dependent coverage. This increase mainly reflects the decline in coverage as an ESI policyholder and small decline in nongroup coverage.

*Sensitivity Analyses*

We implemented three different analyses to test the robustness of our main findings for ESI dependent coverage. First, we used a different clustering adjustment to compute standard errors for our DD estimates because these were obtained using the household as the clustering unit, do not account for possible error correlation at the geographic unit for which the expansion policies were implemented, and thus could be understated (Moulton 1990). Additionally, as Angrist and Pischke (2009) report, clustering standard errors at the state level has the advantage of correcting for any serial correlation that might be present in our data from unobserved shocks that persist over time in specific geographic areas. Implementing this alternative, we found very small differences in the standard errors of our DD coefficient compared with our original adjustment. Only in one case (for ESI dependent coverage in sample A) were standard errors elevated sufficiently to reduce significance from $p < .05$ to $p < .10$.

Next, we considered whether the magnitude of our findings might be sensitive to functional form by fitting logit models and predicting the likelihood of ESI dependent coverage. We applied methods described by Norton, Wang, and Ai (2004) to obtain the marginal effect of interaction terms comprised of dummy variables (i.e., the impact of our POLICY $\times$ TARGET variable). Our predictions obtained in this way were within a standard error of the DD coefficient in the linear probability models. Using sample A, we obtained a predicted 1.50 percentage point change in the probability of ESI dependent coverage from the logit compared with a 1.52 increase from the linear probability model. The results for samples B were a 2.53 percentage point increase from the logit and a 2.58 increase from the linear probability model, for sample C the estimates were 2.13 and 2.77 percentage point increases for the logit and linear probability models, respectively, and those for sample D were 3.56 and 3.84, respectively.

As a final sensitivity test, we examined findings from unweighted linear probability models with standard errors clustered at the household level. We applied this test because disagreement remains among researchers regarding the use of survey population weights in econometric analyses (see Deaton 1997 and Angrist and Pischke 2009 for contrary views). The results from this test yielded statistically significant estimates for all DD coefficients ($p < .01$), ranging in magnitude from 20 percent less to nearly 60 percent in excess of the weighted estimates (a 2.39 percentage point increase in dependent coverage for sample A; a 3.62 increase for B; a 3.29 increase for C; and a 4.74 increase for D). Thus, the unweighted estimates confirm our original econometric findings albeit at an elevated impact.

## Discussion

Thirty-eight states have enacted laws requiring health insurance carriers to extend dependent coverage to young adults. Policy maker motivations for adopting these laws are clear. Nationwide, more than one in four young adults are without health insurance, more than any other age group. Young adults are less likely than their older counterparts to be offered ESI, and state regulations to improve affordability of coverage by limiting premium variation by age have made coverage less affordable for young adults. Dependent coverage expansions are also attractive to state governments because they do not require appropriation of new state resources nor impose significant burdens on employers and insurers.

Despite the appeal of these expansions, our econometric analysis of CPS data for calendar years 2000–2008 reveals an increase in young adult dependent coverage of between 1.52 and 3.84 percentage points following policy implementation (an increase of between 8.5 and 11.9 percent over the average rate of dependent coverage among targeted groups in the preimplementation period) that was largely offset by a decline in ESI policyholder coverage. We also found no concomitant decline in the likelihood that young adults targeted by these policies were uninsured. Our results are robust to a variety of empirical specifications including alternative adjustments for sample clustering, unweighted models, and for estimates accounting for the impact of young adult dependent coverage expansions over time following policy adoption.

*Limitations*

Our findings should be considered in light of several limitations in CPS data. First, we cannot identify whether a parent lives in an expansion state for those young adults living away from home. Second, self-funded health benefit plans are exempt from state insurance regulations under ERISA, but the CPS lacks information about the ERISA status of plans. Third, most states limit expanded dependent coverage to young adults who are financially dependent on their parents and the CPS lacks such measures. Models limited to young adults living in a parent's household mitigate these limitations.

## CONCLUSIONS

Our study provides strong evidence that dependent coverage expansions resulted in a small increase in young adult dependent coverage but did not reduce their likelihood of being uninsured as policy makers intended. The apparent switching that we observe between own-name ESI and expanded dependent coverage may reflect more attractive out-of-pocket premiums or benefits available through a parent's plans compared with those available to young adults through their own employer. For example, the out-of-pocket premiums required to add an additional dependent to a parent's family coverage may be very low or even zero for some parents with ESI. This outcome is also not surprising given young adults are more likely to be employed in low-wage jobs than their parents. In addition to failing to directly address the problem of affordability, the success of young adult dependent expansions may be limited by other barriers, including ERISA preemption of self-insured health benefits, eligibility rules in many states that exclude married young adults and those with dependents, as well as state young adult dependent coverage rules requiring "creditable prior coverage" in order to reduce adverse risk selection.

At this writing, expanded dependent coverage for young adults up to age 26 is to be implemented within 6 months of the March enactment of the Patient Protection and Affordable Care Act (P.L. 111–148). The impact of this national reform provision on uninsured young adults is likely to be greater than the voluntary state programs that we evaluated. In the near term (before 2014), by superseding the ERISA preemption of large, self-insured firms from state dependent coverage expansions, the national reform law will make such employers subject to its dependent coverage expansion. Thus, as part of the newly enacted comprehensive health reform law, expanded dependent coverage will reach more employer-sponsored health plans and could provide a comparatively affordable coverage option for thousands of young adults. In 2014, with few exceptions, the national reform law will mandate that individuals purchase coverage and it will provide premium subsidies to moderate-income families. In some circumstances, adult dependent coverage may provide an attractive vehicle for complying with the mandate.

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## Notes

1. States typically enacted expansion policies in the year before implementation. Among nonexpansion states, insurers typically offer dependent coverage up to age 18 and to full-time students to age 23.
2. Data from the Medical Expenditure Panel Survey—Insurance Component indicate that in 2009, 56.1 percent of private sector workers enrolled in an employer-sponsored health plan were in a self-insured plan and that in establishments of 1,000 or more employees, 82.9 percent of are in self-insured plans. Available at http://www.meps.ahrq.gov/mepsweb/data_stats/summ_tables/insr/state/series_2/2009/tiib2b1.pdf. Last accessed July 22, 2010.

3. Buchmueller and DiNardo (2002) and Monheit and Schone (2004) apply linear probability models in studies of insurance market reform. Coefficients from these models provide direct estimates of the marginal effects. Because our DD estimator is an interaction term, computing marginal effects and standard errors using logit or probit models is not straightforward (Ai and Norton 2003) and extremely computationally time intensive for the large samples used in our analysis.
4. We apply CPS population weights so that our regression coefficients reflect underlying population behavior and adjust standard errors for the CPS complex survey design following Davern et al. (2007). We define our strata to be metropolitan core-based statistical area in which the household resides and cluster standard errors at the household level.
5. We find suggestive evidence that these changes in dependent coverage were accompanied by declines in uninsured rates in the year of implementation (data not displayed): a 4.14 and 3.85 percentage point decline in uninsured rates for samples B and D ($p<.10$ for the latter). However, these results are offset by increased uninsured rates in subsequent years.

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## SUPPORTING INFORMATION

Additional supporting information may be found in the online version of this article:

Appendix SA1: Author Matrix.
Appendix SA2. Results of Linear Probability Models for Samples A & D.

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