Research Report

The Effect of the 2014 Medicaid Expansion on Insurance Coverage for Newly Eligible Childless Adults

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Preface

Following full implementation of the Affordable Care Act's (ACA's) coverage provisions on January 1, 2014, the share of U.S. adults without insurance fell substantially more in Medicaid expansion states than in nonexpansion states. These figures suggest that Medicaid expansion is succeeding at reducing uninsurance. However, existing estimates of the Medicaid expansion's effects do not specifically examine take-up of Medicaid among adults who became newly eligible following the Medicaid expansion. This report used data from the 2009–2014 National Health Interview Survey with restricted-use state geocodes to measure the effect of state Medicaid expansion decisions on insurance coverage and the source of coverage among childless adults who became newly eligible for Medicaid in 2014. This report uses a differences-in-differences approach to compare newly eligible adults to similar adults in nonexpansion states who were not eligible for subsidized coverage through the Health Insurance Marketplace in 2014.

This study addressed the following research questions:

- How did Medicaid expansion affect insurance status for low-income childless adults who became newly eligible?
- Of the newly eligible adults gaining Medicaid coverage because of the expansion, how many would otherwise have been uninsured, and how many would otherwise have been covered by private insurance?
- Which subgroups of the newly eligible population were more or less likely to take up Medicaid coverage in 2014?

This report may be of interest to state and federal health policymakers, as well as other analysts evaluating the ACA and the effects of Medicaid eligibility on insurance coverage.

This research was funded by the Robert Wood Johnson Foundation (RWJF) through its State Health Access Reform Evaluation (SHARE) initiative. SHARE is an RWJF national program that supports rigorous research on health reform issues at a state level, with a focus on state-level implementation of the ACA and other efforts designed to increase coverage and access. SHARE operates out of the State Health Access Data Assistance Center, an RWJF-funded research center in the Division of Health Policy and Management, School of Public Health, University of Minnesota. More information about SHARE is available at www.shadac.org/share. Support for activities related to this project prior to the award from RWJF was provided by the Bing Center for Health Economics at the RAND Corporation.

The research was conducted in RAND Health, a division of the RAND Corporation. A profile of RAND Health, abstracts of its publications, and ordering information can be found at www.rand.org/health.

Abstract

The authors used the National Health Interview Survey (NHIS) to estimate how the Affordable Care Act Medicaid expansion affected health insurance enrollment, by source of coverage, among childless adults who became newly eligible for Medicaid in 2014. The NHIS data allowed the authors to report changes in enrollment by source of coverage and to conduct subgroup analyses of Medicaid take-up by gender, age, and other characteristics. Newly eligible childless adults in expansion states were 8.9 percentage points more likely to be insured in 2014 relative to similar adults in nonexpansion states, reflecting gains in Medicaid with little to no offsetting decrease in private coverage. Subgroup patterns of take-up among the newly eligible differed from findings previously reported for the wider low-income population, many of whom were previously eligible. Because these estimates isolate the behavior of newly eligible adults, these findings may be useful for anticipating take-up if nonexpansion states with limited Medicaid eligibility under current law choose to expand in the future. Similarly, because the control group excludes adults who became eligible for subsidized insurance coverage through the Health Insurance Marketplace, these findings provide insight into the effects of Medicaid expansion relative to a counterfactual involving neither Medicaid expansion nor Marketplace subsidies.

Acknowledgments

We wish to thank Xiaoyu Nie for excellent research assistance; April Grady at the Medicaid and CHIP Payment and Access Commission for providing feedback on our coding of state Medicaid rules; Daniel Levy, Robin Cohen, and Sandra Decker at the National Center for Health Statistics (NCHS) for answering questions about the National Health Interview Survey (NHIS); Jing Tian at NCHS for programming and output review; John Sullivan and Gary Gates at the University of California, Los Angeles (UCLA) Census Research Data Center (RDC) for providing a venue in which we could work with the restricted-use NHIS; Brian Littenberg at the University of Southern California Census RDC for allowing us to continue work during a temporary closure at the UCLA RDC; three anonymous journal referees for reviews of an earlier version of this study; and Katie Carman, Rosalie Pacula, Chapin White, Sandra Berry, Nicole Maestas, and Peter Huckfeldt for providing helpful suggestions and feedback.

The analysis reported in this paper was conducted at the UCLA Census RDC using restricted-use data provided by NCHS. The findings and conclusions in this paper are those of the authors and do not necessarily represent the views of the RDC, NCHS, or the Centers for Disease Control and Prevention.

Contents

Abbreviations

ACA	Affordable Care Act
AFDC	Aid to Families with Dependent Children
CHIP	Children's Health Insurance Program
CMS	Centers for Medicare & Medicaid Services
CPS ASEC	Current Population Survey Annual Social and Economic Supplement
FPL	federal poverty level
GAMC	General Assistance Medical Care
IPUMS-CPS	Integrated Public Use Microdata Series, Current Population Survey
LPM	linear probability model
MAGI	modified adjusted gross income
NCHS	National Center for Health Statistics
NHIS	National Health Interview Survey
OLS	ordinary least squares
RDC	Research Data Center
RWJF	Robert Wood Johnson Foundation
SCHIP	State Children's Health Insurance Program
SHARE	State Health Access Reform Evaluation
SSDI	Social Security Disability Income
SSI	Supplemental Security Income
UCLA	University of California, Los Angeles

Introduction

Following full implementation of the Affordable Care Act's (ACA's) coverage provisions on January 1, 2014, the share of U.S. adults without insurance fell substantially more in Medicaid expansion states than in nonexpansion states (Cohen and Martinez, 2015b). Between 2013 and 2014, Medicaid enrollment increased by 10.75 million nationwide (Centers for Medicare & Medicaid Services [CMS], 2015). These figures suggest that Medicaid expansion is succeeding at reducing uninsurance, but more detailed analysis is needed to distinguish the impact of the Medicaid expansion from the impact of other policies, like the establishment of the Health Insurance Marketplace. Further, the extent to which Medicaid expansion reached the newly eligible population, versus increasing enrollment among people who would have been eligible under the old rules (the welcome-mat effect), remains unclear. Finally, there is limited evidence on the extent to which increases in Medicaid enrollment resulting from the ACA expansion may have been offset by reductions in private insurance coverage.

We used data from the 2009–2014 National Health Interview Survey (NHIS) with restricted-use state geocodes to measure the effect of state Medicaid expansion decisions on insurance coverage and the source of coverage among childless adults who became newly eligible for Medicaid in 2014. Because there are important pre-ACA differences in the demographic, health policy, and economic environments of these two groups of states, we used a differences-in-differences research design to distinguish changes in outcomes resulting from the Medicaid expansion from permanent differences between states and from nationwide changes associated with ACA implementation.

Our study addressed the following research questions:

- How did Medicaid expansion affect insurance status for low-income childless adults who became newly eligible?
- Of the newly eligible adults gaining Medicaid coverage because of the expansion, how many would otherwise have been uninsured, and how many would otherwise have been covered by private insurance?
- Which subgroups of the newly eligible population were more or less likely to take up Medicaid coverage in 2014?

While a growing number of studies have examined the Medicaid expansion, no currently published studies estimate the impact of Medicaid expansion on newly eligible adults or use a control group that was not directly targeted by other ACA coverage expansions. We limited our sample to nondisabled childless adults and excluded from our analysis 13 states in which these adults were categorically eligible (i.e., they had the opportunity to qualify for Medicaid if their incomes were sufficiently low) for Medicaid before 2014. We also focused on adults in poverty—those with family incomes below the federal poverty level (FPL)—because adults with income above the FPL became eligible for Marketplace subsidies in nonexpansion states in 2014. While nondisabled childless adults in poverty constitute only about 9 million of the estimated 196 million adults aged 18–64 in 2014, this population is of particular policy interest because it is a group of newly eligible adults who can readily be identified from survey data on poverty status, family structure, and sources of income.

Our narrower sample definition allows us to examine the effect of new Medicaid eligibility on coverage relative to a counterfactual scenario without Medicaid expansion or Marketplace subsidies. Currently published research addresses the impact of Medicaid expansion on the entire population covered by the new eligibility group—most of whom were already eligible for Medicaid before 2014—relative to a counterfactual scenario that includes Marketplace subsidies for those with incomes above the FPL (Sommers, Gunja, et al., 2015b; Wherry and Miller, 2016). Our analysis complements these studies by providing new evidence on patterns of take-up behavior and the effectiveness of the Medicaid expansion in reaching populations previously excluded from public insurance coverage.

Among newly eligible childless adults, we found that the 2014 Medicaid expansion led to an 8.9-percentage-point increase in the rate of insurance coverage, which was driven by take-up of Medicaid with limited crowd-out of private insurance. Take-up among the newly eligible varied significantly across age groups and racial/ethnic groups in ways that differed from previously available estimates. In addition, adults in worse health experienced larger gains in coverage, suggesting that individuals with greater health care needs were the first to enroll after expansion.

Related Literature

Most of the literature on state-level Medicaid expansions prior to the ACA suggests that coverage expansions increase Medicaid enrollment even though take-up of Medicaid coverage among the newly eligible is often low (Long, Zuckerman, and Graves, 2006; Sommers, Baicker, and Epstein, 2012; Sommers, Kenney, and Epstein, 2014). The literature is more mixed regarding the degree to which expansions lead to offsetting decreases in private coverage, a phenomenon known as *crowd-out*. Several studies of pre-ACA coverage expansions for low-income adults found limited crowd-out for adults in poverty, with crowd-out rates increasing for higher-income families (Gruber and Simon, 2008; Hamersma and Kim, 2013; Long et al., 2006). More recent studies of pre-ACA expansions yield larger estimates of crowd-out for low-income adults, although researchers studying California's early implementation of the ACA expansion found no significant changes in private coverage (Garthwaite, Gross, and Notowidigdo, 2014; Golberstein, Gonzales, and Sommers, 2015).

Researchers have now begun to measure the effects of the 2014 Medicaid expansion. Several early studies that used tracking surveys to analyze the effect of the ACA found larger increases in coverage in Medicaid expansion states, including one study that found this to be the case specifically for low-income adults (Carman, Eibner, and Paddock, 2015; Karpman and Long, 2015; Sommers, Gunja, et al., 2015b). However, these surveys may not reliably identify the source of coverage, and they have substantially lower response rates than the NHIS.

More recently, several papers have used large federal surveys to compare changes in insurance coverage between expansion and nonexpansion states. Our work is most closely related to a recent study that used the 2010–2014 NHIS (Wherry and Miller, 2016). That study found that Medicaid expansion was associated with a 7.4-percentage-point increase in overall insurance coverage for adults with incomes below 138 percent of the FPL, reflecting a 10.5-

percentage-point increase in Medicaid coverage. The gap between the increase in Medicaid coverage and the increase in overall insurance coverage may reflect a reduction in private health insurance coverage, but this effect is imprecisely estimated and was not statistically distinguishable from zero at the 95-percent confidence level. Another study using the 2013–2014 American Community Survey found that coverage gains among adults in poverty were larger in expansion states than in nonexpansion states (Courtemanche, Marton, and Yelowitz, 2016).

Findings have also begun to emerge about which demographic groups benefited most from the 2014 coverage expansions. Across all states and income groups, young adults experienced larger reductions in uninsurance between 2013 and 2014 than did older adults, and larger gains in insurance coverage have also been documented for Black and Hispanic adults relative to white adults (Buchmueller et al., 2016; Courtemanche, Marton, and Yelowitz, 2016; McMorrow et al., 2015a; McMorrow et al., 2015b).

Our study adds to the existing literature by narrowly focusing on take-up behavior among the newly eligible, whereas the studies described above included parents and childless adults who were categorically eligible for Medicaid in 2013. To be clear, outcomes for the broader population studied elsewhere are of considerable policy interest. However, it is challenging to infer from currently published studies how the Medicaid expansion affected insurance coverage relative to a baseline without other coverage expansions; it is also not straightforward to distinguish take-up among the newly eligible from enrollment among previously eligible populations, such as low-income parents or participants in federal disability programs. Estimates that address these narrower questions may be of value to researchers wishing to assess whether the Medicaid expansion has succeeded in reaching newly eligible populations. Information about the take-up behavior of newly eligible childless adults, meanwhile, may be useful to policymakers in nonexpansion states: Most of these states have very limited Medicaid eligibility for childless adults in comparison with states that adopted the Medicaid expansion in 2014 or earlier, meaning that the overall impact of Medicaid expansion in these states would reflect the experiences of the newly eligible to a greater extent than in expansion states that have historically made Medicaid available to a larger population.

Similarly, we view our subgroup analyses—which are driven by take-up behavior among the newly eligible—as answering a different question from studies that measure the ACA's impact on health insurance disparities more broadly. Existing reports of demographic patterns in insurance coverage between 2013 and 2014 either do not distinguish between expansion and nonexpansion states (Courtemanche, Marton, and Yelowitz, 2016; Cohen and Martinez, 2014) or do not stratify on income and pre-ACA eligibility in a way that allows conclusions to be drawn about take-up among the newly eligible (Cohen and Martinez, 2015b; McMorrow et al., 2015a; McMorrow et al., 2015b). This is true even of the estimates reported by Buchmueller et al. (2016), which captured changes in the uninsurance rate by race and ethnicity specifically for adults with family income below 138 percent of the FPL. The estimates in that study pooled parents and childless adults together and included Medicaid expansion states that had already expanded Medicaid eligibility to this income range prior to the ACA. The analysis presented here complements existing estimates of the ACA's impact on health insurance disparities by isolating one mechanism—differences in Medicaid take-up and crowd-out for the newly eligible—that contributes to the overall impact of the ACA.

Data and Methods

We used individual-level data from the NHIS, a household survey designed by the National Center for Health Statistics (NCHS) and collected by the Census Bureau. Our research design used repeated cross-sectional data from 2009 through 2014, encompassing years before and after the 2014 Medicaid expansion. Our analysis focused on the average effect of the 2014 Medicaid expansion in the 14 expansion states in which childless nondisabled adults were categorically ineligible for Medicaid in 2013.

In order to obtain a sample of adults who became newly eligible for Medicaid in 2014, we restricted our analysis to childless adults in poverty who were not beneficiaries of Supplemental Security Income (SSI) or Social Security Disability Income (SSDI). We focused on childless adults because low-income parents were categorically eligible for Medicaid in all states prior to the ACA. We excluded disabled individuals because SSI recipients are categorically eligible for Medicaid in most states, while SSDI recipients frequently qualify through other pathways.

In addition to limiting the sample to nondisabled childless adults in poverty, we excluded 13 states with categorical eligibility for nondisabled childless adults before 2014; our methodology for classifying states is presented in the appendix. The resulting sample consists of adults who were not eligible to enroll in comprehensive Medicaid coverage prior to 2014.

We classified insurance coverage into three categories: private, Medicaid, and non-Medicaid public insurance (such as Medicare and TRICARE). These categories provided an exhaustive classification of the types of insurance reported in the NHIS and allowed us to examine whether adults gaining Medicaid transitioned from other coverage.

We used a regression-adjusted differences-in-differences approach to model the effect of Medicaid eligibility on our sample, comparing changes in insurance over time between expansion and nonexpansion states. This approach controlled for permanent differences between states as well as nationwide changes over time that affect all states, including ACA reforms other than Medicaid expansion and the Marketplace subsidies. We also controlled for gender, age, marital status, race, educational attainment, and employment status. All estimates were weighted to represent the civilian noninstitutionalized population.

The basic differences-in-differences framework assumes that trends in outcomes would have been identical in expansion and nonexpansion states. We relaxed this assumption by controlling for a separate linear time trend in expansion states in addition to nationwide time effects that varied from quarter to quarter. Estimates without differential trends were qualitatively similar to our main specification, but, as we discuss in the next section, the magnitudes of the estimated effects were sensitive to assumptions about differential trends.

Limitations

Although we attempted to exclude adults who were categorically eligible for Medicaid prior to the ACA expansion, a small percentage of our sample (7 to 13 percent) reported having

Medicaid coverage prior to 2014. We cannot determine why these individuals reported having Medicaid coverage: Some might have been eligible through special pathways (e.g., pregnancy, breast or cervical cancer), some might have been enrolled in a limited benefits package, some might have been covered through programs that were closed to new enrollees, some might have enrolled in Medicaid as parents or caretakers prior to a change in family status, and some might have been misreporting their insurance type. Similarly, we did not observe immigration status, and so our sample may include some individuals who were ineligible for Medicaid because they were not lawful permanent residents. We wish to interpret our estimates as the effect on insurance coverage of becoming newly eligible for the comprehensive Medicaid benefits package available under the ACA expansion, and so nonzero rates of Medicaid coverage and unobserved immigration status may be problematic for this interpretation of our estimates.

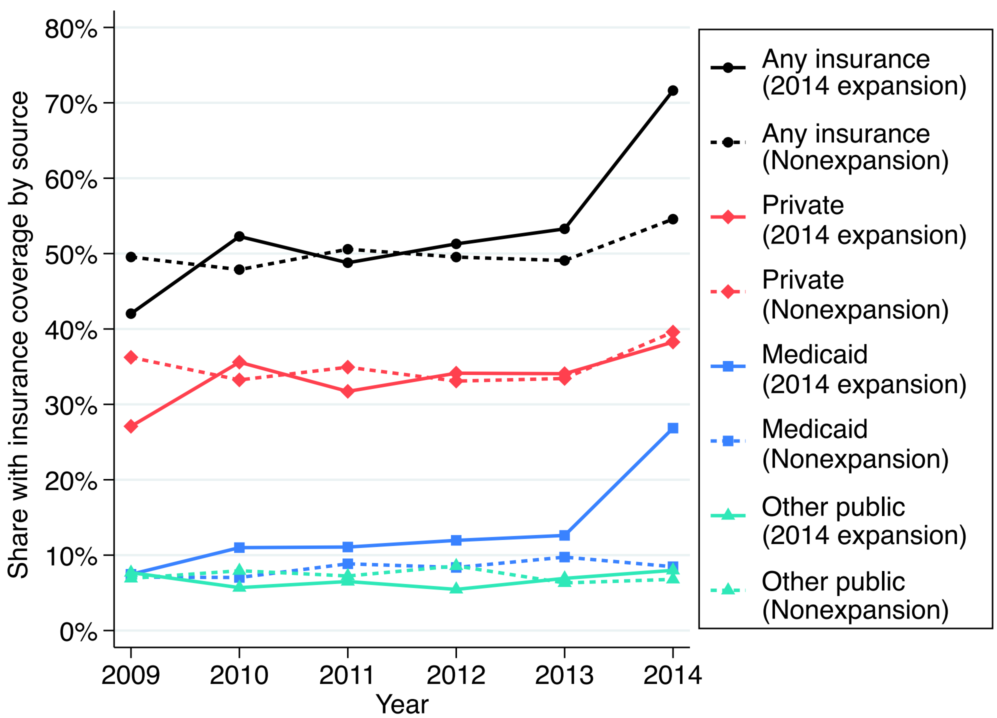
Despite the small percentage of seemingly ineligible childless adults reporting Medicaid coverage, the NHIS, like all other household surveys, undercounts participation in Medicaid relative to administrative data (Call, Davern, et al., 2013). The NHIS has several features—including a point-in-time coverage question and the use of state-specific plan names—that should mitigate limitations identified in other surveys (Cantor et al., 2007; Klerman et al., 2009). Previous audit studies suggested that survey reporting error may have a limited impact on estimates of the uninsurance rate in our setting (Call, Davidson, et al., 2008). However, we were also interested in the source of health insurance, and reporting error may have a larger impact on estimates of take-up and crowd-out. To the extent that the NHIS undercounts Medicaid coverage and overcounts private insurance, we may have tended to underestimate crowd-out. A more serious concern would be if Medicaid expansion leads to a reduction in Medicaid underreporting that is larger in expansion states than in nonexpansion states, which might bias our estimates of Medicaid take-up upward while biasing our estimates of crowd-out downward. We do not have any evidence that reporting styles have been affected by Medicaid expansion, but we cannot rule out this possibility, either.

A further limitation of our study is that our main model assumed that the Medicaid expansion occurred abruptly at the beginning of 2014 (or later in Michigan and New Hampshire). State and private outreach efforts, as well as public awareness of the ACA, could have changed insurance coverage or reporting behavior prior to the official implementation date. For example, expansion state Medicaid agencies may have reduced the frequency or stringency of recertification of eligibility immediately before expansion. Even though we defined our analysis sample to exclude states in which childless adults were eligible for Medicaid before 2014, we cannot rule out a priori the possibility that such preimplementation activities could affect our estimates. Any such effects may have caused us to underestimate the impact of the expansion, particularly in models with differential trends. To guard against bias caused by such preimplementation effects, we ran alternative models that excluded 2013 data, using 2012 and earlier years to define the pre-ACA baseline. We also estimated several models that used alternative assumptions about the presence and functional form of differential trends in coverage in expansion and nonexpansion states. These sensitivity analyses are discussed in the next section.

Results

Figure 1 shows unadjusted time trends in insurance coverage by state expansion status, along with trends in coverage for the three subtypes of insurance considered in this analysis. In 2014 expansion states, the fraction of nondisabled childless adults in poverty covered by health insurance increased by 18.3 percentage points between 2013 and 2014. There was a smaller increase in coverage of 5.5 percentage points in nonexpansion states. These changes in overall insurance coverage closely mirror the pattern observed for Medicaid coverage, which increased by 14.2 percentage points in expansion states with no meaningful change in nonexpansion states. Without adjusting for covariates or differential pre-ACA trends in coverage, these data imply that the Medicaid expansion increased the probability of any insurance coverage by 12.6 percentage points and increased the probability of Medicaid coverage by 14.9 percentage points. These unadjusted differences-in-differences are highly statistically significant ($p < 0.001$).

Figure 1. Insurance Coverage by Source for Nondisabled Childless Adults in Poverty, 2014 Expansion States Versus Nonexpansion States, 2009–2014



SOURCE: Authors' calculations, 2009–2014 NHIS Person File.
NOTES: This figure reports the probability of insurance coverage by source, year, and state Medicaid expansion status for nondisabled childless adults in families with income less than or equal to 100 percent of the FPL. Estimates use final annual person weights. 2014 expansion states are defined as states that began implementation of the ACA Medicaid expansion during 2014 and where childless adults were not categorically eligible for comprehensive Medicaid benefits in 2013 and earlier years. These 14 states are Arizona, Arkansas, Illinois, Kentucky, Maryland, Michigan, Nevada, New Hampshire, New Mexico, North Dakota, Ohio, Oregon, Rhode Island, and West Virginia.

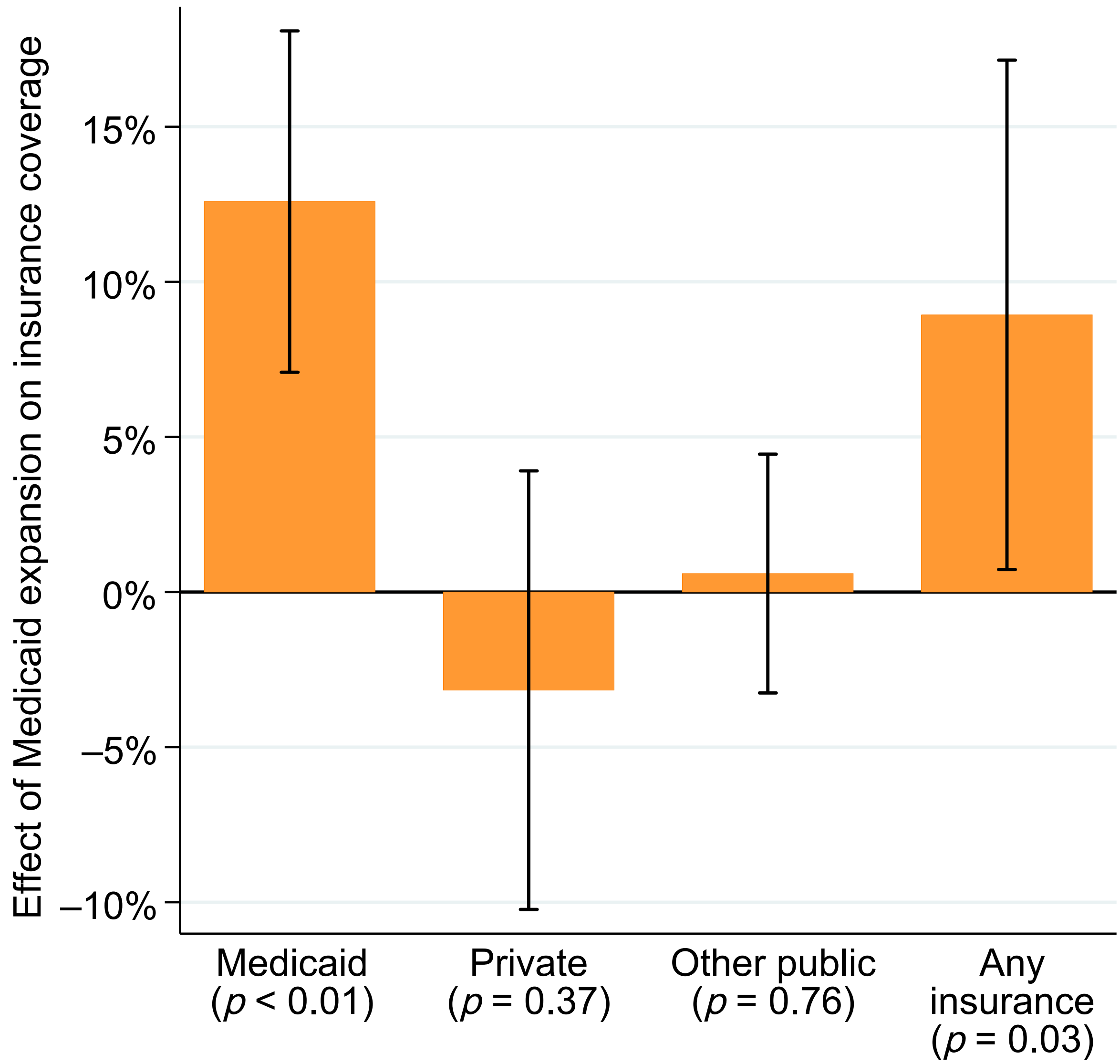
Nonexpansion states are defined as states that had not moved forward with the Medicaid expansion as of the fourth quarter of 2014. These 24 states are Alabama, Alaska, Florida, Georgia, Idaho, Indiana, Kansas, Louisiana, Maine, Mississippi, Missouri, Montana, Nebraska, North Carolina, Oklahoma, Pennsylvania, South Carolina, South Dakota, Tennessee, Texas, Utah, Virginia, Wisconsin, and Wyoming.

Turning to the remaining forms of insurance, Figure 1 shows that rates of private and non-Medicaid public coverage were similar in expansion and nonexpansion states in all years, with the exception of a lower private coverage rate in 2009 for expansion states. Both expansion and nonexpansion states show an increase in private coverage between 2013 and 2014. The estimated increase is slightly larger in nonexpansion states, but unadjusted differences-in-differences estimates for an effect on private coverage were not significant ($p = 0.97$). Non-Medicaid public coverage did not change meaningfully in either group of states over the study period.

While Figure 1 clearly shows that differential increases in Medicaid coverage and overall insurance coverage were associated with Medicaid expansion, we also observed a slight increase in coverage between 2009 and 2013 in the expansion states and a slight decrease in nonexpansion states. Statistical tests reported in the appendix indicate that differential trends in insurance coverage and Medicaid were just barely insignificant at the 10-percent level. To ensure robustness to differential pre-ACA trends, we included a linear trend specific to the expansion states in our main regression model.

Figure 2 presents regression-adjusted estimates of the impact of Medicaid expansion on coverage for low-income childless adults. Overall, insurance coverage increased by 8.9 percentage points ($p = 0.03$), and Medicaid coverage increased by 12.6 percentage points ($p < 0.001$). Private coverage declined by a statistically insignificant 3.2 percentage points ($p = 0.37$). There was no effect on non-Medicaid public insurance ($p = 0.76$).

Figure 2. Effects of 2014 Medicaid Expansion on Insurance Coverage for Newly Eligible Childless Adults



SOURCE: Authors' calculations, 2009–2014 NHIS.
NOTES: This figure reports regression-adjusted differences-in-differences effects of ACA Medicaid expansion on insurance coverage by type. The effect is estimated as a coefficient on a dummy variable equal to 1 in expansion states after the implementation date and 0 otherwise. Effects are reported in percentage points.
Effects were estimated using linear regressions controlling for gender, age, marital status, race, educational attainment, employment status, time (year-quarter) fixed effects, state fixed effects, and a linear time trend specific to expansion states.
2014 expansion states are defined as states that began implementation of the ACA Medicaid expansion during 2014 and where childless adults were not categorically eligible for comprehensive Medicaid benefits in 2013 and earlier years. These 14 states are Arizona, Arkansas, Illinois, Kentucky, Maryland, Michigan, Nevada, New Hampshire, New Mexico, North Dakota, Ohio, Oregon, Rhode Island, and West Virginia. The treatment dummy for Michigan switches from 0 to 1 in the second quarter of 2014. The treatment dummy for New Hampshire switches from 0 to 1 in the fourth quarter of 2014. Data from New Hampshire in the third quarter of 2014 were dropped as a wash-out period for New Hampshire expansion implementation.
Nonexpansion states are defined as states that had not moved forward with the Medicaid expansion as of the fourth quarter of 2014. These 24 states are Alabama, Alaska, Florida, Georgia, Idaho, Indiana, Kansas, Louisiana, Maine,

Mississippi, Missouri, Montana, Nebraska, North Carolina, Oklahoma, Pennsylvania, South Carolina, South Dakota, Tennessee, Texas, Utah, Virginia, Wisconsin, and Wyoming.
Error bars report 95-percent confidence intervals based on *t*-distribution with 37 degrees of freedom and standard errors clustered on state.

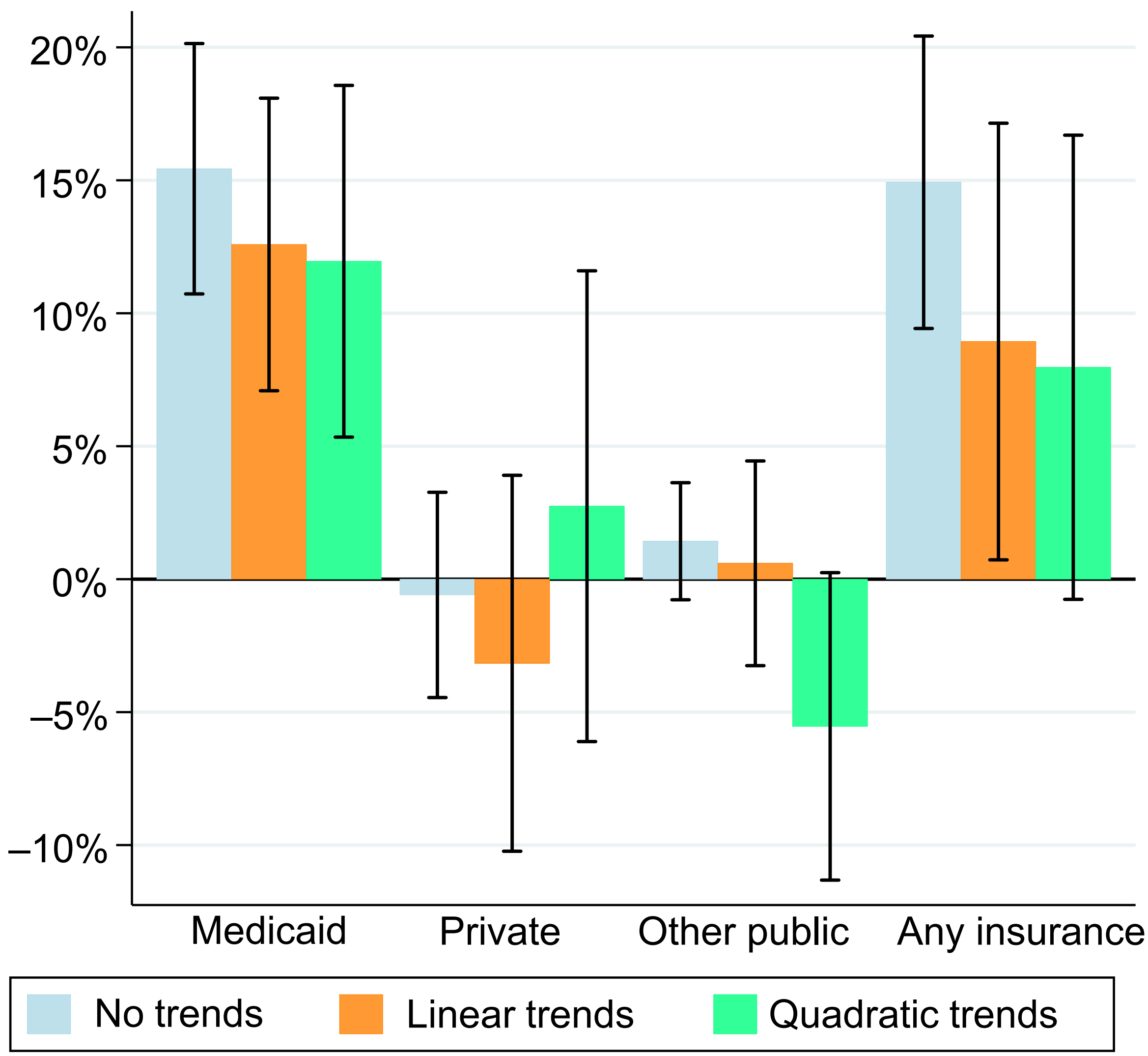
Although our estimate of the Medicaid expansion's impact on private insurance is statistically insignificant, it is negative. The point estimate is large enough to be substantively meaningful, suggesting that 25 percent of newly eligible childless adults gaining Medicaid coverage would have had private coverage in the absence of the Medicaid expansion. However, this estimate is imprecise: We cannot rule out decreases in private insurance coverage as large as 10.2 percentage points or increases as large as 3.9 percentage points at the 95-percent confidence level.

Sensitivity Analysis

To evaluate more fully whether newly eligible adults gaining Medicaid coverage in 2014 would have been privately insured in the absence of the Medicaid expansion, we estimated several additional models that rely on different assumptions about the presence and functional form of differential trends in coverage in expansion and nonexpansion states. Figure 3 presents estimates from these models side by side with our main specification. These alternative models did not provide any evidence that Medicaid expansion crowded out private coverage among newly eligible adults; the specification that includes linear differential trends yielded the most negative impacts on private coverage of any of the models we estimated.

Figure 3 indicates that the estimated change in private coverage is fairly sensitive to assumptions about differential trends in coverage between expansion and nonexpansion states. Evidence of decreases in private coverage for newly eligible adults appears to be limited to an imprecise point estimate in one specification. Estimated increases in Medicaid coverage were more robust to different assumptions about differential trends in expansion and nonexpansion states. While the impact of Medicaid expansion on overall coverage became slightly smaller and less significant when we controlled for quadratic trends, the point estimate was very close to the estimate in our main specification. These models and results are presented in greater detail in the appendix.

Figure 3. Estimates of Medicaid Expansion Impacts Under Alternative Modeling Assumptions



SOURCE: Authors' calculations, 2009–2014 NHIS.
NOTES: This figure reports regression-adjusted differences-in-differences effects of ACA Medicaid expansion on insurance coverage by type under alternative specifications for differential time trends in expansion and nonexpansion states. Effects were estimated as a coefficient on a dummy variable equal to 1 in expansion states after the implementation date and 0 otherwise. Effects are reported in percentage points.
Effects were estimated using linear regressions controlling for gender, age, marital status, race, educational attainment, employment status, time (year-quarter) fixed effects, state fixed effects, and a linear time trend specific to expansion states.
2014 expansion states are defined as states that began implementation of the ACA Medicaid expansion during 2014 and where childless adults were not categorically eligible for comprehensive Medicaid benefits in 2013 and earlier years. These 14 states are Arizona, Arkansas, Illinois, Kentucky, Maryland, Michigan, Nevada, New Hampshire, New Mexico, North Dakota, Ohio, Oregon, Rhode Island, and West Virginia. The treatment dummy for Michigan switches from 0 to 1 in the second quarter of 2014. The treatment dummy for New Hampshire switches from 0 to 1 in the fourth quarter of 2014. Data from New Hampshire in the third quarter of 2014 were dropped as a wash-out period for New Hampshire expansion implementation.
Nonexpansion states are defined as states that had not moved forward with the Medicaid expansion as of the fourth quarter of 2014. These 24 states are Alabama, Alaska, Florida, Georgia, Idaho, Indiana, Kansas, Louisiana, Maine, Mississippi, Missouri, Montana, Nebraska, North Carolina, Oklahoma, Pennsylvania, South Carolina, South Dakota,

Tennessee, Texas, Utah, Virginia, Wisconsin, and Wyoming.
Error bars report 95-percent confidence intervals based on *t*-distribution with 37 degrees of freedom and standard errors clustered on state.

Figure 3 also highlights a methodological point that is relevant to other studies that exploit the policy variation created by state Medicaid expansion decisions: The treatment of preexisting trends can have a substantial effect on the conclusions that analysts draw about the impact of Medicaid expansion. The assumption that outcomes in expansion and nonexpansion states would have moved in parallel in the absence of Medicaid expansion should be evaluated carefully.

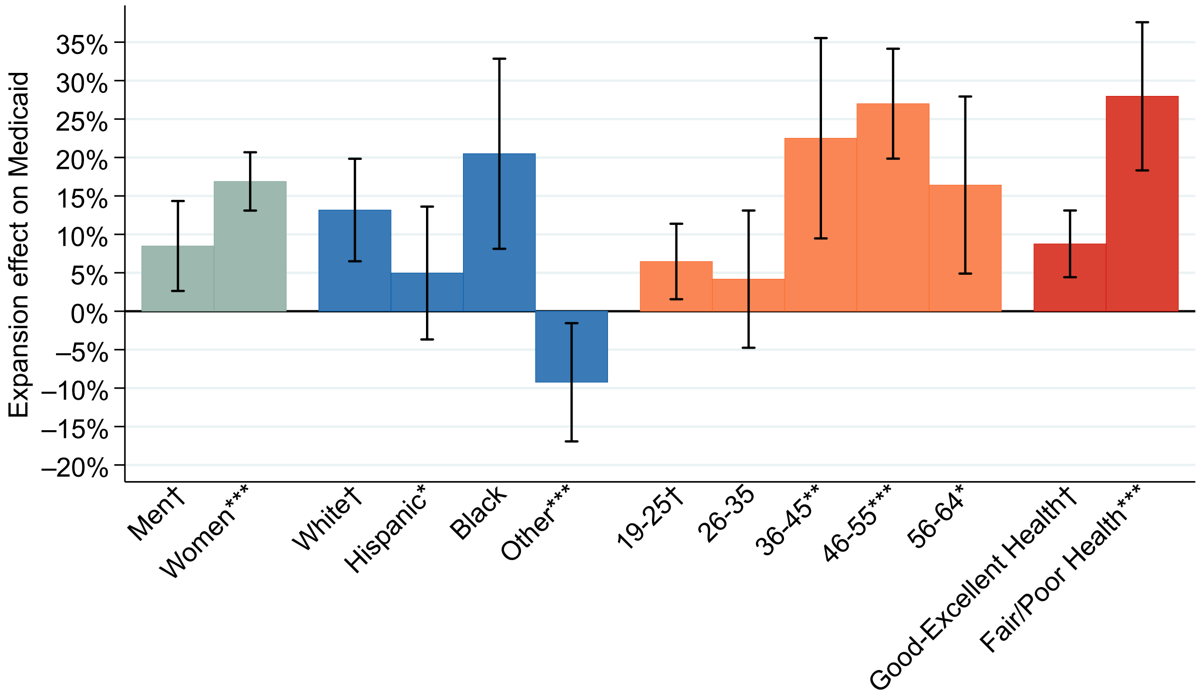
We estimated several further alternative specifications of our regression model to assess the robustness of our results to the preexpansion time period used as a baseline and our choice of estimation sample. Because of the concerns raised above about preimplementation activities, we ran several models that excluded 2013 data, using 2012 and earlier years as the pre-ACA baseline. These results, available in the appendix, show that our estimates are not sensitive to the inclusion or exclusion of data from the year preceding the expansion's implementation.

Readers may also be concerned that it is inappropriate to include employment status as a control variable because labor supply may respond to Medicaid eligibility (Garthwaite et al., 2014). In the appendix, we report estimates for models that did not control for employment status. These estimates are nearly identical to our main specification. We focused on a narrow group of expansion states to isolate the newly eligible. Estimates for a sample that includes four states with eligibility for some childless adults in poverty in 2014 are very similar to our main results (see the appendix).

Demographic Differences in Medicaid Take-Up

To better understand which individual characteristics were associated with take-up of Medicaid, we analyzed changes in Medicaid coverage by gender, race/ethnicity, age, and health status (Figure 4). Regression coefficients and estimates for other types of coverage can be found in the appendix.

Figure 4. Effects of 2014 Medicaid Expansion on Medicaid Coverage for Newly Eligible Childless Adults by Gender, Race, Age, and Health Status



SOURCE: Authors' calculations, 2009–2014 NHIS.
NOTES: This figure reports regression-adjusted differences-in-differences effects of ACA Medicaid expansion on Medicaid coverage for subgroups. Base effects are estimated as a coefficient on a dummy variable equal to 1 in expansion states after the implementation date and 0 otherwise. Interaction effects are estimated as a coefficient on interaction between an expansion dummy variable and a dummy variable for subgroup membership. Effects are reported in percentage points.
Effects are estimated using linear regressions controlling for gender, age, marital status, race, educational attainment, employment status, time (year-quarter) fixed effects, state fixed effects, and a linear time trend specific to expansion states.
† Indicates the base category in the regression model; other bars report the sum of base and interaction effects.
P-values for significance of pairwise difference from base category effect are indicated as follows: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.
Error bars report 95-percent confidence intervals based on *t*-distribution with 37 degrees of freedom and standard errors clustered on state. Confidence intervals for interaction effects treat base coefficients as known.

While Medicaid coverage increased for nearly all of the subpopulations examined, there were meaningful differences between demographic groups in the size of the increase. Men were 8 percentage points less likely than women to gain Medicaid coverage. The effect of the expansion did not differ significantly between Black and white non-Hispanic adults. Medicaid take-up among Hispanic adults was marginally significantly lower than among white non-Hispanic adults ($p = 0.06$), though these groups experienced identical gains in overall insurance coverage because Hispanic adults experienced larger increases in private coverage. Adults in the non-Hispanic "other" racial category (primarily Asian Americans) were significantly less likely to gain Medicaid coverage than other racial groups, and actually experienced reductions in Medicaid coverage relative to similar adults in nonexpansion states.

Our subgroup findings differ from corresponding nationwide changes in coverage between 2013 and 2014. For instance, the NHIS showed similar reductions in uninsurance for men (4.2 percentage points) and women (4 percentage points) (Cohen and Martinez, 2014, 2015a). Even though we were not able to condition on previous insurance status in the NHIS, the gender difference we observed in Medicaid take-up is similar to early estimates of insurance changes among previously uninsured adults (Vistnes and Cohen, 2015).

More striking differences were observed among subgroups defined by age and health status. Adults over age 35 were 10 to 21 percentage points more likely to gain Medicaid coverage than younger adults, and people in poor or fair health were 19 percentage points more likely to gain Medicaid coverage than those in good, very good, or excellent health.

Subgroup effects on any insurance coverage are similar to the results shown for Medicaid, except that the size of the Medicaid expansion effect on any insurance coverage for Hispanic adults is very close to the size of the effect for non-Hispanic white adults. We found no evidence of a statistically significant reduction in private insurance coverage for any of the groups considered. These estimates are reported in the appendix.

To provide additional insight into the demographic differences in take-up behavior implied by these estimates, Table 1 presents the subgroup effects underlying Figure 4 alongside the 2013 uninsurance rate for each subgroup in the expansion states included in our sample.

Table 1. Medicaid Coverage Gains, Preexpansion Uninsurance Rates, and Implied Medicaid Take-Up Rates by Demographic Group for Nondisabled Childless Adults in Poverty

Demographic Group	Expansion Effect on Medicaid Coverage	2013 Uninsurance Rate in Expansion States	Implied Medicaid Take-Up Rate
Total			
Nondisabled childless adults in poverty	13%	42%	30%
Gender			
Men (base)	8%	36%	23%
Women***	17%	33%	52%
Race/ethnicity			
White non-Hispanic (base)	13%	39%	34%
Hispanic*	5%	35%	14%
Black non-Hispanic	20%	33%	62%
Other non-Hispanic***	–9%	25%	–37%
Age			
19–25 (base)	6%	30%	22%
26–35	4%	47%	9%
36–45**	23%	48%	47%
46–55***	27%	36%	75%
55–64*	16%	25%	66%
Health status			
Good or better health (base)	9%	39%	23%
Fair or poor health***	28%	26%	108%

SOURCE: Authors' calculations, 2009–2014 NHIS.
The expansion effect on Medicaid coverage is the adjusted differences-in-differences point estimate presented in Figure 2 (for all nondisabled childless adults in poverty) or Figure 3 (for subgroups).
The 2013 uninsurance rate was calculated for the estimation sample of nondisabled childless adults in poverty for the expansion states included in our analysis.
The implied take-up rate is the increase in Medicaid coverage caused by Medicaid expansion divided by the 2013 uninsurance rate.
P-values for significance of pairwise difference from base category effect are indicated as follows: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.
See the appendix for details of estimation and standard errors on Medicaid coverage effects.

The NHIS does not capture information on the past insurance status of most currently insured respondents, making it infeasible to calculate a take-up rate among the population of previously uninsured individuals. However, we can give readers a sense of how the change in Medicaid coverage compares to the size of the uninsured population by scaling our estimates by the 2013 uninsurance rate in expansion states for each subgroup we examine. We labeled this ratio the "implied take-up rate" in order to distinguish it from the take-up rate among the previously insured (which we cannot estimate). If only individuals who were uninsured in 2013 enrolled in Medicaid following the expansion and there was no churn in insurance status for reasons unrelated to the ACA, this ratio would yield the take-up rate among previously uninsured individuals. However, we caution that insurance transitions are widespread even in the absence of policy changes, and that some reductions in private coverage may have resulted from Medicaid expansion.

The implied take-up rate for the overall population of nondisabled childless adults in poverty suggests that the majority of uninsured individuals in this population who gained Medicaid eligibility in 2014 did not enroll; dividing the 12.9-percentage-point increase in Medicaid coverage by the 42-percent uninsurance rate for this group in 2013 yields an implied take-up rate of 30 percent.

As indicated by Figure 4, implied take-up varied widely across subgroups. While 2013 uninsurance rates were roughly similar across genders and racial/ethnic groups, older adults and those in fair or poor health had sharply higher implied take-up rates than younger and healthier adults because the former groups had both lower rates of uninsurance and larger Medicaid coverage gains. Thus, while the overall implied take-up rate was 30 percent, the implied take-up rates for adults 35 and older ranged from 47 percent to 75 percent. For adults in fair or poor health, meanwhile, the increase in Medicaid coverage was slightly larger than the 2013 uninsurance rate, leading to an implied take-up rate just above 100 percent.

Discussion

We estimate that Medicaid expansion made low-income childless adults in Medicaid expansion states 8.9 percentage points more likely to be insured than they would have been without the expansion. Virtually all of this increase came from enrollment in Medicaid, with no significant evidence of offsetting reductions in private coverage. Our estimates for overall take-up and crowd-out are qualitatively similar to the findings of other differences-in-differences studies of the 2014 Medicaid expansion (see the appendix for further discussion). Our findings confirm that Medicaid take-up by the newly eligible contributed meaningfully to the increases in coverage observed in expansion states.

Assuming that our sample excludes adults previously eligible for comprehensive Medicaid benefits, our results can be interpreted as the effect of becoming newly Medicaid eligible on the probability of coverage among nondisabled childless adults in poverty. Most uninsured adults in our sample did not gain coverage in the first year of the expansion, however: Our estimated 12.6-percentage-point increase in Medicaid coverage represents only 30 percent of the 2013 uninsurance rate in this population (see the appendix). While low, this estimate is consistent with prior estimates of take-up among newly eligible adults (Busch and Duchovny, 2005).

We note several possible explanations for low take-up in the first year of the expansion. This relatively low take-up rate may result, in part, from the fact that there is no open enrollment period for Medicaid, so eligible individuals can delay enrollment until they seek care. The individual mandate, which could incentivize take-up among higher-income Medicaid-eligible individuals, is not applicable to most adults in poverty: Individuals with income below the federal tax income filing threshold are exempt from the mandate, and very few adults in poverty are required to file tax returns. In addition, because we did not observe immigration status, it is possible that some people whom we classify as newly eligible are in fact ineligible because of Medicaid's residency and citizenship requirements. Administrative reports from CMS suggest that Medicaid take-up continued to increase in 2015 and 2016 (CMS, 2016).

Our finding that newly eligible whites and Blacks experienced the biggest increase in Medicaid coverage is also noteworthy. Between 2013 and 2014, the nationwide uninsurance rate decreased more for Hispanics (6.9 percentage points) and non-Hispanic Blacks (4.3 percentage points) than for non-Hispanic whites (2.9 percentage points) (Cohen and Martinez, 2014, 2015a). Our findings suggest that larger gains in insurance coverage for some minority groups observed among the overall adult population were not driven by differential take-up of Medicaid among newly eligible adults. Estimates from the American Community Survey reported in a recent study that simultaneously stratifies on race, income, and state expansion status also suggest that differences in Medicaid take-up across racial and ethnic groups did not closely resemble changes in the uninsurance rate for individuals of all income levels in all states (Buchmueller et al., 2016).

Similarly, the age gradient we observed in reductions in uninsurance differs from the nationwide pattern found in the early-release NHIS data, which showed that the uninsurance rate fell most for adults aged 18–24 (6.1 percentage points), followed by adults aged 25–34 (4.4 percentage points) (Cohen and Martinez, 2014, 2015a). While the nationwide population of adults aged 45–64 had a small reduction in uninsurance between 2013 and 2014, we found that childless nondisabled adults in poverty in this age range experienced the *largest* coverage gains due to the Medicaid expansion. While we reiterate that the implied take-up rates reported above are a very rough calculation that should not be interpreted as the probability that an uninsured adult gained Medicaid coverage, Table 1 indicates that take-up in the first year of the Medicaid expansion was higher among some subgroups—older adults and those in fair or poor health—who were more likely to be insured prior to Medicaid expansion. We note that our subgroup estimates do not contradict previously reported subgroup findings for the nationwide population, as our population of interest is a small fraction of the nationwide adult population and our sample excludes California and several other large states with high pre-ACA Medicaid eligibility.

We also found that newly eligible adults in fair or poor health were more likely to gain Medicaid coverage than healthier adults. We caution that, because self-rated health is evaluated at the time of the survey, these results could potentially be biased due to reverse causation from Medicaid coverage to self-rated health status. In a study with a similar differences-in-differences research design to ours, Simon, Soni, and Cawley (2016) found that the 2014 Medicaid expansion led to a small but significant improvement in self-rated health for childless adults in poverty. If Medicaid coverage improves self-rated health status, our estimated interaction effect between *poor* health and Medicaid coverage will be biased downward (away from our finding that poor health predicted Medicaid take-up). We note that the short-term effect of coverage on self-rated health is theoretically ambiguous because increased access to care could make adults newly aware of health problems without yielding immediate improvements in health. The analysis of the 2014 NHIS by Wherry and Miller does not indicate that Medicaid expansion was associated with any changes in self-rated health status for the overall low-income adult population (Wherry and Miller, 2016). While we cannot rule out reverse causation, there is no evidence to date that the Medicaid expansion harmed self-rated health status among newly eligible adults, which is the relationship that would be needed to generate bias in the direction of our estimates.

When juxtaposed with our finding that older adults had higher take-up of Medicaid, our finding that worse self-rated health was associated with Medicaid take-up adds to the evidence that individuals with greater health care needs were the first to enroll in coverage under the 2014 expansion. This pattern is consistent with pre-ACA research on Medicaid take-up, and it mirrors a recent study showing that early enrollees in Marketplace policies were older and more likely to use medication than later enrollees (Donohue et al., 2015; Kenney et al., 2012).

An alternative explanation is that, unlike those with private insurance, Medicaid-eligible individuals can delay enrollment until they need to interact with the health care system. This phenomenon—which has been termed *conditional coverage*—is one mechanism that could lead to a mechanical association between health care utilization and Medicaid take-up among the newly eligible population. The importance of conditional coverage is likely to be most pronounced in the period immediately following Medicaid expansion, since a higher proportion of newly eligible individuals is likely to have encounters with health care providers as time passes. While our analysis does not distinguish between conditional coverage and other mechanisms that might generate higher take-up among those in fair or poor health, the contribution of conditional coverage to observed patterns of Medicaid take-up under the ACA is a question that may warrant further attention from researchers and policymakers. For instance, the policy implications of early data on per-enrollee Medicaid spending may depend on assumptions about whether conditionally covered individuals who do not seek care in a given time frame should be included in or excluded from the risk pool. Analysis of data from 2015 and subsequent years will help determine the extent to which the association between poor health and Medicaid coverage documented here is a short-run or a long-run phenomenon.

Conclusion

Our results corroborate findings from other data sources that insurance coverage increases were larger in states that expanded Medicaid. Our estimates add to the existing literature because our treatment and control groups were more narrowly defined to exclude those previously eligible for Medicaid and those directly affected by other coverage expansions. By focusing on the group most likely to gain Medicaid eligibility (nondisabled low-income childless adults) and by focusing on states that implemented the Medicaid expansion in 2014, this study clarifies the important role played by increased Medicaid coverage in previously reported coverage gains. We found very little evidence of differential changes in private coverage between expansion and nonexpansion states. Although our best estimate of the increase in overall insurance coverage was somewhat smaller than the increase in Medicaid coverage, our estimates by insurance type yielded no significant evidence that the Medicaid expansion crowded out private coverage in the first year of the 2014 coverage expansion.

While we found substantial increases in Medicaid coverage, our estimates nevertheless suggest that take-up over the first year of the 2014 Medicaid expansion was limited relative to the size of the Medicaid-eligible uninsured population. Our subgroup analyses indicated that take-up was lower among men (compared with women), Hispanics and members of other racial/ethnic groups (compared with non-Hispanic whites and Blacks), and adults under age 36 (compared with older adults). Finally, we found that newly eligible adults in fair or poor health were more likely to gain Medicaid coverage than healthier adults.

Our subgroup analyses suggest that heterogeneity across demographic groups in coverage gains and Medicaid take-up among nondisabled childless adults in poverty were often quite different from the patterns observed among all adults, with women and older adults more likely to gain Medicaid coverage than men and younger adults. Differences in immigration status, which we did not include in our data, may account for some of these subgroup patterns. However, the differences are quite large in magnitude, and we suspect that they primarily reflect enrollment and take-up behavior.

Policymakers interested in increasing Medicaid take-up among the newly eligible may wish to consider these findings in targeting future outreach efforts. A clearer understanding of take-up and crowd-out among the newly eligible may be especially important for anticipating the likely outcome if nonexpansion states adopt the Medicaid expansion in the future, since many of these states have very limited Medicaid eligibility under current law.

Appendix: Methods and Sensitivity Analyses

Sample Construction and Descriptive Statistics

The population of interest in this study consisted of childless adults aged 19–64 living in families with income below 100 percent of the applicable U.S. Department of Health and Human Services Federal Poverty Guideline who were not receiving SSI or SSDI benefits. We used the edited family structure and relationship variables in the final-release NHIS to classify families by the number of children aged 18 or under. The NHIS family structure variables define 18-year-olds as adults even though 18-year-olds count as children for the purposes of determining categorical Medicaid eligibility for parents and caretakers. Using the family relationship codes, we defined *children* as persons aged 18 and under with a parent, caretaker, aunt/uncle, or grandparent aged 19 or over present in the household and then classified adults according to the presence of children in the family.

Income is measured in the NHIS over the calendar year preceding the survey. Our research design requires classification of households by poverty status but does not require a precise measure of the level of income. The final-release NHIS contains a categorical variable reporting previous-year poverty status as defined by the Census Bureau poverty threshold. We used this variable as a criterion for inclusion in our analysis sample. While the NHIS suffers high item nonresponse—typically 22 to 32 percent in recent years—for the exact income question, since 2007 the NHIS has used unfolding brackets to elicit a range for family income from respondents unable to report an exact amount. Unfolding brackets are a survey method that seeks to reduce nonresponse by asking respondents who do not report an exact amount to answer a series of yes-or-no questions about whether their income is above or below a certain threshold. Critically for our research design, the bracket boundaries have been based on Census poverty thresholds since 2011. Unfolding brackets have reduced nonresponse substantially even though many NHIS respondents remain unable or unwilling to report an exact income amount: Only 7.5 percent of persons in the 2014 NHIS (7 percent unweighted) failed to report any bracketed information on family income.[1] We observed whether family income was above or below 100 percent of the FPL for all other individuals in the NHIS. Although we did not use the NHIS imputed income files, the relatively low rate of missing data in our sample is likely to limit the scope for sample selection bias as a result of our choice to exclude individuals with no information on poverty status.

We also excluded all individuals who reported receiving SSI income or Social Security income for their own disability (which is a reasonable proxy for SSDI) at any time in the previous calendar year.

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[1] Authors' calculations, 2014 public-use NHIS. An additional 3 percent of families have a poverty ratio edited to "undefinable" because the number of persons under age 18 is equal to the family size. These families would be excluded from our sample based on age.

Panel A of Table A.1 reports the impact of each step in our sample definition on the number of observations available in the public-use NHIS. For all 50 states and the District of Columbia, the pooled public-use NHIS from 2009 through 2014 contains 16,907 childless adults living in families with income below 100 percent of the FPL and not receiving SSI or SSDI. Panel B of Table A.1 presents unweighted sample sizes by year.

Table A.1. Sample Construction Step-by-Step and Final Sample Sizes by Year and State Group

A. Number of observations in 2009–2014 public-use file	
Sample criteria	N
Adults aged 19–64 at time of survey in person file	362,981
Restrict to families with income below 100 percent of the FPL	51,518
Restrict sample to childless adults	21,572
Restrict to those not receiving SSI or SSDI benefits	16,907

Year	Nonexpansion states	2014 expansion states	Estimation sample	Excluded expansion states	All states
2009	1,050	384	1,434	688	2,122
2010	1,134	447	1,581	776	2,357
2011	1,393	602	1,995	1,001	2,996
2012	1,589	688	2,277	948	3,225
2013	1,341	665	2,006	942	2,948
2014	1,470	821	2,291	968	3,259
All years	7,977	3,607	**11,584**	5,323	16,907

B. Number of observations by year in analysis sample

The sample sizes for the public-use NHIS in Panel A of Table A.1 include the entire country and therefore overstate the sample size available for our analysis. Panel B of Table A.1 shows the sample size available by year and group of states. The sample size increases substantially over the first three years of the sample in part because adult poverty rates climbed sharply following the 2008–2009 Great Recession (DeNavas-Walt and Proctor, 2014). The NHIS's increased sample size after 2011 is also apparent. Our final sample contained 11,584 individuals, of which just over two-thirds were in nonexpansion states. The sample of expansion state residents in 2014 who represent the treated group in our differences-in-differences models contains 821 individuals.

Sample Means

Table A.2 presents sample averages for our insurance coverage and demographic variables for all years (2009–2014) pooled together. This table reports summary statistics separately for the 2014 expansion states and the nonexpansion states. In this table, as in all other calculations reported in this appendix, final person sampling weights are used unless otherwise noted.

Table A.2. 2009–2014 Sample Means for Nondisabled Childless Adults in Poverty

Insurance Coverage		2014 Expansion States	Nonexpansion States
	Any insurance	54.1%	50.2%
	Private insurance	33.8%	35.1%
	Medicaid	14.1%	8.3%
	Any public insurance	20.8%	15.6%
	Non-Medicaid public	6.7%	7.3%
	SCHIP	0.2%	0.4%
	Other public insurance	1.4%	1.5%
	Other government insurance	1.2%	1.0%
	Military insurance	2.1%	2.8%
	Medicare	3.4%	3.2%
Demographics			
	Age	36.2	36.6
	Married	16.4%	18.9%
	Female	46.8%	47.4%
Education			
	Less than high school/GED	18.9%	20.5%
	High school diploma/GED	28.3%	29.7%
	Some college	31.6%	29.3%
	College or more	21.2%	20.5%
Employment			
	Employed	44.5%	47.4%
	Unemployment	20.6%	16.2%
	Not in labor force	34.9%	36.3%
Race/ethnicity			
	Non-Hispanic white	65.2%	56.0%
	Hispanic	10.5%	14.8%
	Non-Hispanic Black	18.7%	23.4%
	Other	5.6%	5.8%
Self-reported health status			
	Good, very good, or excellent	68.0%	68.4%
	Fair or poor	32.0%	31.6%
N (unweighted)		3,607	7,977
Final person sampling weights used			

NOTE: SCHIP = State Children's Health Insurance Program.

Table A.2 indicates that nondisabled adults in poverty are roughly 4 percentage points more likely to have insurance coverage in 2014 expansion states, though these figures include the impact of the Medicaid expansion. As shown in Figure 1 in the main text, both groups of states have similar levels of private insurance coverage and overall coverage, but expansion states had higher levels of Medicaid coverage prior to the expansion. We classified the type of insurance coverage into three categories: private insurance, Medicaid, and non-Medicaid public insurance. The non-Medicaid public insurance category includes all forms of federal, state, and local government insurance coverage other than Medicaid. Respondents were assigned to this category only if they reported public insurance coverage but did not report Medicaid coverage. Inclusion of this category in the analysis gave us an exhaustive classification of the types of insurance reported in the NHIS and allowed us to examine whether adults gaining Medicaid coverage under the ACA might represent transfers of previously insured adults from other government programs. Table A.2 tabulates the specific insurance types that constitute the "Other public" category reported in the figures in the main text. Coverage from other sources of public insurance is similar between 2014 expansion and nonexpansion states, although military insurance is slightly more common in nonexpansion states.

The remainder of Table A.2 reports the average demographic characteristics of our sample. On some basic demographics, including age and gender, respondents from two groups of states are very close to being balanced. On other characteristics, however, including marital status, education, and employment, moderate differences on the order of 1 to 4 percentage points are apparent. The sharpest differences are apparent on race and ethnicity: 65 percent of our expansion state population are non-Hispanic whites versus 56 percent of nonexpansion state respondents. The proportions of respondents that are Black or Hispanic are correspondingly higher in nonexpansion states. We controlled for all demographic characteristics listed in Table A.2 in our regressions.

Protocol for Defining 2014 Expansion States

This section describes our procedure for identifying states with pre-ACA categorical eligibility for childless adults.

An incremental contribution of this study is our focus on a treatment group consisting only of adults who became newly Medicaid-eligible in 2014. At present, published differences-in-differences estimates of the Medicaid expansion's impact on insurance status use less narrowly drawn samples and report differences between all states moving forward with the Medicaid expansion and those not moving forward, or average take-up. While a binary classification of states into expansion and nonexpansion states is informative, generous pre-ACA Medicaid eligibility in many of the expansion states makes it challenging to assign a behavioral interpretation to a differences-in-differences coefficient estimated on the full group of expansion states. One issue is that parents were categorically eligible for Medicaid in all states, and many of the expansion states had relatively generous income limits for parents in 2013 and earlier years: Among families with income below 100 percent of the FPL in the 27 states that had adopted the ACA expansion by the end of 2014, we estimate that roughly 90 percent of parents were

Medicaid-eligible in 2013.[2] In order to isolate the newly eligible, we restricted attention to childless adults throughout this study.

Restricting attention to childless adults was not sufficient to isolate the newly eligible population, however, because childless adults were eligible for coverage in many of the states that adopted the Medicaid expansion by the end of 2014. We estimated that 55 percent of childless adults in these states were Medicaid-eligible in 2013.[3] In order to isolate the newly eligible population, we coded Medicaid income limits for parents and childless adults in all states and excluded states in which childless adults were categorically eligible for comprehensive Medicaid benefits. This led us to exclude the six "early-expansion" states that had implemented the ACA Medicaid expansion before 2014, but it also led us to exclude an additional seven states. We refer to the 14 remaining expansion states as "2014 expansion states" to distinguish them from the 13 "excluded expansion states" in which at least some childless nondisabled adults were eligible for some form of comprehensive Medicaid benefits in 2013. The 2014 expansion states served as the treatment group in our differences-in-differences analysis, while all 24 of the states that had not adopted the ACA Medicaid expansion by the end of 2014 served as the control group.

Table A.3 lists our classification of states as 2014 expansion states, excluded expansion states, and nonexpansion states, along with the highest income limit for each state in 2013. Income limits are reported in modified adjusted gross income (MAGI) as a percentage of the FPL.

The remainder of this subsection explains our procedure for collecting and coding state Medicaid eligibility rules.

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[2] Authors' calculations, 2013 Current Population Survey Annual Social and Economic Supplement (CPS ASEC). Eligibility was imputed by comparing 2012 MAGI approximated following Czajka (2013) to the relevant 2013 MAGI income limits coded as discussed below. We used the Integrated Public Use Microdata Series, Current Population Survey (IPUMS-CPS) for this analysis (Flood et al., 2015).

[3] Authors' calculations, 2013 CPS ASEC.

Table A.3. Classification of State Expansion Status

2014 Expansion States			Excluded Expansion States			Nonexpansion States		
	2013 Income Limits			2013 Income Limits			2013 Income Limits	
State	Parents	Childless Adults	State	Parents	Childless Adults	State	Parents	Childless Adults
AR	17	n.a.	CA†	211	210	AL	11	n.a.
AZ	106	n.a.	CO†	107	10	AK	106	n.a.
IL	195	n.a.	CT†	198	56	FL	30	n.a.
KY	20	n.a.	DC†	216	210	GA	34	n.a.
MD	123	n.a.	DE	107	108	ID	22	n.a.
MI*	54	n.a.	HI	208	208	IN	208	n.a.
ND	54	n.a.	IA	246	240	KS	30	n.a.
NH**	59	n.a.	MA	138	138	LA	18	n.a.
NM	47	n.a.	MN†	102	75	ME	208	n.a.
NV	30	n.a.	NJ†	61	25	MS	24	n.a.
OH	90	n.a.	NY	150	100	MO	18	n.a.
OR	50	n.a.	VT	195	157	MT	47	n.a.
RI	179	n.a.	WA†	133	133	NE	56	n.a.
WV	19	n.a.				NC	41	n.a.
						OK	41	n.a.
						PA	32	n.a.
						SC	62	n.a.
						SD	52	n.a.
						TN	99	n.a.
						TX	15	n.a.
						UT	42	n.a.
						VA	47	n.a.
						WI	201	n.a.
						WY	54	n.a.

Table reports highest income limit for 2013 reported in state MAGI conversion plans in terms of MAGI as a percentage of FPL.
Table reflects eligibility for comprehensive benefits (Medicaid state plan, Medicaid Managed Care, or waiver plans covering primary care, specialists, and hospital care). Plans with enrollment closed in 2011 or earlier are excluded.
* Michigan ACA expansion was implemented beginning in the second quarter of 2014.
** New Hampshire ACA expansion was implemented on August 15, 2014. We excluded data from the third quarter of 2014 for New Hampshire as a wash-out.
† indicates "early-expansion" states adopting the ACA expansion in part or in full prior to 2014.

MAGI Conversion Plans as Data Source on Medicaid Limits

In addition to the Medicaid coverage expansion, the ACA mandated major changes in how household income would be defined for the purposes of establishing Medicaid eligibility. Beginning in 2014, the ACA mandated that eligibility determinations for Medicaid, the Children's Health Insurance Program (CHIP), and exchange subsidies be made on the basis of MAGI.[4] Because pre-ACA net income concepts varied dramatically from state to state, the ACA required every state to calculate MAGI-based eligibility standards in 2013. MAGI conversion was required for groups involving parents/caretaker relatives, pregnant women, children under age 19, and childless adults. Even income standards that might seem irrelevant under the ACA Medicaid expansion (i.e., MAGI income limits for parents and children below 138 percent of the FPL in states moving forward with the expansion) were converted for several reasons, most notably that the health care costs of newly eligible and previously eligible beneficiaries are reimbursed at different Federal Medical Assistance Percentage rates.

Our primary sources for coding state Medicaid eligibility rules were the MAGI conversion plans submitted by the states to CMS. We collected MAGI conversion plans for 49 states and the District of Columbia from the CMS website (Medicaid.gov, undated[a]).[5] Two members of the research team reviewed the MAGI conversion plans available from the state web pages on the CMS website and independently entered the net income and MAGI income standards for every pathway covering families or childless adults. We did not enter pathways that applied to pregnant women or children. According to the Kaiser Family Foundation's survey of Medicaid program rules, the net income standard for pregnant women was above the poverty line in every state by 2013, meaning that the 2014 Medicaid expansion does not induce any variation in the eligibility of pregnant women (Heberlein, Brooks, Alker, et al., 2013).

The MAGI conversion plans often did not provide many details on the covered population or the type of benefits provided for each eligibility group. For the mandatory Section 1931 eligibility pathways (i.e., families meeting the former Aid to Families with Dependent Children [AFDC] income standards), it is clear that families with children are covered and traditional Medicaid benefits or Medicaid managed care is provided. However, there was significant ambiguity regarding the 1115 waiver pathways because the 1115 waivers encompass a very diverse group of programs and populations. We coded all 1115 waivers that were not clearly restricted to narrow benefits (e.g., family planning services) or targeted only to pregnant women or children.

Drawing guidance from the Kaiser Family Foundation's surveys of Medicaid eligibility (Heberlein, Brooks, Guyer, et al., 2011; Heberlein, Brooks, Guyer, et al., 2012; Heberlein, Brooks, Alker, et al., 2013), the first reviewer examined approved 1115 waivers obtained from the CMS website to identify the target population and scope of benefits for 1115 waivers

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[4] MAGI is a tax term that includes total gross income minus allowed deductions, plus certain tax-exempt income types, such as Social Security income, interest, and foreign income.

[5] All conversion plans were downloaded on March 13, 2015. All states except Massachusetts provided a conversion plan; we discuss our handling of Massachusetts below.

reported in the conversion plans that were ambiguous on these dimensions.[6] We determined that the MAGI conversion plans included a number of benefit plans that were significantly narrower than Medicaid in one of several ways, and we excluded from consideration 1115 pathways that met any of the following restrictions:

- **Limited-benefit plans that do not cover both primary care and inpatient hospital care.** Some 1115 plans cover only family planning services, or they cover only primary care in limited settings but do not cover hospital care. Our view is that these plans should be excluded from consideration because they are much less comprehensive than traditional Medicaid.
    - Maryland: The Primary Adult Care program provided "a limited primary care health benefit package to uninsured adults" (CMS, 2013a).
    - Michigan: The Adult Benefit Waiver program provided "a limited ambulatory care benefit package" (CMS, 2014).
    - Missouri: Gateway to Better Health is a program limited to St. Louis County providing "a limited primary care package" to beneficiaries who receive care at specified Federally Qualified Health Centers (CMS, 2013b).
    - Vermont: The Catamount health waiver program, available to those with incomes up to 300 percent of the FPL, is described as limited. (It also does not affect our analysis because another 1115 waiver goes up to 150 percent of the FPL.)
- **Plans that provide premium assistance for private employer-sponsored coverage or the individual market.** Some of these plans are limited to employees of participating employers. We also suspect that premium support for employer-sponsored insurance or individual market coverage is likely to be reported as private insurance in the NHIS rather than Medicaid.
    - Arkansas: The Safety Net Benefit Program provides premium assistance for employer-sponsored insurance (CMS, 2012b).
    - Oklahoma: Insure Oklahoma provides premium assistance for employer-sponsored insurance or individual market coverage (CMS, 2011).
- **Plans that were closed to new enrollment in 2011 or earlier.** While there may be a stock of enrollees in some closed plans prior to 2014, we assumed that closed plans were likely to have negligible enrollment.
    - Arizona: The 1115 program for childless adults with incomes up to 100 percent of the FPL was phased out in mid-2011 (CMS, 2012a).
    - Indiana: In the Healthy Indiana program, enrollment for childless adults was closed.
    - Maine: The 1115 program for childless adults up to 100 percent of the FPL was closed prior to 2014 (the timing is unclear from the waiver).
    - Wisconsin: BadgerCare Core covered childless adults up to 200 percent of the FPL with a limited benefits package. (The benefits package meets our criterion of covering both primary and inpatient care.) However, a binding enrollment cap was imposed shortly after coverage was expanded to childless adults, and the state

[6] All waiver applications cited were obtained from the Medicaid website on May 22–25, 2015 (Medicaid.gov, undated[b]).

reported in a waiver application that the childless adult waiting list was approximately seven times larger than the enrolled population in 2013.

Other plans that deviate in some way from traditional Medicaid but offer a comprehensive benefit were included in our database. In particular, we included

- Medicaid managed care
- state plans that charge premiums to some beneficiaries.

There are two states for which no family or childless adult income standards were reported in the MAGI conversion reports. A MAGI conversion plan for Massachusetts was not submitted to CMS. This may be due to the fact that Massachusetts had already implemented health care reform prior to the ACA. In any event, Massachusetts does not contribute any identifying variation to our research design because all low-income adults were eligible for Medicaid or subsidized coverage prior to 2014 (Heberlein et al., 2011). We assigned Massachusetts MAGI limits of 138 percent of the FPL for families and childless adults in all years so that all adults in our analysis sample were classified as Medicaid-eligible.

The other state missing MAGI conversion information was Texas. Texas submitted a conversion report, but the conversion report described the income limit for Section 1931 families as “AFDC” without further elaboration. We used the 2014 MAGI income standard of 15 percent for families reported for Texas by CMS as our proxy for the MAGI limit, and we used this standard to impute pre-ACA eligibility throughout the 2009–2013 period.[7]

The MAGI conversion plans are a valuable data source, but it was necessary to draw on additional references and data sources to verify the accuracy and completeness of the income standards reported in the conversion plans. For instance, one limitation of using the MAGI conversion plans is that states could choose not to provide conversion estimates for Medicaid pathways that would not be used in 2014 and later years.[8] The MAGI conversion plans also do not report pathways covering childless adults that were established in two states (Connecticut and Minnesota) following enactment of the ACA, nor do they report pathways in several other states (New York and Washington). We assigned MAGI limits to these states as follows:

- Connecticut: Childless adults, 56 percent of the FPL (early ACA expansion as reported in Sommers, Arntson, et al. [2013] and Herz [2012])
- Minnesota: Childless adults, 75 percent of the FPL (early ACA expansion as reported in Sommers, Arntson, et al. [2013] and Herz [2012])
- New York: Childless adults, 100 percent of the FPL; parents, 150 percent of the FPL (Family Plus taken from Kaiser Family Foundation)
- Washington: Childless adults and parents, 133 percent of the FPL (conversion of the state Basic Health plan to 1115 waiver, coded from Kaiser Family Foundation).

We compared the results of our coded income standards to the Kaiser Family Foundation’s widely used survey of state Medicaid eligibility to identify any pathways that were omitted from

[7] The only change in eligibility for Texas parents reported by the Kaiser Family Foundation is a 1-percentage-point decline in the standard for working parents between 2012 and 2013. Texas does not report offering any optional eligibility groups for families in the MAGI report or the other sources we consulted.

[8] In practice, very few of the discontinued pathways represented the highest income limit for families, so this limitation did not affect our ability to classify state Medicaid policies in earlier years.

the MAGI conversion reports and to assist in locating additional sources clarifying the status of pathways that were incompletely described in the conversion plans (Heberlein, Brooks, Guyer, et al., 2012; Heberlein, Brooks, Alker, et al., 2013) We also drew on a Congressional Research Service report to identify the statutory authority used by certain states to support early implementation of the ACA expansion (Herz, 2012). For 1115 eligibility pathways, we also attempted to compare our database to the 2012 program rules coded in the Urban Institute's TRIM3 microsimulation model (TRIM3, 2012, accessed May 18–27, 2015). The TRIM3 rules agreed with our coded limits except for the following cases, which we resolved in favor of the MAGI conversion reports:

- Hawaii: TRIM3 reported 100 percent for parents versus 208 percent in the MAGI conversion plan.
- Minnesota: TRIM3 reported 215 percent for parents versus 102 percent in the MAGI conversion plan. We follow Sommers, Kenney, and Epstein (2014) in coding Minnesota's income limit as 75 percent of the FPL.[9]
- New Jersey: TRIM3 reported 200 percent for parents versus 27 percent (61 percent MAGI) in the MAGI conversion reports. The 200-percent threshold corresponds to the NJ FamilyCare waiver program. However, NJ FamilyCare's 2013 annual report states that enrollment is closed for all pathways applicable to adults with income above 100 percent of the FPL. The same source reports that the number of beneficiaries in the state was essentially flat between December 2010 and December 2013, consistent with enrollment being closed to higher-income adults. We ignored the 200-percent eligibility pathway for New Jersey.
- Tennessee: TRIM3 reported 400 percent for parents, which we did not find documented elsewhere and therefore ignored.

Statistical Methods

We applied a regression-adjusted differences-in-differences approach to estimate the effect of expanded Medicaid eligibility. Because we used multiple years of pre-2014 data, our main specification relaxed the assumption that outcomes would have evolved identically in expansion and opt-out states if the expansion states had also opted out. We modeled outcome of interest y_{ist} for individual *i* in state *s* in quarter *t* as the sum of a state-specific intercept term (μ_s) for people in state *s*, a time effect (v_t) that varied freely by year and quarter but was common to all states, the effect of *i*'s demographic characteristics (X_{it}), the effect of a linear time trend (t, coded as the number of quarters elapsed since the end of 2008) that was present only for the expansion

[9] "Based on a comparison of MinnesotaCare and General Assistance Medical Care (GAMC) populations that were transitioned to Medicaid at the outset of the expansion, versus March 2013 expansion enrollment statistics provided by the Minnesota Department of Human Services. GAMC had an income cutoff of 75 percent of poverty but also used an asset test (which was eliminated under the 2010 expansion). MinnesotaCare offered coverage as high as 250 percent of poverty but had a state cap on total spending that often limited enrollment" (Sommers, Kenney, and Epstein, 2014, p. 79).

states, the effect of a binary variable (T_{st}) that equaled 1 if the Medicaid expansion was in effect in state *s* during quarter *t* and equaled 0 otherwise, and a mean-zero error term (ε_{ist}):

$$y_{ist} = \tau_T T_{st} + X_{it}\beta_X + \beta_T t\mathbf{1}\{s \text{ is expansion state}\} + \mu_s + \nu_t + \varepsilon_{ist},$$

where β_X is a vector of coefficients on individual characteristics X_{it} and β_T is the coefficient on the differential time trend. The parameter of interest is τ_T, which can be interpreted as the average effect of the Medicaid expansion on outcome *y*, holding an individual's demographic characteristics constant. The coefficient τ_T directly captures the predicted change in the probability of y_{ist} associated with the Medicaid expansion and is scaled as a fraction; for example, if τ_T in a model for any insurance coverage were 0.1, we would conclude that the Medicaid expansion had increased the probability that a nondisabled childless adult in poverty had insurance coverage by 10 percentage points.

The basic differences-in-differences approach effectively compares trends over time in expansion and nonexpansion states to assess whether trends diverged after the Medicaid expansion took effect on January 1, 2014. Because the approach focuses on changes within states rather than absolute differences, it enables us control for any fixed differences between expansion and nonexpansion states, including permanent unobserved differences. For example, the approach controls for underlying differences in the quality of hospitals and physicians across states, time-invariant differences in the stigma or hassle associated with Medicaid enrollment, fixed differences in individuals' awareness of the law, and any other time-invariant factors that may influence outcomes. The inclusion of individual characteristics X_{it} corrects the estimated differences-in-differences effect τ_T for differences in the composition of our population of interest across states and over time and improves the statistical power of the model by reducing the unexplained variation in the outcome y_{ist}.

The potential advantage of our model over differences-in-differences without trends is that the benchmark for changes in outcomes in expansion states takes into account trends observed in those states in 2009 through 2013, as well as any deviation from 2009 through 2013 trends observed in the opt-out states in 2014. While the differences-in-differences model with group-specific time trends can be unreliable when the number of time periods observed before policy implementation is small relative to the number of time periods observed in postimplementation data (Wolfers, 2006), we have enough preimplementation time periods (20 quarters) relative to postimplementation time periods (four quarters) that this is unlikely to be a major concern. Another risk of using trends is that unmodeled preimplementation activities (e.g., outreach starting in 2013) could affect our estimated trends, potentially biasing the estimated effects of Medicaid expansion on coverage downward. We address this possibility below by estimating alternative models that drop 2013 data as a preimplementation wash-out period.

For most of the 2014 expansion states, T_{st} is equal to 1 in 2014 and 0 in all other years. Two states in our sample—Michigan and New Hampshire—began the Medicaid expansion after January 1, 2014. Michigan's ACA expansion was implemented beginning in the second quarter of 2014. We coded the treatment indicator T_{st} for Michigan to 0 in the first quarter of 2014 and to 1 in subsequent quarters. New Hampshire's ACA expansion was implemented on August 15,

2014. We excluded data from the third quarter of 2014 for New Hampshire as a wash-out period and coded the treatment indicator T_{st} for New Hampshire to 1 only in the fourth quarter of 2014.

Alternative Specifications

In this appendix, we also report basic differences-in-differences estimates that do not control for a differential time trend in the expansion states:

$$y_{ist} = \tau_T T_{st} + X_{it}\beta_X + \mu_s + \nu_t + \varepsilon_{ist} .$$

This model makes stronger assumptions than the model in our main specification, and so it delivers more precise estimates, but it is not robust to violations of the assumption that expansion and nonexpansion states would have had parallel trends in coverage in the absence of the ACA.

To address the possibility that Medicaid expansion activities would affect coverage or reporting behavior prior to the 2014 implementation date, we estimated models both with and without differential trends that dropped 2013 data.

Estimation and Statistical Inference

Our regression specification is a linear probability model (LPM), and we estimated this model by ordinary least squares (OLS) regression. While the LPM does not explicitly account for the binary nature of the outcome variable, it can be justified as an approximation to the conditional expectation function of y_{ist} given the explanatory variables (Angrist and Pischke, 2008). Because all of our explanatory variables except the linear time trend are all binary or categorical, inconsistency arising from values of the regression function outside the unit interval may be limited in our specification. Our choice of the LPM over a fully parametric binary choice model, such as logistic regression, was motivated largely by our concerns about clustering—i.e., unmodeled correlation of the error term within states and over time. Standard maximum likelihood estimates of the logistic regression model are inconsistent in this setting, while OLS remains consistent as long as the error term is uncorrelated with the regressors.

Classical inference assuming independent and identically distributed error terms is inappropriate in our setting for several reasons. First, our outcomes are binary variables, leading to conditional heteroskedasticity. Second, the stratified sampling design of the NHIS requires estimation procedures that account for sampling design using the stratum and primary sampling unit variables on the NHIS files in addition to sampling weights. However, standard survey estimation procedures using the NHIS design variables are inappropriate for our application because these procedures assume statistical independence of error terms across survey strata, and many states contain multiple strata—the public-use NHIS identifies 300 unique strata.

We believe that error terms are most likely correlated across time periods within states. To the extent that state health policy, economic conditions, labor market regulation, and other factors that vary at the state level cause states to experience different dynamics of insurance coverage prior to ACA implementation, error terms will be correlated across states within time. This can be the case even if our regression model is correctly specified. Furthermore, our identifying variation exists at the state level. While our regression estimates would remain consistent in such a scenario, inference that uses the NHIS survey design variables may lead to underestimates of standard errors and overrejection of null hypotheses.

We used standard errors clustered by state to account for arbitrary correlation of the error term within state as well as conditional heteroskedasticity. Survey strata are nested within states, so our inference approach should be viewed as more conservative than inference using the NHIS design variables. Our main estimation sample contains 38 clusters (states). We calculated *p*-values and confidence intervals using a *t*-distribution with 37 degrees of freedom, as suggested by Cameron and Miller (2015).

Results and Robustness Checks

Table A.4 presents our main regression estimates (Column 5) in addition to a range of other specifications used to assess the robustness of our results. Column 1, which includes no covariates except for state and time fixed effects, is used to assess the significance of the "unadjusted differences-in-differences" effects of Medicaid expansion referenced in the main text. Column 2 presents analogous estimates using all available pre-ACA years (2009–2013) as the baseline instead of only 2013 data.

Column 3 reports basic regression-adjusted differences-in-differences estimates that control for individual covariates but do not include differential trends in coverage in the expansion states. Omitting the differential trends yields estimated increases in Medicaid and private coverage that are each roughly 3 percentage points higher than in our main specification. However, the point estimates for our main model (Column 5) fall well within the 95-percent confidence intervals implied by Column 3, and vice versa, and we view these findings as being qualitatively similar. The model without differential time trends (Column 3) does yield a meaningfully larger increase in overall insurance coverage (14.9 percentage points versus 8.9 percentage points in our main model), though this point estimate also falls within the 95-percent confidence interval for our main model.

Table A.4. Differences-in-Differences and Regression-Adjusted Estimates of Medicaid Expansion Effect

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Years	2013–2014	2009–2014	2009–2014	2009–2012, 2014	2009–2014	2009–2012, 2014	2009–2014	2009–2014	2009–2014
Control for covariates?			Y	Y	Y	Y	Y	Y	Y
State group–specific time trends?					Linear	Linear	Quadratic		Linear
Model excludes employment?								Y	Y
Any insurance	0.126	0.173	0.149	0.160	0.089	0.100	0.080	0.161	0.106
	(0.030)	(0.028)	(0.027)	(0.028)	(0.041)	(0.043)	(0.043)	(0.028)	(0.042)
	0.000	0.000	0.000	0.000	0.034	0.025	0.072	0.000	0.016
Medicaid	0.149	0.158	0.154	0.155	0.126	0.108	0.120	0.163	0.135
	(0.033)	(0.023)	(0.023)	(0.024)	(0.027)	(0.028)	(0.033)	(0.023)	(0.026)
	0.000	0.000	0.000	0.000	0.000	0.000	0.001	0.000	0.000
Private insurance	-0.001	0.016	-0.006	-0.002	-0.032	-0.027	0.027	-0.008	-0.029
	(0.033)	(0.023)	(0.019)	(0.018)	(0.035)	(0.045)	(0.044)	(0.020)	(0.034)
	0.967	0.501	0.758	0.926	0.371	0.559	0.534	0.700	0.386
Non-Medicaid public insurance	-0.010	0.013	0.014	0.021	0.006	0.029	-0.055	0.019	0.011
	(0.019)	(0.010)	(0.011)	(0.010)	(0.019)	(0.021)	(0.029)	(0.011)	(0.020)
	0.611	0.232	0.197	0.050	0.755	0.178	0.060	0.095	0.592
N	4,297	11,584	11,584	9,578	11,584	9,578	11,584	11,584	11,584

SOURCE: 2009–2014 National Health Interview Survey.
NOTES: Each cell reports the differences-in-differences regression coefficient for the effect of the 2014 Medicaid expansion from a separate OLS regression model.
Standard errors clustered on state (38 clusters) are in parentheses; two-sided *p*-values for equality to 0 are reported below standard errors. *P*-values were calculated based on a *t*-distribution with 37 degrees of freedom.
The sample consists of nondisabled childless adults aged 19–64 with annual family income below 100 percent of the FPL.
States with any categorical eligibility for nondisabled childless adults in 2013 are excluded from the sample. These 13 states are California, Colorado, Connecticut,

Delaware, the District of Columbia, Hawaii, Iowa, Massachusetts, Minnesota, New Jersey, New York, Vermont, and Washington.
All regressions include state and quarter fixed effects. Covariates include age categories (19–25, 26–35, 46–55, and 56–64), gender, binary marital status, four racial/ethnic groups (non-Hispanic white, Hispanic, non-Hispanic Black, and other), educational attainment (less than high school, high school diploma or GED, some college, college diploma, and more than college), and employment status (employed, unemployed, not in labor force, and unknown).

Columns 4 and 6 report regression specifications that are identical to those in Columns 3 and 5 but that are estimated on a sample that excludes data from 2013 as a wash-out period that could potentially be affected by state-specific activities related to the 2014 Medicaid expansion. These estimates do not differ meaningfully from the corresponding estimates that include 2013 data.

Column 7 adds a quadratic term to the specification of the differential time trend in the expansion states. The estimated impacts on overall insurance coverage and Medicaid are very close to those estimated in our main model, as reported in Column 5. The estimated effect of expansion on private coverage is insignificant and positive, while the effect on non-Medicaid public coverage becomes negative and statistically significant. This specification yields estimated impacts on private and other public coverage that do not resemble any of the other specifications we estimated. We view these results as a warning that the quadratic specification of the differential trends may be undesirably sensitive to the data from 2009, when the expansion states had lower rates of private coverage and higher rates of other public coverage relative to the nonexpansion states than in later years.

Finally, Columns 8 and 9 report regression specifications that are identical to those in Columns 3 and 5 but that omit the control variable for current employment status; some readers may be concerned that it is inappropriate to include employment status as a control variable because labor supply may respond to Medicaid eligibility (Garthwaite, Gross, and Notowidigdo, 2014). These estimates do not differ meaningfully from the corresponding estimates based on our main specification.

Comparison with Other Published Differences-in-Differences Estimates

We stated in the main text that our basic findings about take-up with limited crowd-out were qualitatively consistent with the other published differences-in-differences estimates using large federal surveys. Here we provide a more detailed comparison of our results and the two peer-reviewed studies to date that used large federal surveys to estimate differences-in-differences models for the coverage effects of Medicaid expansion. An earlier paper by Sommers et al. also used a differences-in-differences research design with tracking survey data from the Gallup-Healthways Well-Being Index (Sommers, Gunja, et al., 2015b). Sommers et al. were not able to reliably distinguish between different sources of coverage, and they did not distinguish between expansion states with different pre-ACA Medicaid policies, but they also found significant reductions in uninsurance (5.2 percentage points) among adults with family income of 138 percent of the FPL or below.

Compared with the 2016 paper by Wherry and Miller that also analyzed the 2014 NHIS, we found larger increases in Medicaid and overall coverage. Wherry and Miller estimated that Medicaid expansion increased overall insurance coverage by 7.4 percentage points and increased Medicaid coverage by 10.5 percentage points. We also found slightly weaker evidence of reductions in private insurance, which Wherry and Miller calculate decreased by a marginally significant 3.7 percentage points ($p = 0.07$) in their main estimates.

There are two major differences between our research design and Wherry and Miller's, which should have offsetting impacts on the magnitude of our estimated effects. On the one hand, because we focus on a population that experienced sharper increases in Medicaid eligibility, our estimates may tend to be larger than those reported by Wherry and Miller. Specifically, we excluded adults with incomes between 100 percent and 138 percent of the FPL, we excluded parents, we used data collected in all quarters of 2014 while they dropped the first two quarters as a wash-out period, and we excluded eight expansion states included in their analysis (California, Colorado, Connecticut, Hawaii, Iowa, Minnesota, New Jersey, and Washington).

On the other hand, because, unlike Wherry and Miller, we control for differential linear trends in expansion and nonexpansion states, our estimates may tend to be smaller. The models reported in Column 3 of Table A.3, which omit pre-trends and thus correspond most closely to Wherry and Miller's specification, yield much larger estimated increases in overall coverage and Medicaid coverage with no evidence of reductions in private coverage.

Courtemanche, Marton, and Yelowitz (2016) analyzed data from the 2013–2014 American Community Survey to track changes in insurance by poverty ratio and state expansion status. They estimated that adults in poverty in expansion states became 8.8 percentage points more likely to be insured between 2013 and 2014, versus 4.7 percentage points in nonexpansion states, suggesting that Medicaid expansion was associated with a 4.1-percentage-point increase in insurance coverage for adults in poverty. They also reported rates of public and private coverage. Those estimates imply that Medicaid expansion was associated with a 6.1-percentage-point increase in public coverage and a 1.9-percentage-point decrease in private coverage. However, this study is not directly comparable to ours because it groups childless adults and parents together, and because the main models report results for all 50 states and the District of Columbia.

Take-Up and Crowd-Out Rates

Our regression coefficient for Medicaid coverage may be interpreted directly as a take-up rate because it captures the change in the probability of Medicaid coverage that resulted from gaining eligibility. To provide additional context for interpreting this estimate, we can scale this estimate by the 2013 uninsurance rate for nondisabled childless adults in poverty in our sample, which we estimated to be 42.0 percent. The increase in Medicaid coverage was equivalent to 30 percent of the pre-ACA uninsurance rate. Similarly, the reduction in uninsurance was equivalent to 21.3 percent of the pre-ACA uninsurance rate.

We can also scale our estimated change in private coverage by the change in Medicaid coverage to obtain a crowd-out rate in terms of the change in private coverage per new Medicaid enrollee. The point estimate from the model controlling for differential trends is 25 percent crowd-out, but we note that the change in private coverage is not significantly different from 0, and so the crowd-out ratio is too imprecisely estimated to be very informative: The 95-percent confidence interval based on delta-method standard errors covers crowd-out rates from –26 percent to 76 percent. The crowd-out estimate from the model without differential trends is close to 0 (4 percent), with a delta-method 95-percent confidence interval covering [–21 percent, 28.5 percent].

Because the ratio of two normal random variables can be highly non-normal if the denominator is close to 0, we also used the variance-covariance matrix of our regression estimates for private coverage and Medicaid to examine the distribution of the crowd-out ratio by simulation. The resulting 95-percent confidence interval was [–38 percent, 71 percent], which is not meaningfully different from the delta-method estimate. This is not surprising; the estimated increase in Medicaid coverage is highly significant, so there were essentially no draws that were very close to 0. The estimated increase in Medicaid is even larger when pre-trends are omitted, so the delta-method confidence interval falls within 1 percentage point of the simulated confidence interval.

Assessing the Importance of Differential Time Trends

The only specification choice that had a substantial effect on our estimates was the inclusion of a differential time trend in the expansion states: The estimated effect of the expansion on Medicaid and private coverage was slightly less positive when differential trends were included, and the impact on overall coverage was lower. To assess the importance of differential trends, Table A.5 reports the coefficients β_T for the differential time trends, standard errors, and *p*-values for equality to 0. Coefficients and standard errors were multiplied by 4 so that they could be interpreted as the predicted annual increase in insurance coverage in expansion states relative to nonexpansion states after controlling for individual covariates and state and time fixed effects.

As suggested by Figure 1 in the main text, insurance coverage was trending upward in expansion states relative to nonexpansion states by 2 percentage points per year prior to 2014. Point estimates in Column 1 suggest that this increase was accounted for by both Medicaid coverage and private coverage, although the private coverage pre-trend was imprecisely estimated.

In our main model (Column 1 of Table A.5), the coefficients β_T on the linear pre-ACA time trends were not statistically significant when compared with the *t*-distribution that we considered appropriate for our clustering scheme. However, the pre-trends for any insurance ($p = 0.102$) and Medicaid coverage ($p = 0.105$) were extremely close to being significant at the 10-percent level. In fact, comparison to critical values from the normal distribution (1.64 for the 10-percent level) rather than the *t*-distribution with 37 degrees of freedom (1.69) would identify both of these coefficients as marginally significant. Thus, even though the differential trends were insignificant at conventional levels, they were close enough to significance that we did not feel comfortable relying on the parallel trends assumption for our main estimates.

A potential limitation of our main specification is that the pre-ACA time trend could be contaminated by postimplementation dynamics if 2014 Medicaid expansion led to a trend break in insurance coverage, with effects increasing over the course of 2014. We viewed this as a minor concern for our main specification because our dataset contained five years of preimplementation data but only one year of postimplementation data. To examine whether failure to allow for dynamics following implementation affected our estimates of differential time trends by expansion status, we estimated a model that allowed the differences-in-differences treatment effect to vary freely from quarter to quarter in 2014. In this specification, any postimplementation trend in outcomes was absorbed by the quarter-specific treatment effects and therefore could not affect our estimate of the differential trend in insurance for expansion states. The resulting estimates of the time trend are reported in Column 2 of Table A.5. The estimates are very close to the time trends estimated in our main specification, suggesting that postimplementation dynamics do not account for the differential trends we estimate.

Finally, Column 3 of Table A.5 reports coefficients on trends that omitted 2013 data as a wash-out period, since it is reasonable to attribute differential trends to the early effects of activities associated with Medicaid expansion. The magnitudes of the preexisting trends were similar to the estimates including 2013 data, however, and the trend in Medicaid coverage was larger and clearly significant (p = 0.034) when 2013 data were excluded. We have included Table A.5 to allow interested readers to understand our motivations for favoring the model with differential pre-ACA trends as our main specification and to place the estimates with and without differential trends in context.

Table A.5. Coefficients on Preexpansion Linear Time Trends

	(1)	(2)	(3)
Years	2009–2014	2009–2014	2009–2012, 2014
Control for covariates?	Y	Y	Y
Quarter-specific effects of 2014 expansion?*		Y	
Preimplementation trend: Any insurance	0.021	0.022	0.018
	(0.013)	(0.012)	(0.013)
	0.102	0.067	0.185
Preimplementation trend: Medicaid	0.010	0.013	0.014
	(0.006)	(0.006)	(0.007)
	0.105	0.037	0.034
Preimplementation trend: Private insurance	0.009	0.007	0.008
	(0.009)	(0.009)	(0.012)
	0.312	0.467	0.536
Preimplementation trend: Non-Medicaid public insurance	0.003	0.004	-0.002
	(0.004)	(0.004)	(0.005)
	0.506	0.281	0.656
N	11,584	11,584	9,578

Table reports coefficients on linear time trends specific to expansion states included in models from columns 5–6 in Table A.4.

Time trend coefficients and standard errors scaled up by 4 to represent predicted increase in insurance coverage over 1 year for expansion states relative to nonexpansion states.

Regressions also control for sociodemographics, year-quarter time effects, state fixed effects, and dummy variable for expansion states in 2014.

* Model contains four dummies for the quarters of 2014 interacted with the indicator for expansion state status.

Standard errors clustered by state (38 clusters) are in parentheses.

P-values for equality of time trend to 0 are below standard errors.

P-values are based on *t*-distribution with 37 degrees of freedom and standard errors clustered on state.

Subgroup Analyses

We estimated that the majority of nondisabled childless adults in poverty who gained Medicaid eligibility in 2014 did not enroll in Medicaid. In order to learn more about the characteristics of those individuals who gained insurance because of the ACA Medicaid expansion, we conducted several subgroup analyses. We interacted the treatment variable in our main differences-in-differences regression models with indicators for group membership to yield the following model, which allows the effect of the 2014 Medicaid expansion to vary freely across groups indexed by G:

$$y_{ist} = \tau_T T_{st} + \sum_{G \in \mathcal{G}} \tau_G T_{st} \mathbf{1}\{i \text{ is in } G\} + X_{it}\beta_X + \mu_s + \nu_t + \varepsilon_{ist}$$

where $\mathcal{G}$ is a set of subgroups excluding a base category. We estimated this model for the following groups:

- race (base category: non-Hispanic white)
- gender (base category: male)
- age (base category: ages 19–25)
- self-reported health status (base category: good, very good, or excellent health).

Table A.6 reports coefficients from these four models for three sets of outcomes: any insurance coverage, Medicaid, and private coverage. The Medicaid estimates are reported in Figure 3 in the main text. Similar figures for the other insurance types are presented as Figures A.1 and A.2. As discussed in the main text, subgroup patterns for any insurance are very similar to the patterns observed for Medicaid, while none of the subgroups experienced a significant decrease in private coverage.

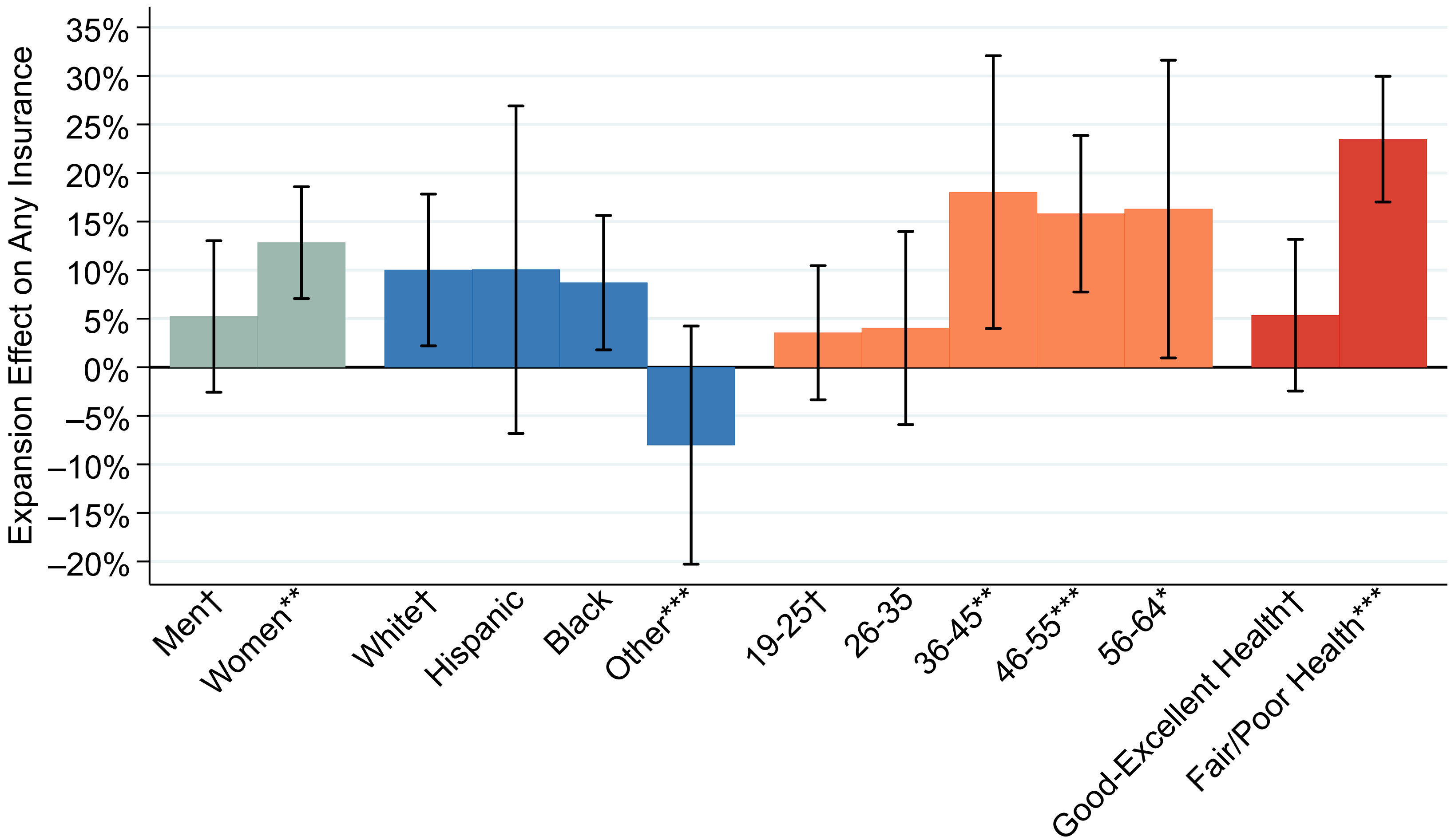
Table A.6. Subgroup Effects of Medicaid Expansion on Insurance Coverage

Years	2009–2014	2009–2014	2009–2014	2009–2014	2009–2014	2009–2014	2009–2014	2009–2014	2009–2014	2009–2014	2009–2014	2009–2014	2009–2014
Control for covariates?	Y	Y	Y	Y	Y	Y	Y	Y	Y	Y	Y	Y	Y
State group–specific time trends?	Linear	Linear	Linear	Linear	Linear	Linear	Linear	Linear	Linear	Linear	Linear	Linear	Linear
Subgroup model	Gender	Gender	Race	Race	Race	Race	Age	Age	Age	Age	Age	Health	Health
Subgroup	Men	Women	Non-Hispanic white	Hispanic	Non-Hispanic Black	Other non-Hispanic	19–25	26–35	36–45	46–55	55–64	Good or better health	Fair or poor health
Effect type	Base	Interaction	Base	Interaction	Interaction	Interaction	Base	Interaction	Interaction	Interaction	Interaction	Base	Interaction
Any insurance	0.052	0.076	0.100	0.000	-0.013	-0.180	0.036	0.005	0.145	0.123	0.127	0.054	0.181
	(0.038)	(0.028)	(0.039)	(0.083)	(0.034)	(0.061)	(0.034)	(0.049)	(0.069)	(0.040)	(0.076)	(0.039)	(0.032)
	0.183	0.011	0.013	0.997	0.704	0.005	0.304	0.922	0.044	0.004	0.101	0.173	0.000
Medicaid	0.085	0.084	0.132	-0.082	0.073	-0.224	0.065	-0.023	0.160	0.205	0.099	0.088	0.192
	(0.029)	(0.019)	(0.033)	(0.043)	(0.061)	(0.038)	(0.024)	(0.044)	(0.064)	(0.035)	(0.057)	(0.021)	(0.048)
	0.006	0.000	0.000	0.062	0.238	0.000	0.011	0.605	0.017	0.000	0.089	0.000	0.000
Private insurance	-0.026	-0.011	-0.028	0.103	-0.071	-0.040	-0.029	0.032	-0.003	-0.075	0.039	-0.032	0.004
	(0.032)	(0.037)	(0.026)	(0.070)	(0.068)	(0.064)	(0.037)	(0.087)	(0.051)	(0.056)	(0.052)	(0.033)	(0.048)
	0.421	0.765	0.282	0.148	0.303	0.534	0.432	0.710	0.958	0.187	0.466	0.342	0.937
N	11,584	11,584	11,584	11,584	11,584	11,584	11,584	11,584	11,584	11,584	11,584	11,584	11,584

SOURCE: 2009–2014 National Health Interview Survey.
NOTES: Subgroup analysis was conducted using regression specifications in which a differences-in-differences treatment variable was interacted with dummies for subgroup membership.
The base category coefficient is the effect of the expansion, but the interaction coefficient is the difference in the effect of the expansion between the subgroup at hand and the base category.
Standard errors clustered on state (38 clusters) are in parentheses; two-sided *p*-values for equality to 0 are reported below standard errors. *P*-values were calculated based on a *t*-distribution with 37 degrees of freedom.
The sample consists of nondisabled childless adults aged 19–64 with annual family income below 100 percent of the FPL.
States with any categorical eligibility for nondisabled childless adults in 2013 are excluded from the sample. These 13 states are California, Colorado, Connecticut, Delaware, the District of Columbia, Hawaii, Iowa, Massachusetts, Minnesota, New Jersey, New York, Vermont, and Washington.
All regressions include state and quarter fixed effects. Covariates include age categories (19–25, 26–35, 36–45, 46–55, and 56–64), gender, binary marital status, four racial/ethnic groups (non-Hispanic white, Hispanic, non-Hispanic Black, and other), educational attainment (less than high school, high school diploma or

GED, some college, college diploma, and more than college), and employment status (employed, unemployed, not in labor force, and unknown).

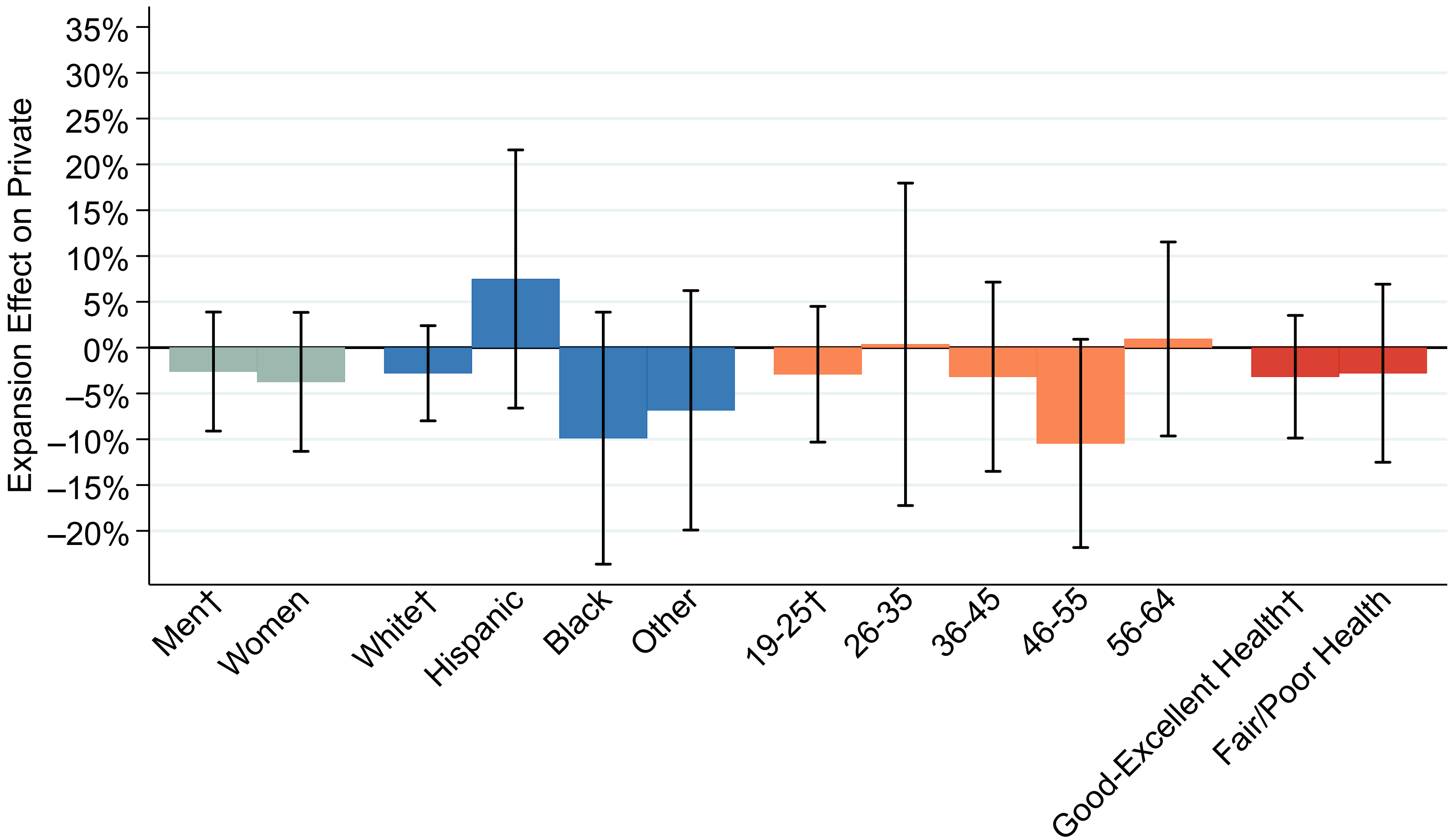
Figure A.1. Subgroup Effects on Probability of Any Insurance Coverage



NOTES: This figure shows the regression-adjusted differences-in-differences effects of ACA Medicaid expansion on any insurance coverage for subgroups. Base effects are estimated as a coefficient on a dummy variable equal to 1 in expansion states after the implementation date and equal to 0 otherwise. Interaction effects are estimated as a coefficient on interaction between the expansion dummy variable and a dummy variable for subgroup membership. Effects are reported in percentage points.

Effects are estimated using linear regressions controlling for gender, age, marital status, race, educational attainment, employment status, time (year-quarter) fixed effects, state fixed effects, and a linear time trend specific to expansion states.
† This indicates the base category in the regression model; other bars report the sum of base and interaction effects.
P-values for difference from base category effect are indicated as follows: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.
Error bars report 95-percent confidence intervals based on *t*-distribution with 37 degrees of freedom and standard errors clustered on state. Confidence intervals for interaction effects treat base coefficients as known.

Figure A.2. Subgroup Effects on Probability of Private Coverage



NOTES: This figure shows the regression-adjusted differences-in-differences effects of ACA Medicaid expansion on private insurance coverage for subgroups. Base effects are estimated as a coefficient on a dummy variable equal to 1 in expansion states after the implementation date and equal to 0 otherwise. Interaction effects are estimated as a coefficient on interaction between the expansion dummy variable and a dummy variable for subgroup membership. Effects are reported in percentage points.
Effects are estimated using linear regressions controlling for gender, age, marital status, race, educational attainment, employment status, time (year-quarter) fixed

effects, state fixed effects, and a linear time trend specific to expansion states.
† This indicates the base category in the regression model; other bars report the sum of base and interaction effects.
P-values for difference from base category effect are indicated as follows: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.
Error bars report 95-percent confidence intervals based on *t*-distribution with 37 degrees of freedom and standard errors clustered on state. Confidence intervals for interaction effects treat base coefficients as known.

Differences-in-Differences Effects of Medicaid Expansion for Wider Groups of States

In order to isolate a population of low-income childless adults with limited Medicaid eligibility prior to implementation of the 2014 Medicaid expansion, we excluded from our analysis 13 of the 27 states that had implemented the expansion by the end of 2014. To place our main coverage estimates and subgroup findings in context, Table A.7 reports estimates for several expanded samples that encompass these excluded expansion states.

Column 1 of Table A.7 reproduces our main estimates for reference, and Column 4 reproduces estimates for our main sample from a model that omits differential trends. Columns 2 and 5 of Table A.7 present estimates of our model with and without differential time trends for an expanded sample of 2014 expansion states that adds to the treatment group four expansion states that provided Medicaid eligibility to childless adults in 2013 but that imposed income limits below 100 percent of the FPL. We refer to these four states (Colorado, Connecticut, Minnesota, and New Jersey) as *partial expansion states*. We prefer to exclude them from our main specification because pre-2014 eligibility undermines the interpretation of our differences-in-differences model as capturing the effect of Medicaid expansion among the newly eligible. Adding these states to our main specification (Column 2) yields estimates that are close to the estimates for the 2014 expansion states. However, the effects of Medicaid expansion on Medicaid and private insurance are slightly lower, resulting in a lower estimate of the increase in overall insurance coverage that is no longer significant at the 5-percent level ($p = 0.073$).

However, because all four of these partial expansion states used early implementation of the ACA to expand Medicaid between 2010 and 2014, controlling for pre-2014 trends in a sample that includes these states seems inappropriate. Estimates that omit pre-2014 differential time trends were indistinguishable from estimates of the same specification for our sample of 2014 expansion states.

Columns 3 and 6 of Table A.7 expand the sample further to include all 50 states and the District of Columbia. The increase in Medicaid coverage is highly significant and is very close to the estimate for the 2014 expansion states in both the model with differential trends (13.1-percentage-point increase) and the model without differential trends (15.2-percentage-point increase). The change in private coverage becomes more negative (5.8-percentage-point decrease) and statistically significant ($p = 0.033$) when differential trends are included, but inclusion of differential trends is even less appropriate in this sample because this differential trend is likely to reflect coverage gains resulting from early implementation of the ACA, most notably in California (Sommers, Chua, et al., 2015a; Golberstein, Gonzales, and Sommers, 2015). Even if differential trends are omitted from the model, however, we find a marginally statistically significant ($p = 0.057$) reduction in private coverage (3.3-percentage-point reduction) associated with the 2014 Medicaid expansion when all expansion states are included in the sample.

Comparison of these results (Table A.7, Column 6) with the our main estimates for the 2014 expansion states (Table A.7, Column 4) and the sample including partial expansion states (Table A.7, Column 5) suggests strongly that this reduction in private coverage is driven by the nine expansion states where all adults with income below 100 percent of the FPL were Medicaid-eligible in 2013. Because this population was fully eligible prior to 2014, estimates for this group of states reflect welcome-mat effects. These findings warrant further investigation, but we did not focus on them in this study because we suspect that insurance coverage changes in 2014 among the previously Medicaid-eligible are likely to reflect fundamentally different mechanisms than coverage changes among the newly eligible. That is, we think that welcome-mat effects and take-up among the newly eligible should be studied separately. We are exploring these differences further in work in progress using the NHIS, but larger surveys, such as the CPS ASEC or the American Community Survey, may be better suited to examining state-specific differences among the low-income adult population. (We are also limited by the terms of our access agreement with NCHS, which does not provide for disclosure of state-specific estimates.)

Table A.7. Coverage Effects for Broader Groups of States

	(1)	(2)	(3)	(4)	(5)	(6)
Years	2009–2014	2009–2014	2009–2014	2009–2014	2009–2014	2009–2014
Main sample included*	Y	Y	Y	Y	Y	Y
Sample includes partial 2014 expansion states**		Y	Y		Y	Y
Sample includes all expansion states***			Y			Y
Control for covariates?	Y	Y	Y	Y	Y	Y
State group–specific time trends?	Linear	Linear	Linear			
Any insurance	0.089	0.066	0.044	0.149	0.138	0.104
	(0.041)	(0.036)	(0.037)	(0.027)	(0.024)	(0.026)
	0.034	0.073	0.238	0.000	0.000	0.000
Medicaid	0.126	0.118	0.131	0.154	0.142	0.152
	(0.027)	(0.026)	(0.028)	(0.023)	(0.022)	(0.024)
	0.000	0.000	0.000	0.000	0.000	0.000
Private insurance	-0.032	-0.040	-0.058	-0.006	0.001	-0.033
	(0.035)	(0.032)	(0.026)	(0.019)	(0.019)	(0.017)
	0.371	0.220	0.033	0.758	0.968	0.057
Non-Medicaid public insurance	0.006	-0.007	-0.024	0.014	0.004	-0.008
	(0.019)	(0.020)	(0.015)	(0.011)	(0.012)	(0.011)
	0.755	0.716	0.132	0.197	0.766	0.469
Number of states in model	38	42	51	38	42	51
N	11,584	†	16,907	11,584	†	16,907

* This sample was restricted to nonexpansion states and expansion states with no categorical eligiblity for

nondisabled childless adults in 2013.
** Partial 2014 expansion states had categorical eligibility for childless adults with an income limit below 100 percent of the FPL in 2013. These states are Colorado, Connecticut, Minnesota, and New Jersey.
*** This sample includes all expansion states.
† Sample size including partial expansion states was not released from the NCHS RDC.
SOURCE: 2009–2014 National Health Interview Survey.
NOTES: Each cell reports the differences-in-differences regression coefficient for the effect of the 2014 Medicaid expansion from a separate OLS regression model.
Standard errors clustered on state are in parentheses; two-sided *p*-values for equality to 0 are reported below standard errors. *P*-values were calculated based on a *t*-distribution with *G* minus 1 degree of freedom, where *G* is the number of clusters.
All samples consist of nondisabled childless adults aged 19–64 with annual family income at or below 100 percent of the FPL.
All regressions include state and quarter fixed effects. Covariates include age categories (19–25, 26–35, 36–45, 46–55, and 56–64), gender, binary marital status, four racial/ethnic groups (non-Hispanic white, Hispanic, non-Hispanic Black, and other), educational attainment (less than high school, high school diploma or GED, some college, college diploma, and more than college), and employment status (employed, unemployed, not in labor force, and unknown).

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