Health insurance transitions and use of fringe banks: Evidence from the Affordable Care Act

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Abstract
The Affordable Care Act (ACA) increased health insurance enrollment, potentially improving financial security. We test whether this insurance increase relates to changes in use of “fringe banks” (e.g., payday lenders, check cashers, and pawn shops). Using the panel structure of the Current Population Survey (CPS), we link 5 years of a Federal Deposit Insurance Corporation (FDIC)-sponsored supplement to March CPS data. We find declines in fringe bank use, specifically fringe credit (e.g., pawn loans), associated with new insurance coverage with larger declines for households affected by the ACA’s Medicaid expansion. These results suggest that health insurance reduces reliance on these controversial financial products.

KEYWORDS
ACA, fringe banks, health insurance, Medicaid, nonbank products, pawn loans, payday loans

JEL CLASSIFICATION
D12; D14; I13; J18

INTRODUCTION

Prior to the 2010 enactment of the patient protection and Affordable Care Act (ACA), nearly 50 million individuals lacked insurance and 29 million nonelderly adults were underinsured, meaning that their health insurance coverage did not adequately protect them from financial risk (Collins et al., 2015). Both uninsurance and underinsurance led to financial distress. Approximately 41% of working-age individuals struggled with medical bills, medical debt, or both and those in low-income households experienced this struggle to a greater extent (Doty et al., 2008). This lack of protection from financial risk threatened the ability of these households to withstand economic shocks, save, and make investments that promote upward economic mobility.

The ACA sought to improve protection from medical risk for low- and middle-income households by requiring health insurance for all Americans, as well as addressing insurance access and affordability. It expanded insurance access by encouraging (or requiring) employers to offer health insurance coverage, created new state-based private insurance markets with progressive subsidies, and offered new eligibility for public insurance to lower-income households.
households through Medicaid. These changes, in conjunction with the individual mandate, caused the uninsured rate to fall to a historic low level of 8.8% by 2016 (Barnett & Berchick, 2017).

The low- and middle-income households most likely to be uninsured before the ACA were also those most likely to be among the one-in-four households that utilized nonbank financial services in the past 12 months (Burhouse & Osaki, 2012). These “fringe banks,” nonbank providers of financial services, including payday lenders, pawn lenders, and check cashers, fill demand not being met by mainstream financial providers, but continued use could threaten long-term financial well-being. Utilizing nonbanks for financial services may inhibit savings; prevent the building of a credit record, often a necessary condition for employment, rental housing, and obtaining credit from banks; and, increase the price of available credit.

Yet, fringe banks are widely used due to ongoing financial insecurity for lower income Americans. An estimated 40% of Americans do not have enough savings to cover a $400 emergency expense, with higher rates among low- and middle-income households (Federal Reserve, 2019). Financial insecurity is related to out-of-pocket medical expenses. Approximately 20% of adults had unexpected medical bills of between $1,000 and $5,000 in the previous year and 40% of these have unpaid debt from these expenses; insured individuals are less likely to experience medical debt (Board of Governors 2019). If health insurance reduces reliance on fringe banks by protecting households from high medical expenses, then health insurance expansions may increase financial security and improve economic mobility.

This research examines if changes in health insurance coverage arising from the ACA affect the use of nonbank financial products. Our empirical approach employs 5 years of the Federal Deposit Insurance Corporation (FDIC)-sponsored Unbanked and Underbanked Supplement to the Current Population Survey (FDIC-CPS) linked across time to health insurance from the March CPS before and after the household was observed in the FDIC-CPS. This nationally representative data with both fringe bank use and changes in health insurance coverage, spans the period before and after the ACA's 2014 implementation of the individual mandate, the creation of state-based insurance exchanges, and the Medicaid expansion.

We use our novel dataset to first document several empirical facts. First, we replicate previous findings that income is negatively correlated with both use of fringe banks and lack of a bank account. In the pre-ACA period, income is positively correlated with health insurance enrollment; post-ACA, the relationship between income and health insurance is substantially flatter, particularly for individuals from households at the income-to-poverty levels targeted by the ACA. Second, we calculate health insurance transition rates to document that between 2013 and 2017, the likelihood that an individual was uninsured in two adjacent years fell from 10% to 4.6%. We also demonstrate that after the ACA, individuals were more likely to maintain consistent private or public health insurance. Finally, we verify that individuals most likely to gain health insurance after the ACA were more likely to come from lower-to-middle income households.

We then link changes in individual-level health insurance coverage induced by the ACA to changes in the use of fringe banks. We use two different empirical specifications, one that uses time variation from multiple ACA provisions, and the second that uses state and time-variation in enrollment due to the Medicaid expansion. We then examine correlations between individuals who gained health insurance in the past year and their reported use of fringe bank products. We find new health insurance enrollment after the ACA is associated with a 7.2 percentage point (24%) decrease in any fringe product use, and a 3.8 percentage point reduction (40%) in fringe credit use. Individuals newly enrolled in Medicaid expansion states reduce fringe bank use by 5.8 percentage points (18%), and fringe credit use by 5.3 percentage points (38%). Fringe credit decreases are driven by a 9.6 percentage point decline (59%) individuals at or below the Medicaid eligibility threshold in Medicaid expansion states.

Our study builds upon a growing literature that suggests that households financially benefit from health insurance enrollment. Health insurance may improve the financial stability and credit worthiness through reductions in out-of-pocket medical spending and medical debt, ultimately reducing bankruptcy filings, and improving credit scores (Brevoort et al., 2017; Caswell & Waidmann, 2019; Finkelstein et al., 2012; Golberstein et al., 2015; Gross & Notowidigdo, 2011; Hu et al., 2018; Mazumder & Miller, 2016; Miller et al., 2018). We build upon these findings, contributing new insight into the effect of health insurance on fringe bank products that are omitted from the existing literature but are key to understanding the economic well-being of households below median income. Together with A. Fitzpatrick and Fitzpatrick (2019), as well as work by Allen et al. (2017) on the effect of the California Medicaid expansion on use of payday loans, we find that health insurance reduces reliance on fringe banks. Health insurance may, therefore, potentially provide the newly insured a way to reduce their use of relatively expensive sources of credit and allow their financial behaviors to look more like that of middle-class Americans.

We make three main contributions to the literature. First, we extend the literature on the impact of health insurance on financial stability by examining the impact of health insurance enrollment on fringe bank use of low- and middle-income households. While these products are key to the financial lives of lower income Americans, there is less...
research on what factors drive demand for these products. Second, we contribute to the evaluation of the ACA, including ongoing state-level debates on the Medicaid expansion, and the larger debate over the role of public health insurance coverage in the United States. We show that the ACA improved financial outcomes of the newly insured by reducing their reliance on expensive sources of credit. Finally, our work highlights the strong correlation between health insurance and inclusion into the financial mainstream. We show that lower income households will enroll in health insurance if it is offered to them and change their financial behavior to look more like that of higher income households.

This paper proceeds as follows. Section 2 provides background on the ACA and non-bank financial products. Section 3 provides our conceptual framework. Section 3 presents the data and methods. Section 4 presents the results. Section 5 includes the discussion and conclusion.

2 | BACKGROUND

2.1 | The Affordable Care Act

The ACA was the largest health insurance reform since the creation of Medicare and Medicaid in 1965. The motivations for the ACA were twofold: (1) to move toward universal health insurance coverage and (2) to improve individual financial protections against medical risk, including indebtedness due to medical bills and forgoing care due to a lack of ability to pay (Obama, 2009). Struggles to afford health care or pay medical bills were related to being insured, as well as lack of adequate insurance (Doty et al., 2008). Reasons for lack of health insurance varied (Abraham et al., 2009; Levy & DeLeire, 2008). Voluntary employer-provided health insurance resulted in many employers not offering coverage, even to full-time workers. Without employer-provided coverage, high premiums and pre-existing conditions excluded individuals from individual market eligibility, if it was even available for purchase. In addition, many lower-income adults, particularly those without children, were ineligible for publicly provided health insurance through Medicaid.

Signed into law on March 23, 2010, the ACA sought to achieve universal health insurance coverage through individual and employer mandates, subsidies to purchase health insurance on newly created state-based insurance exchanges, and an expansion of Medicaid eligibility. The ACA also set minimum requirements for health insurance and made other regulatory changes to industry practices.

One of the largest expansions of access to low-cost health insurance coverage for lower income adults was the Medicaid expansion. Historically, Medicaid had federally determined minimum eligibility criteria based on income in relationship to the federal poverty level (FPL) and age. Prior to the ACA, states could and, often did, adopt more generous eligibility criteria, especially for children. But, before the ACA, no federal eligibility requirement existed for adults and few states provided coverage to adults, especially childless adults.

The implementation of the individual mandate, the Medicaid expansion, and the creation of state-based insurance markets for individuals with progressive subsidies began in 2014. The goal of each of these policies varied: the individual mandate was to expand the insurance pool, Medicaid expansion was to help individuals at or just above the poverty line (138% of FPL) gain public health insurance coverage, and households above Medicaid eligibility but up to middle income (138%–400% of FPL) were targeted progressive subsidies to purchase insurance in the individual market. The 2012 Supreme Court decision in National Federation of Independent Business v. Sebelius allowed states to opt-out of the expansion. Originally 26 states elected to expand their Medicaid programs; an additional 11 states have since implemented the expansion.

As intended, the ACA increased health insurance coverage. By 2017, approximately 23 million people gained coverage due to the ACA, including 9.5 million people through the Medicaid expansion (Carman et al., 2015). The largest increases in health insurance coverage occurred for individuals that were most likely to be uninsured prior to the ACA, including minorities, lower income households, those without a college degree, the young, single, and childless (Buchmueller et al., 2016; Kominski et al., 2017). These demographics overlap with populations most likely to use fringe bank products.

2.2 | “Fringe bank” products: Pawn loans, payday loans, and other nonbank financial products

Nonbank financial products include transaction products ranging from nonbank remittances, money orders, and check cashing services to credit products such as pawn loans, payday loans, auto-title loans, and rent-to-own (RTO) contracts.
Fringe bank products come with high costs, either through fees and/or interest. Fringe bank transaction products convert checks into cash or income into payments for a fee, while fringe bank credit provide short-term access to small-dollar loans. For example, the typical payday loan is a 2-week loan of $100 to $500 that, with $20 of fees per $100, imply an annual percentage rate of over 400%. Fringe bank credit has the potential to create debt or loss of an asset.

Households may change their demand for fringe bank products based upon unforeseen and urgent expenses, such as uninsured medical emergencies. While higher-income households can typically rely upon either savings in bank accounts or credit cards to meet these demands, lower-income households often lack or underutilize both bank accounts and mainstream credit options. Use of fringe banks is related to low or moderate income, unbanked status, education, and age; non-Whites, Hispanics, single adults, and those with children are more likely to use these products (Apaam et al., 2018; Barr, 2004; Caskey, 2002; McKernan et al., 2003; Zinman, 2010). The regulatory environment and local area characteristics influence the mix and type of fringe bank providers in an area (Fowler et al., 2014; Graves, 2003).

Despite their frequent use and need, the availability and use of fringe bank products remain controversial. A new literature empirically examines the relationship between fringe bank use and whether it improves economic well-being. This literature tends to focus on payday lending, measured by either access to or use of payday loans. It finds mixed effects of payday loans on economic financial and material well-being, with some work finding that these providers help households avoid financial distress while other work finds that increases bankruptcy filings and problems with housing and utility bills (Campbell et al., 2012; Fitzpatrick & Coleman-Jensen, 2014; Melzer, 2011; Melzer & Morgan, 2009; Morgan & Strain, 2008; Morse, 2011; Skiba & Tobacman, 2019; Zinman, 2010). Whether these products are affected by health insurance transitions is largely unknown, partly due to data limitations. For example, providers are not federally regulated, and do not report to credit bureaus, resulting in decentralized data on their availability and use.

Even if the empirical literature does not give a clear finding on these measures of well-being, using these nonbank products could inhibit opportunities to improve another measure of economic well-being: economic mobility. These providers typically do not report to credit bureaus so these households may not be able to create or improve their credit scores, a necessary condition to access lower cost mainstream credit. Additionally, owning and fully utilizing a bank account can improve the ability of households to save and access mainstream credit, which typically comes at a lower cost than fringe bank credit (K. Fitzpatrick, 2015a, 2015b). Thus, unbanked and underbanked households may face impediments to economic mobility due to fewer opportunities to save and make investments in education, entrepreneurship, and other goods.

3 | DATA AND METHODS

3.1 | Data

To examine the link between health insurance transitions and outcomes potentially correlated with economic mobility, we combine several waves of CPS supplements. The March Annual Social and Economic Supplement contains detailed information on the self-reported health insurance status of each individual in the household, as well as detailed household income. The FDIC-sponsored supplement to the CPS (FDIC-CPS) queries respondents on their household’s use of bank accounts and nonbank financial providers; the data provide a rare and detailed look at the financial services utilized by low-income households, as well as the frequency and reasons for use.

The CPS is a nationally representative monthly survey. Each wave contains a rotating panel (“rotation group”), where a specific housing unit is in the survey for 4 months, spends 8 months out of sample, before returning for 4 months. We analyze the effect of insurance transitions on fringe bank use with five waves (2009, 2011, 2013, 2015, and 2017) of the FDIC-sponsored supplement to the basic CPS linked to two waves of the March CPS (both the prior and current year of their FDIC-CPS participation).

One challenge is that the FDIC-CPS questions refer to household-level financial decisions, and insurance status can vary within household. Therefore, for each household in the matched FDIC-CPS, we select the most employed adult within the household to identify health insurance status. We then create that individual’s health insurance unit by assigning individuals and their dependents to families based on eligibility rules that relate to employer, individual market, and Medicaid coverage. We use the aggregate income of these health insurance units to approximate eligibility for Medicaid and or tax subsidies.
We further limit the sample to households most likely affected by the ACA by dropping households with all members age 65 or older from our sample. These households are likely insured through Medicare and thus unaffected by the ACA’s insurance expansions. In some specifications, we further limit the sample by creating “low income” subsamples with incomes below either 138% of FPL or below 400% of FPL. These subsamples are used to analyze how insurance transitions affect the financial behavior for individuals directly impacted by either the Medicaid expansion (below 138% of FPL) and those affected by either the Medicaid expansion or tax credits (below 400% of FPL).

Another challenge to our approach is that the March CPS redesigned both the income and health insurance questionnaire in 2014. First, we account for this by relying primarily on a sample not selected by income and limiting the sample based on age since older households were more affected by the income redesign (Semega et al., 2015). Second, the income redesign primarily affects our subsample analysis by potentially changing the composition of households in our low-income sample. We perform a robustness check of selecting on low educational attainment, rather than income. Income and education are highly correlated, but the educational attainment questions did not change. Results are robust to this alternative measure of exposure to the policy change.

The larger challenge is changes to the questions on health insurance status. In 2014, the Census Bureau implemented a redesign of the CPS health insurance questions due to concerns that the CPS question format overestimated the uninsured rate. To improve measurement, a redesign was implemented in 2014. This redesign primarily changed the time frame for the health insurance questions by querying the individual’s current health insurance coverage rather than asking about health insurance coverage in the last calendar year. Previous studies found that this change primarily affected aggregate estimates of insured rates, specifically employer-based health insurance coverage (Brault et al., 2014; Pascale et al., 2016). While Brault et al. (2014) find statistically significant differences in overall enrollment, as well as significant differences in Medicaid enrollment between the old and new question format, the magnitude of the changes is small (1–2 percentage points). Pascale et al. (2016) also find small (between 1 and 2 percentage points) but statistically insignificant differences in enrollment in the individual market or Medicaid between the old health insurance questions and the revised questions. For our analysis, the likely effect of the redesign would be to over-estimate the newly insured rate, resulting in smaller estimates on the impact of fringe bank use. In other words, by assigning new enrollment to those who were continuously enrolled, we would under-estimate the impact of new coverage on fringe bank use because these individuals would be less likely to change their financial behavior. To reduce potential bias, we incorporate year fixed effects in our specification. We conduct a robustness check by limiting the sample to households observed after the 2014 redesign and report these results in Appendix S1.

### 3.2 Measures

Our key variable of interest is the change in health insurance of the most employed adult in the household. We construct this measure by examining the most employed adult’s insurance coverage in the previous March CPS and comparing it to their insurance coverage in the subsequent March CPS. Our main measure is “newly insured” to denote adults that were uninsured in the previous March CPS but reported health insurance coverage in the subsequent March survey.

Our dependent variables measure use of fringe banks reported in the FDIC-CPS. Because the FDIC-CPS does not contain a consistent series of questions across years, we harmonize variables to form a consistent series over time. Our first measure is use of any fringe bank product over the previous 12 months, constructed by examining if anyone in the household reports using a nonbank money order, check casher, rent-to-own contract, pawn loan, or payday lender. While transaction products (nonbank money orders and check cashing) facilitate bill payment, credit products provide liquidity to households who lack sufficient savings. We hypothesize that credit products would be mostly affected by health insurance enrollment and measure use of fringe bank credit over the previous 12 months with an indicator of whether anyone in the household used a pawn loan, payday loans, and rent-to-own contract. We also consider fringe bank credit products separately, adding auto title loans as an additional outcome. Auto title loans are not included in our main measure because it was only first queried about in 2013.

Finally, we examine two measures of mainstream financial services: banking and credit card ownership. We first measure whether a household is currently unbanked and whether the household used a credit card in the previous 12 months, although the credit card measure is only available in 2015 and 2017. We hypothesize that insurance coverage may relax budget constraints and improve credit scores, both of which can facilitate bank account ownership and allow households to seek credit from lower cost, mainstream sources like credit cards (K. Fitzpatrick, 2015b).
3.3 | Summary statistics and descriptive characteristics

In Table 1, we present summary statistics of the sample overall, as well as disaggregated by time period. We begin with the characteristics of the most employed adult in the household, our focal individual for our analysis of changes in health insurance. While the majority of these individuals are White (69.7%), rather than African-American (11.4%) or Hispanic (12.9%), the sample becomes more non-White in the post-period. Other characteristics do not differ over time. The mean age is 45 years old. A small portion (14.4%) were born outside the United States. In terms of education, few (7.6%) have less than a high school degree, 27.7% have at least a high school degree and 29.0% have some college education but no degree. The remaining third (35.8%) of individuals have at least a college degree. The lower income sample (not shown), consists of more non-White, foreign born, younger, and less educated adults.

Moving to characteristics of the household in Panel B of Table 1, slightly more than a third (35.2%) of the sample are single adults without children, almost a third (28.7%) are married without children, and almost another third are married with children (27.1%). The remaining sample is in unmarried households with kids. Although there are not differences in household type across time, as households continued to recover from the Great Recession, mean real household income rises over time by $5,928 (not shown) and mean state unemployment rates falls by 1.7 percentage points. In terms of other characteristics of the household’s economic environment, 14.7% of our sample lives in a rural area and 27.0% lives in the central city, the remaining portion lives in a metropolitan area. Our sample becomes slightly more urban over time.

In Panel C of Table 1, we examine the financial services of interest. Overall, one-fifth (20.5%) of our sample had used any fringe bank product (rent-to-own contract, pawn loan, payday loan, nonbank money order, or check casher) in the previous 12 months. But, only 5% of the sample used a fringe bank credit source (rent-to-own contract, pawn loan, or payday loan) in the previous 12 months. Over time, use of both any fringe bank product and fringe bank credit decline over time, potentially due to the recovery from the Great Recession as well as increases in insurance coverage from the ACA.

Looking at specific financial services, few households report using any individual fringe bank product: RTO contract (1.4%), pawn loan (2.0%), or payday loan (2.6%). Small declines in payday loan use can be seen over time. Use of auto-title loans is also quite rare (1.1%) but small increases in use between 2013 and the post-period. The portion of unbanked households is 6.1% and does not change significantly over time. Credit cards, only available in the post-period, are owned by 71.2% of households. Households in the lower-income sample have slightly higher rates of fringe product use, but lower rates of credit card ownership (not shown).

Use of financial services, and health insurance enrollment rates, are closely related to income and poverty status over time. Figure 1 plots these three variables by the income of the household relative to the FPL in 2013 (the final year prior to the implementation of the ACA) and then again in 2017, the most recent data available. In 2013, there is a clear negative relationship between poverty level and health insurance enrollment rates. There is also a striking negative relationship between poverty level and unbanked status or use of fringe bank products. By 2017, the relationship between health insurance enrollment and poverty had weakened, particularly for the poorest households who were eligible for the Medicaid expansion (under 138% of FPL) and subsidies for health insurance in the individual market (100% of FPL – 400% of FPL). At the same time, unbanked status declined, with particularly large changes for the poorest households. Among households below 138% of FPL, bank account ownership rates increased by nearly 5 percentage points. Average fringe bank use also fell, but not as sharply.

Health insurance status is not static over time; insurance transitions are also common, particularly in the pre-period. Table 2 presents insurance transition probability matrices. In the final year prior to the enactment of the ACA, 9.6% of individuals who were uninsured in the previous year were still uninsured the following year. However, by 2017, only 4.4% of individuals who were uninsured in the 2016 March CPS were still uninsured in 2017. While part of this decrease is likely due to the ongoing recovery from the Great Recession, it is notable that public health insurance coverage also expanded. In both 2012 and 2013, 8.7% of individuals report public health insurance coverage, between 2016 and 2017 this figure rises to 12.6% of individuals.

Indeed, over time individuals with lower incomes, who are more likely to be eligible for Medicaid, have made up increasing shares of the newly insured population. As Figure 2 shows, in all years, adults in poorer households are more likely to report gains in health insurance coverage. However, in 2015—the first year of the ACA implementation in our data—there was a notable spike in health insurance enrollment among lower-income individuals. The highest rates of enrollment were among those eligible for Medicaid (i.e., under 138% of FPL) and those eligible for the largest
TABLE 1 Summary statistics of key variables

|                | Overall (1) | Pre-ACA (2009, 2011, 2013) (2) | Post-ACA (2015, 2017) (3) | Diff (4) |
|----------------|------------|-------------------------------|----------------------------|---------|
| **Panel A: Individual characteristics** |            |                               |                            |         |
| White (non-Hispanic) | 0.697 | 0.713 | 0.646 | 0.067*** |
| African-Americans | 0.114 | 0.108 | 0.131 | −0.023*** |
| Hispanic | 0.129 | 0.122 | 0.151 | −0.029** |
| Other race/ethnicity | 0.061 | 0.057 | 0.072 | −0.015* |
| Number of children | 0.672 | 0.673 | 0.669 | 0.004 |
| Age | 45.172 | 45.099 | 45.396 | −0.297 |
| Foreign born | 0.144 | 0.139 | 0.161 | −0.021 |
| Less than high school degree | 0.076 | 0.073 | 0.073 | 0.000 |
| High school degree | 0.277 | 0.271 | 0.271 | 0.000 |
| Some college | 0.290 | 0.291 | 0.291 | 0.000 |
| College graduate | 0.358 | 0.355 | 0.365 | −0.010 |
| **Panel B: Household characteristics** |            |                               |                            |         |
| Unmarried without kids | 0.352 | 0.351 | 0.356 | −0.005 |
| Unmarried with kids | 0.095 | 0.088 | 0.095 | −0.007 |
| Married without kids | 0.287 | 0.288 | 0.284 | 0.005 |
| Married with kids | 0.271 | 0.272 | 0.265 | 0.006 |
| Household size | 2.758 | 2.750 | 2.784 | −0.034 |
| Poverty level | 5.659 | 5.604 | 5.832 | −0.23 |
| State unemployment rate | 6.812 | 7.218 | 5.546 | 1.672*** |
| Rural | 0.147 | 0.163 | 0.097 | 0.066*** |
| Central city | 0.270 | 0.263 | 0.292 | −0.030** |
| Payday loan ban | 0.327 | 0.321 | 0.346 | −0.025 |
| **Panel C: Financial services** |            |                               |                            |         |
| Any use of fringe banks | 0.205 | 0.211 | 0.185 | 0.026*** |
| Any use of fringe bank credit | 0.052 | 0.054 | 0.044 | 0.010** |
| Rent-to-own contract | 0.014 | 0.015 | 0.014 | 0.001 |
| Pawn loan | 0.020 | 0.021 | 0.016 | 0.004 |
| Payday loans | 0.026 | 0.028 | 0.020 | 0.007* |
| Auto title loan | 0.011 | 0.007 | 0.013 | −0.005** |
| Credit card ownership | 0.712 | — | 0.712 | — |

Notes: Above are weighted averages of select variables from all years of data (column 1), pre-ACA (column 2), and post-ACA (column 3). The sample is the most employed adult in all households that completed the FDIC-CPS and could also be linked with the previous survey in March (n = 16,704). Next to column 3 is the t-test of differences between pre-ACA and post-ACA implementation, controlling for a state fixed effect. “Poverty level” is based on the most employed adult’s health insurance unit. See text for more details. Standard errors are clustered at the state level. Auto-title loans are only available in the 2013, 2015, and 2017 FDIC-CPS. Credit card ownership is only available in the 2015 and 2017 FDIC-CPS.

*p < .1; **p < .05; ***p < .01.

Subsidies in the individual market (i.e., between 100% and 200% of FPL). However, all those in households under 400% of FPL were particularly likely to report health insurance coverage gains in 2015. Gains were lower in 2017, and similar to that of 2013, partly because adults were more likely to retain insurance coverage from previous years as opposed to gain coverage.

To quantify the effect of income on the likelihood of becoming newly insured, we estimate a multivariate regression of the likelihood that an individual becomes newly enrolled in health insurance on a large dummy variable group of
categories of income as measured as a percent of the federal poverty line, controlling for factors correlated with income. Figure 3 plots point estimates on the coefficients from each of these income bins, where the omitted category is individuals in households above 600% FPL. In 2015 and 2017, individuals in households below 138% of FPL are 4.0 percentage points more likely to gain coverage than high income individuals, while individuals between 138 and 200%, and between 200% and 250% of FPL are 7.8–7.1 percentage points more likely to gain coverage. Individuals between 250% and 300% FPL are 3.9 percentage points more likely to gain coverage, while individuals 300%–350% FPL are 2.7 percentage points more likely to gain coverage. Individuals in households above 350% of FPL report similar new enrollment rates as the highest income category.

Now that we have established the relationship between household poverty level and changes in health insurance coverage, we turn to examining whether those who were newly enrolled into health insurance also report improved financial security.

**TABLE 2** Insurance transition probabilities by insurance category

|                | Uninsured, t + 1 | Private insured, t + 1 | Public insured, t + 1 |
|----------------|------------------|------------------------|-----------------------|
| Panel A: 2013  |                  |                        |                       |
| Uninsured, t   | 0.096            | 0.045                  | 0.017                 |
| Insured private, t | 0.046          | 0.680                  | 0.044                 |
| Insured public, t | 0.018           | 0.055                  | 0.087                 |
| Panel B: 2017  |                  |                        |                       |
| Uninsured t    | 0.044            | 0.036                  | 0.015                 |
| Insured private t | 0.030           | 0.711                  | 0.056                 |
| Insured public t | 0.013           | 0.074                  | 0.126                 |

Notes: Above are weighted averages of insurance transitions from the March-CPS 2012–2013 and March-CPS 2016–2017.
Regression approach

We test whether health insurance enrollment through the ACA is associated with increased financial security as measured by changes in fringe bank services. Because there were several channels through which individuals could have gained health insurance coverage, we consider two different approaches. First, we estimate whether post-ACA, newly enrolled individuals report changes in fringe banking use in the following difference-in-difference specification:

\[
\text{AnyFringeProduct}_{ist} = \alpha_0 + \alpha_1 \text{Newly Insured}_{ist} + \alpha_2 \text{Post2014}_{i} + \alpha_3 \text{Post2014}^* \text{Newly Insured} + \gamma_s + \eta_t + \delta X + \epsilon_{ist} \tag{1}
\]

where \(\text{AnyFringeProduct}\) for household \(i\) from state \(s\) at time \(t\) is an indicator for whether the household had used any fringe bank product (nonbank money order, check cashing service, pawn loan, payday loan, or RTO contract). We
include $\gamma$, a state fixed effect to account for time-invariant state-level characteristics, and $\eta$, a time fixed effect to account for any year-specific factors. Our coefficient of interest, $\alpha_3$, captures the total effect of ACA provisions related to the individual and employer mandate, the creation of the individual health insurance marketplace with progressive subsidies, the Medicaid expansion (in states that expanded), as well as the other policy changes in the ACA.

To account for confounding factors we include a vector of control variables describing the most employed adult in the household: White, Black, Hispanic (“other” is the omitted group); education (less than a high school education, high school degree, and some college; those with a college degree or more education serve as the omitted group); a continuous measure of age, as well as a set of indicators for age in 10-year age categories to capture cohort effects; the number of children of the adult that live in the household and the number of individuals that live in the individuals household; and, indicators for whether the adult was born in another country and whether the adult is a US citizen. We also include a vector of controls describing the household: unmarried without children; married without children; married with at least one child; married with at least one child. Finally, we capture spatial differences in the economic environment of the household: an indicator that the household lives in a rural area or a central city (with suburban area as the omitted group), the state’s unemployment rate in that year, and if the state prohibited payday lending. We cluster standard errors at the state level.

Equation (1) captures the total effect of various ACA provisions: mandates, individual market expansions with progressive subsidies, and the Medicaid expansion. However, one concern with the above regression is that we may not be adequately isolating the impact of health insurance enrollment due to exogenous factors. For example, over the post-2014 period, national economic conditions improved, variation that is likely endogenous to both health insurance enrollment status and fringe bank use. Therefore, we also consider a specification treating the Medicaid expansion as plausibly exogenous variation in health insurance enrollment. We estimate:

$$\text{FringeProduct}_{ist} = \beta_0 + \beta_1 \text{Newly Insured}_{ist} + \beta_2 \text{Medicaid Expansion}_{s} + \beta_3 \text{Medicaid Expansion} \ast \text{Newly Insured}_{ist} + \gamma_s + \eta_t + \delta X + \mu_{ist}$$

where all variables are the same as above except for the inclusion of MedicaidExpansion state instead of Post2014. MedicaidExpansion is a dummy variable that takes on a value of one if the state had implemented the Medicaid expansion by March of that year, when the March-CPS was fielded. In both regressions, our coefficient of interest is on the interaction term. This interaction compares newly insured individuals across time (Equation (1)) or by state of residence and time (Equation (2)). Equation (2) is our preferred specification because it utilizes state and time variation in the implementation of the Medicaid expansion, although it omits households responding to the creation of the individual health insurance market.

Several assumptions need to hold for our difference-in-difference estimates to produce unbiased results. Namely, states which chose to adopt the Medicaid expansion need to be an appropriate counterfactual for states that chose not to adopt the expansion. We test this assumption by examining whether the time path of nonexpansion states was similar to expansion states in the period prior to Medicaid expansion, the “parallel trends” assumption. These results, presented in Appendix S1, show that the parallel trends assumption is satisfied, and more generally the plausibility of this identification strategy. Still, concerns may remain of omitted variables bias. We acknowledge this as a potential limitation of assigning causality to our conclusions.

4 | RESULTS

In Table 3 we present estimates from our regression analysis. Panel A presents estimates from Equation (1). We find that post-2014, following the implementation of the ACA’s Medicaid expansion and the creation of state-based insurance exchanges for nongroup insurance coverage, those in households who were newly insured reduced their use of fringe credit products by 7.2 percentage points (24%), significant at the 5% level (column 1). We also find a significant decrease of 3.8 percentage points (70%) in use of fringe credit products among the newly insured after 2014 (column 2). Point estimates are negative, but insignificant, among households below 138% of FPL (columns 3 and 4) and below 400% of FPL (columns 5 and 6), possibly due to a much smaller sample size, some households falling in the “coverage gap” in states that did not expand Medicaid, or requirements to pay insurance premiums, copays, and/or deductibles for insurance.
Given that income may be measured with error we also examine a sample of households with low educational attainment. Among households without a college degree, newly insured individuals are 7.5 percentage points (23%) less likely to have used any fringe product in the past year, and 5.0 percentage points (45%) less likely to use fringe credit products. Both are significant at the 5% significance level.

In Panel B of Table 3 we present estimates from Equation (2) that utilize state and time variation in the implementation of the Medicaid expansion. We find a similar pattern of point estimates as in Panel A, although our estimates are slightly larger and more precisely estimated. Among the full sample, the point estimate for the newly insured in Medicaid expansion states is a 5.8 percentage point (18%) decline in fringe bank use and a statistically significant 5.3 percentage point (38%) decline in fringe credit use. Among newly insured, households that would be eligible for Medicaid

| TABLE 3 | Regression estimates for the relationship between health insurance transitions and any use of fringe banks |
| --- | --- |
| | Full sample | Below 138% FPL | Below 400% FPL | Without a college degree |
| Any fringe product | Any fringe credit | Any fringe product | Any fringe credit | Any fringe product | Any fringe credit | Any fringe product | Any fringe credit |
| (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| Newly insured | 0.037** | 0.027** | 0.030 | 0.018 | 0.019 | 0.029** | 0.027 | 0.032** |
| | (0.017) | (0.011) | (0.055) | (0.030) | (0.020) | (0.013) | (0.019) | (0.013) |
| Post-2014 | −0.029** | −0.021*** | −0.085** | −0.031 | −0.051** | −0.027* | −0.021 | −0.017** |
| | (0.012) | (0.007) | (0.036) | (0.029) | (0.021) | (0.014) | (0.013) | (0.008) |
| Newly insured * post-2014 | −0.072** | −0.038* | −0.035 | −0.045 | −0.060 | −0.039 | −0.075** | −0.051** |
| | (0.031) | (0.022) | (0.080) | (0.045) | (0.039) | (0.027) | (0.037) | (0.025) |
| Observations | 16,358 | 16,358 | 2,476 | 2,476 | 8,261 | 8,261 | 14,337 | 14,337 |
| R-squared | 0.103 | 0.051 | 0.093 | 0.069 | 0.085 | 0.046 | 0.100 | 0.050 |
| Mean of dependent variable in pre-period for newly insured | 0.299 | 0.095 | 0.425 | 0.111 | 0.333 | 0.113 | 0.328 | 0.114 |

Panel B: Interaction with Medicaid expansion

| Newly insured | 0.026* | 0.025** | 0.015 | 0.019 | 0.008 | 0.027** | 0.027 | 0.030** |
| | (0.016) | (0.010) | (0.051) | (0.029) | (0.018) | (0.013) | (0.019) | (0.012) |
| Medicaid expansion state | 0.005 | −0.000 | −0.022 | 0.033 | −0.002 | 0.001 | −0.021 | −0.000 |
| | (0.015) | (0.007) | (0.050) | (0.031) | (0.028) | (0.014) | (0.013) | (0.009) |
| Newly insured * Medicaid expansion state | −0.058 | −0.053*** | 0.033 | −0.096* | −0.036 | −0.059** | −0.075** | −0.080*** |
| | (0.041) | (0.020) | (0.124) | (0.056) | (0.052) | (0.027) | (0.037) | (0.020) |
| Observations | 16,358 | 16,358 | 2,476 | 2,476 | 8,261 | 8,261 | 14,337 | 14,337 |
| R-squared | 0.102 | 0.051 | 0.093 | 0.070 | 0.085 | 0.046 | 0.063 | 0.051 |
| Mean of dependent variable, prior to ACA, for newly insured in states that never expanded | 0.317 | 0.138 | 0.431 | 0.164 | 0.339 | 0.154 | 0.331 | 0.155 |

Notes: Above are linear probability models where the outcome is whether or not the respondent had used (even columns) any fringe credit product in the past 12 months (rent-to-own contract, pawn loan, or payday loan) or (odd columns) (credit products or rent-to-own contracts, pawn loan, payday loan, nonbank money order or check cashing outlets) in the past 12 months. The sample is all households in the merged FDIC-CPS/March CPS file (columns 1 and 2); all households where the most employed adult is in a health insurance unit at or below 138% of the FPL (columns 3 and 4); all households where the most employed adult is in a health insurance unit at or below 400% of FPL (columns 5 and 6); and, all households where the most employed adult has less than a college degree (columns 7 and 8). “Newly insured” means that most employed adult reported gaining health insurance coverage between the previous March CPS and the following March CPS. All regressions are weighted and include state and time fixed effects, as well as controls for race, age, foreign born status, and education (for most employed adult), household composition, city/suburban/rural status, state-level unemployment rate, and whether there was a payday loan ban in that state-year. Standard errors clustered at the state level.

*p < .1; **p < .05; ***p < .01.
TABLE 4  Regression estimates for the relationship between health insurance transitions and use of fringe bank credit products

|                      | Rent-to-own (RTO) contract | Pawn loan | Payday loan | Auto title loan | Credit card | Unbanked |
|----------------------|-----------------------------|-----------|-------------|-----------------|-------------|----------|
| Panel A: Full sample |                             |           |             |                 |             |          |
| Newly insured        | 0.003                       | 0.007     | 0.018**     | −0.007          | −0.149***   | 0.025    |
|                      | (0.006)                     | (0.007)   | (0.007)     | (0.005)         | (0.046)     | (0.019)  |
| Medicaid expansion state | 0.002                       | −0.004    | 0.001       | −0.002          | −0.010      | −0.006   |
|                      | (0.005)                     | (0.006)   | (0.007)     | (0.005)         | (0.024)     | (0.014)  |
| Newly insured * Medicaid expansion state | −0.014                     | −0.021**  | −0.021      | 0.015           | 0.109       | 0.021    |
|                      | (0.013)                     | (0.010)   | (0.018)     | (0.010)         | (0.076)     | (0.040)  |
| Observations         | 16,358                      | 16,358    | 16,358      | 6,437           | 3,620       | 16,358   |
| R-squared            | 0.024                       | 0.022     | 0.033       | 0.013           | 0.208       | 0.136    |
| Mean of dependent variable, prior to ACA, for newly insured in states that never expanded | 0.023 | 0.046 | 0.093 | 0.017 | 0.426 | 0.130 |
| Panel B: Below 138% FPL |                             |           |             |                 |             |          |
| Newly insured        | 0.003                       | 0.007     | 0.018**     | −0.013          | −0.190      | 0.039    |
|                      | (0.006)                     | (0.007)   | (0.007)     | (0.012)         | (0.127)     | (0.054)  |
| Medicaid expansion state | 0.002                       | −0.004    | 0.001       | −0.001          | −0.147      | −0.014   |
|                      | (0.005)                     | (0.006)   | (0.007)     | (0.006)         | (0.107)     | (0.051)  |
| Newly insured * Medicaid expansion state | −0.014                     | −0.021**  | −0.021      | 0.031           | 0.090       | 0.103    |
|                      | (0.013)                     | (0.010)   | (0.018)     | (0.019)         | (0.160)     | (0.159)  |
| Observations         | 2,476                       | 2,476     | 2,476       | 1,043           | 630         | 2,476    |
| R-squared            | 0.054                       | 0.047     | 0.064       | 0.175           | 0.324       | 0.180    |
| Mean of dependent variable, prior to ACA, for newly insured in states that never expanded | 0.020 | 0.053 | 0.109 | 0.012 | 0.344 | 0.287 |
| Panel C: Below 400% FPL |                             |           |             |                 |             |          |
| Newly insured        | 0.005                       | 0.007     | 0.019**     | −0.007          | −0.144**    | 0.022    |
|                      | (0.008)                     | (0.009)   | (0.008)     | (0.007)         | (0.057)     | (0.025)  |
| Medicaid expansion state | 0.003                       | −0.008    | 0.008       | −0.001          | 0.003       | −0.018   |
|                      | (0.010)                     | (0.011)   | (0.011)     | (0.005)         | (0.033)     | (0.027)  |
| Newly insured * Medicaid expansion state | −0.018                     | −0.024*   | −0.022      | 0.021           | 0.086       | 0.055    |
|                      | (0.019)                     | (0.013)   | (0.024)     | (0.013)         | (0.104)     | (0.055)  |
| Observations         | 8,261                       | 8,261     | 8,261       | 3,159           | 1,819       | 8,261    |
| R-squared            | 0.026                       | 0.024     | 0.037       | 0.031           | 0.201       | 0.132    |
| Mean of dependent variable, prior to ACA, for newly insured in states that never expanded | 0.030 | 0.061 | 0.094 | 0.023 | 0.493 | 0.164 |

Notes: Above are linear probability models where the outcome is whether or not the respondent had used a specific financial product in the past 12 months. The exception is credit cards, which refer to ownership. The sample is all households in the merged FDIC-CPS/March CPS file (Panel A); all households where the most employed adult is in a health insurance unit at or below 138% of the FPL (Panel B); and, all households where the most employed adult is in a health insurance unit at or below 400% of FPL (Panel C). “Newly insured” means that most employed adult reported gaining health insurance coverage between the previous March CPS and the following March CPS. The mean of credit card ownership is only calculated among the newly insured in nonexpansion states overall because the variable is not available before the ACA. The mean for auto title loans is calculated among nonexpansion states overall due to small sample size. All regressions are weighted and include state and time fixed effects, as well as controls for race, age, foreign-born status, number of children, and education (for most employed adult), household composition, city/suburban/rural status, state-level unemployment rate, and whether there was a payday loan ban in that state-year. Standard errors clustered at the state level.

*p < .1; **p < .05; ***p < .01.
coverage in Medicaid expansion states (columns 3 and 4), there is no significant decline in fringe product use but a 9.6 percentage point (59%) decline in fringe credit use. For households below 400% of FPL—that is, eligible for Medicaid or tax subsidies for health insurance—there is a 5.9 percentage point (38%) decline in fringe credit use. Finally, in columns 7 and 8, the sample of households without a college degree, shows a 7.5 percentage point (23%) decline in fringe bank use and an 8.0 percentage point (52%) decline in fringe bank credit related to the Medicaid expansion.\textsuperscript{16}

We check to ensure that the redesign of the insurance questions does not bias our results by dropping observations before the 2014 redesign and estimating Equation (2).\textsuperscript{17} Estimates are imprecise, possibly due to small sample sizes. The newly insured in Medicaid expansion states increase fringe bank use, consistent with research suggesting that “late expander” states tended to provide less generous Medicaid coverage, by imposing more cost sharing and other requirements on recipients (Brooks et al., 2019). The pattern of results for fringe bank credit, however, is similar to our main estimates. Point estimates suggest that newly insured households in states that expanded Medicaid coverage are less likely to use fringe bank credit; point estimates are largest for those with income levels targeted by the Medicaid expansion. Results are in Appendix S1.

These are large and economically meaningful relationships. The use of financial services by households at the lowest income levels, who were those most likely to use fringe bank products, demonstrated the largest changes contemporaneously with the implementation of major ACA provisions. For example, prior to the ACA, 39% of households from below 138% of FPL used fringe bank products (see Figure 1) and 9.9% of these households had used fringe bank credit. Moreover, comparing Panels A and B gives us insight into the relative impacts of general health insurance enrollment and Medicaid enrollment. Our results indicate that the “total” relationship between the ACA and fringe bank use is similar to the relationship between the Medicaid expansion and fringe bank use. This may be because the population affected by the Medicaid expansion were those most likely to change their use of fringe banks. It may also be because even after the ACA, many individuals above Medicaid eligibility continue to struggle with medical bills.

To further explore these results, we re-estimate Equation (2) for the following outcomes: each type of fringe bank credit product, unbanked status, and whether the household owns a credit card in the past year. We choose Equation (2) because it exploits state and time variation that allows us to estimate the auto title loan use and credit card ownership variables that are only available in later periods.

Table 4 presents the results. Columns 1 through 3 include the three fringe bank credit products in our main measure: RTOs, pawn loans, and payday loans. All point estimates are negative, but we only find a statistically significant decline (2.1 percentage point or 43.5%) for pawn loans. In Column 4, we examine auto title loans, a fringe bank product only available in the FDIC-CPS since 2013, and find a positive but statistically insignificant point estimate. We also find a large but insignificant point estimates for credit card ownership, a measure only available since 2015. Finally, we examine unbanked status and find small and positive but insignificant relationships the portion of unbanked households. Estimates are generally similar among sample of households at 138% FPL (Panel A) and 400% FPL (Panel B).

Together with Table 3, estimates from Table 4 suggest that newly insured individuals affected by the Medicaid expansion specifically report declines in fringe bank credit. One mechanism could be that the Medicaid expansion reduced medical expenses and debt, contributing to reduced demand for fringe bank credit. This is consistent with Skiba and Tobacman (2019) who find that most borrowers used payday loans because they could not borrow from mainstream credit sources.

### 5 Discussion and Conclusion

We estimate the relationship between the ACA, particularly the Medicaid expansion, and the use of fringe banks, including pawn shops and payday lenders. Our novel and detailed dataset allows us to examine how health insurance transitions affect fringe bank use. We find that insurance coverage is associated with reductions in demand for fringe bank credit, especially pawn loans. We demonstrate that health insurance may be an important method to discourage lower income households from using higher-cost sources of credit, potentially helping lower-income households engage in financial behaviors similar to middle class Americans.

By 2017, Medicaid covered 22% of all nonelderly adults—many who may be unable to afford other health insurance—and paid for nearly one-sixth of national spending on health care (Centers for Medicare and Medicaid Services, 2018). Quantifying whether and to the degree to which these individuals are better off compared to alternative insurance arrangements or remaining uninsured is important for policymakers and so this research is timely. We show that expanding public health insurance can improve the financial well-being of lower income households. Changes to
the ACA, as well as state and federal changes to Medicaid, continue to be debated. When considering options ranging from restructuring Medicaid as a block-grant program to implementing the Medicaid expansion in states that have not yet chosen to do so, policymakers should consider the spillovers into other policy areas.

One such area is consumer financial protection. The federal government is emphasizing financial deregulation, including large changes to the Consumer Financial Protection Bureau (CFPB), which may increase the availability of fringe bank products. If Medicaid coverage reduces demand for these products, it could reduce the harms of financial deregulation that many consumer advocates fear. Moreover, because many households above Medicaid eligibility still struggle with medical bills, expanding adequate health insurance may be an effective policy to improve financial outcomes for low- and middle-income households by encouraging less reliance on expensive sources of credit.

Health insurance may also encourage more participation in the mainstream financial system and, over time, increase financial security and economic mobility. Our results indicate that health insurance expansions could promote pathways to the middle class by reducing household reliance on controversial fringe bank products, and ultimately encourage a more inclusive financial system that enhances long-term well-being.

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ENDNOTES
1 Individuals who earn between 100% and 400% of FPL and are ineligible for Medicaid can receive federal tax credits for assistance with the purchase of health insurance in the individual market. With these subsidies, premiums are limited to a fixed percentage of the recipient’s income.
2 Despite the creation of the CFPB and federal regulation of the banking industry, relatively few federal limits exist. Regulation of these products varies widely with some states choosing to tightly regulate or outright ban these products and others enacting little to no regulation. Still, consumers can travel across state lines or use online providers to access these products. While state regulation reduces fringe bank use, they are often not binding.
3 In contrast to traditional longitudinal panels, the CPS randomly samples housing units, not individuals; therefore, the individuals and households completing the survey may change across time for the same identification number. We present analysis in this paper from all observations. We also check the fraction observations determined to be a “match” according to the procedures defined in Madrian and Lefgren (2000). The ID is defined to be a “match” if the individual’s sex and age (with a 1-year allowance) differ and then either race matched or educational attainment matched (with an allowance to account for individuals graduating). Between March and March, 97.2% of all observations are a match. Results are largely similar when we restrict the sample to those who are likely matches (not shown).
4 The 2009 FDIC-CPS was in January while FDIC-CPS for all other years occurred in June. A total of 58,376 household observations were successfully merged across the FDIC-CPS and the March CPS.
5 To create these units, we adapted State Health Access Data Assistance Center sample code. Our focus on the most employed adult results in only one health insurance unit considered per household. For simplicity in the exposition, we use the term household rather than health insurance unit in the text.
6 Results are robust to including these health insurance units in the sample.
7 See Appendix S1. Over the four survey waves, the FDIC-CPS has modified their questions. In particular, in 2009 the survey asked about whether the household had ever used a nonbank check casher, nonbank money order, pawn lender, or rent-to-own agreement, and then asked about the frequency of use (“at least a few times a year; once or twice a year; almost never”). In contrast, in all subsequent surveys the survey specifically inquired about alternative financial service use in the past 12 months. For the 2009 responses, we treat households that report using the AFS product “once or twice a year” or “at least a few times a year” as having used the product in the past 12 months. In contrast, for payday loans in 2009, the survey first asked if anyone ever used a payday loan and then queried the number of times in past 12 months anyone in the household used a payday loan. For 2009, we determine payday loan use in by the number of payday loans used in the past 12 months.
8 We omit tax refund anticipation loans (RALs) because rather than a regular source of liquidity, these products are only available once a year with income tax filing. Additionally, the 2009 survey asked about RAL use in the past 5 years while 2011, 2013, and 2015 surveys asked about the past 12 months. When we include RALs in our measure by treating households that used a RAL in the past 5 years as having used one in the past 12 months, our results do not materially change.
9 The relationship between income and new health insurance enrollment is approximately the same in 2009, and 2011 as in 2013, so those years have been excluded for clarity.
The Medicaid expansion primarily affected individuals in households at or below 138% (133% with a 5% income disregard) FPL. Many individuals also became enrolled through the health insurance exchanges, which offered subsidies for families up to 400% of the FPL.

11 We estimate: \( \text{NewlyInsured}_{it} = \delta_0 + \sum_{i=1}^{10} \delta_i \text{Poverty Category}_i + \gamma_X + \eta_t + \delta_X + \epsilon_{it} \).

12 States that implemented the Medicaid expansion by March 2017 are Alaska, Arizona, Arkansas, California, Colorado, Connecticut, Delaware, District of Columbia, Hawaii, Illinois, Indiana, Iowa, Kentucky, Louisiana, Maine, Maryland, Massachusetts, Michigan, Minnesota, Montana, Nevada, New Hampshire, New Jersey, New Mexico, New York, North Dakota, Ohio, Oregon, Pennsylvania, Rhode Island, Vermont, Washington, and West Virginia. Coverage under the Medicaid expansion became effective January 1, 2014 in all states that have adopted the Medicaid expansion except for the following: Michigan (4/2014), New Hampshire (9/2014), Pennsylvania (1/2015), Indiana (2/2015), Alaska (9/2015), Montana (1/2016), and Louisiana (7/2016). Virginia did not implement expansion until 1/2019. Idaho, Nebraska, and Utah adopted but have not implemented the expansion as of 6/2019.

13 The coverage gap refers to low-income individuals in states that did not expand Medicaid. Some of these individuals may be ineligible for traditional Medicaid in their state but not have access to Marketplace subsidies because their income is below the poverty level.

14 Appendix S1 provides estimates for households with a high school degree or less.

15 The point estimate for any fringe bank use just misses conventional significance levels at \( p = .165 \).

16 Households with a high school degree or less educational attainment are in Appendix S1.

17 We do not estimate Equation 1 because we have no preperiod data.

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**SUPPORTING INFORMATION**

Additional supporting information may be found online in the Supporting Information section at the end of this article.

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