

Offsetting Policy Feedback Effects: Evidence from the Affordable Care Act *

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Abstract

Can politicized, market-based policies generate policy feedbacks? The Affordable Care Act (ACA) is complex, relying partly on new exchanges, subsidies, and penalties to increase health insurance enrollment. Yet prior research focuses on changes to Medicaid, a pre-existing program. We employ several data sets—including a rolling cross-sectional survey (n=117,000) and a novel, population-based panel—to assess the impacts of the exchanges and individual mandate on public opinion. Descriptive analyses demonstrate that exchange users are more supportive of the ACA. However, subsequent analyses using various inferential strategies indicate that the exchanges' overall causal effects were limited. These limited feedback effects are due to heterogeneity, not invisibility: exchange users connected positive experiences with ACA opinions, but these were offset by price spikes and concentrated, negative attitude changes among the uninsured.

Political scientists have long documented the varied ways public policies can reconfigure the political landscape at the elite and mass levels (Patashnik, 2008; Campbell, 2012). While policies that are directly implemented by governments commonly generate feedback effects on public opinion (Soss, 1999; Campbell, 2003; Mettler, 2005; Lerman and Weaver, 2014; Nall, 2018), policies that rely on indirect mechanisms such as private markets or the tax code often do not (Howard, 1999; Mettler, 2011; Morgan and Campbell, 2011; Galvin and Thurston, 2017; Mettler, 2018). In part, that is because policies which are provided through private actors or the tax code are often lower in visibility and less easily politicized (Howard, 1999; Mettler, 2011). These claims are bolstered by a separate body of research concluding that self-interest typically has limited effects on public opinion (Citrin and Green, 1990).

Would market-based policies produce feedback effects on mass opinion if they were more salient? Or would indirect provision remain too difficult to attribute to government-provided benefits, especially within a highly politicized and polarized context?

We consider these questions within the context of the 2010 Affordable Care Act (ACA). The ACA was widely considered the most important American social policy reform in decades. Its primary goal was to expand access to health insurance, as nearly 50 million people lacked insurance upon enactment (Kaiser Health News, 2012). Unlike certain existing programs such as traditional Medicare, key elements of the ACA were designed to act principally through market-based mechanisms. Its centerpiece was the creation of exchanges on which Americans could purchase private health insurance. Initially, 75% of the total gains in health insurance enrollment were expected to come from the exchanges (Congressional Budget Office, 2010). The exchanges were potentially pivotal not only in expanding access to insurance but also in broadening the ACA's political support. The exchanges were designed to serve a large and middle-class constituency, which when combined with the Medicaid expansion offered the possibility of fostering a cross-class coalition in support of the law.

The ACA is similar to previously studied market-based policies in that it relies on incremental policy levers, many of which are individually low-salience measures affecting small and distinct sub-populations (Patashnik and Oberlander, 2018). Yet, unlike other market-based policies, the ACA was an extraordinarily salient and politicized initiative that has been at the forefront of political debates since it was drafted (Hopkins, 2018). The ACA also had profound impacts

on many Americans that were unlikely to go unnoticed: comprehensive health insurance is a substantial benefit that significantly improves individuals' financial security, health care access, and recipients' mental health (Finkelstein et al., 2012; Sommers, Gawande and Baicker, 2017). The law employed sticks as well as carrots, as it entailed substantial new taxes and regulation (Jacobs and Mettler, 2018) that might be viewed favorably or unfavorably depending on personal circumstances. It thus has the potential to generate *negative* policy feedbacks as well, especially if the offerings on the exchanges were unsatisfactory or disappointing. Whether positively or negatively, this combination of political visibility and real-world impact may make the ACA's exchanges more influential than other market-based policies. Yet to date, research on the ACA's impacts has focused primarily on its Medicaid expansion (Clinton and Sances, 2017; Haselswerdt, 2017; Hopkins and Parish, 2019) or its overall impacts (McCabe, 2015; Jacobs and Mettler, 2016, 2018) while research isolating the exchanges' impacts has been quite limited (but see Hosek, 2016).¹

The paper's next two sections develop these hypotheses, first briefly sketching how the exchanges operate and then using prior research to explain our hypotheses. It subsequently turns to three sections which evaluate these conflicting predictions empirically. Our tests exploit various population-based data sets to estimate the impact of the exchanges' implementation on the ACA attitudes of those most likely to use them. Specifically, we focus on those who purchase their own insurance (whether via the exchanges or not) or who lack insurance.

Our initial results are descriptive and set the stage for subsequent causal inferences. Those who purchase their own insurance became more favorable at precisely the moment when the exchanges opened. At approximately the same time, those without insurance became less favorable toward the ACA. To account for the heterogeneity of experiences on the exchanges, we also provide descriptive results from separate KFF surveys of the non-group insured and Kentucky residents in the SI.² Those using the exchanges who received subsidies felt more positively toward the law.

Do the cross-sectional differences reflect causal effects, or are they instead the products of se-

¹Existing research has also considered exchange enrollment as a *dependent* variable, demonstrating that Republicans are less likely to use the exchanges (Lerman, Sadin and Trachtman, 2017; Sances and Clinton, 2019).

²Kentucky saw the largest initial gains in insurance post-ACA (Hopkins and Parish, 2019), making it a valuable state to analyze.

lection into different types of insurance? Building on the call in Campbell (2012) to conduct more policy feedbacks research focusing on causal inference, one central contribution of this manuscript is to provide estimates of the exchanges that are credibly causal effects and not subject to alternative explanations based on selection bias. Rather than relying on a single, decisive test, we assess the impact of the exchanges through several analyses.

In one, we make use of over-time variation. Evidence from a novel, population-based panel uncovers no evidence that exchange users overall became disproportionately more positive toward the ACA after implementation. However, the panel evidence illustrates that those who become uninsured also become dramatically less supportive of the ACA.

Next, to account for selection into purchasing on the exchanges, we consider whether those demographically most likely to use the exchanges became more favorable toward the ACA after implementation. They do not. And we further find that those most likely to be uninsured were significantly *less* favorable toward the ACA—and even less likely to identify as Democrats—after its implementation.

Among its regulations, the ACA limited the premiums that insurers could charge older respondents, making those in their early 60s a most-likely case for positive feedback effects. A regression discontinuity analysis demonstrates positive changes in attitudes among those with newly capped premiums. The ACA helped close the previously observed gap between 64-year-olds (who typically rely on private insurance) and 65-year-olds (who often use Medicare) (Lerman and McCabe, 2017). Finally, using administrative data on exchange pricing and geo-coded HTS respondents, we find that citizens who purchase insurance via the exchanges become more negative when prices on their local exchanges spike.

Overall, these results suggest that Americans shifted their attitudes and even partisanship in response to perceived benefits or costs from the ACA. Feedback effects, however, partially cancelled out in aggregate, due to concentrated perceived costs and backlash among those likely to go without insurance. It's not that the ACA's market-oriented features were politically invisible. Instead, their effects were heterogeneous, varying across places and people.

Background on the ACA's Exchanges

Some social policies can be summarized succinctly—the ACA cannot be. It is thus valuable to provide background on the ACA's design and goals. Doing so enables us to develop targeted hypotheses about its potential feedback effects, contextualize and justify potential estimation strategies, and underscore the law's complexity.

The pre-ACA system of health insurance provision was already a patchwork of public and private insurance, with private insurance typically tax-subsidized and provided primarily through employers (Hacker, 2002). While there was a separate insurance market for individuals, insurers could reject applicants with pre-existing conditions, and the individual market was at once small and highly variable in the products available (Jones, 2017). On the public side, Medicare was a federal program that insured a majority of those over 64, while Medicaid was a means-tested federal-state partnership targeting households with children and varying markedly across states (Michener, 2018).

Rather than develop or expand a single program, the ACA sought to fill in gaps in the existing system. The law's centerpiece was arguably the establishment of exchanges on which individuals could purchase private insurance. Insurers were only allowed to charge older customers three times as much as younger customers, and they could no longer deny coverage due to health histories. Initial projections indicated that by 2019, 24 million Americans would be insured through the exchanges and 16 million through the Medicaid expansion (Congressional Budget Office, 2010). As enacted, the law also expanded Medicaid to 138% of the federal poverty line and allowed adults to participate irrespective of family status. Separate elements ended lifetime caps on insurance company payments, defined essential health benefits plans must include, and enabled children to remain on their parents' insurance until 26.

The exchanges themselves were comprised of various policy levers. As enacted, the exchanges were bolstered by a new tax penalty or "individual mandate" for Americans who didn't have qualifying health insurance, a mandate intended to limit adverse selection. In 2016, 5 million tax returns made such payments, with a mean payment of \$727 (Internal Revenue Service, 2019). The federal exchanges also had a notoriously rocky roll-out. The healthcare.gov website was plagued by long waiting times and other technical difficulties during its fall 2013 opening.

By contrast, for Americans making below 400% of the poverty line, the federal government provided subsidies in the form of the Advanced Premium Tax Credit (APTC). In 2016, 6.1 million American households received this credit, with a mean subsidy of approximately \$4,000 (Internal Revenue Service, 2019).³

Prior Research and Hypotheses

Research on policy feedbacks is well established (Campbell, 2012). The developed state of prior research enables us to begin by stating hypotheses on the ACA's exchanges specifically.

One study which considers whether market-based policy levers can produce mass-level feedback effects is Morgan and Campbell's 2011 analysis of the 2003 Medicare Modernization Act (MMA). Like the ACA, the MMA was a health policy reform that relied substantially on market-based mechanisms, though the MMA sought to add a prescription drug benefit to Medicare. Partly using panel data, Morgan and Campbell show that the MMA's impacts on public opinion were limited: its implementation did not lead beneficiaries to support more market-based policies generally, nor did it improve perceptions of Republicans' handling of health policy.

Is the imprint of the ACA's exchanges on public opinion likely to differ from the MMA's? Although the ACA's exchanges rely on private insurers to deliver a government-defined benefit, the ACA diverges from the MMA in its political profile and benefits provided. The law overall was the subject of sustained, highly salient political battles that lasted for years after implementation, a fact which may heighten its salience (see also Gollust et al., 2014; Fowler et al., 2017; Hopkins, 2018). And prior research provides extensive evidence that acquiring health insurance can shape ACA attitudes, whether through Medicare (Lerman and McCabe, 2017), the ACA overall (McCabe, 2015; Jacobs and Mettler, 2016, 2018), or the Medicaid expansion (Hopkins and Parish, 2019).

³Notably, those with access to qualifying health insurance through employers were ineligible for subsidies on the exchanges. Separately, people whose incomes prove higher than expected are required to repay the excess subsidy.

Potential Limits on the Exchanges' Feedback Effects

Despite the ACA's sometimes sizable benefits, there are several reasons the exchanges might generate muted or even negative feedbacks. We consider three such hypotheses.

First, even if the exchanges benefited the vast majority of users, those benefits might not be attributed to the ACA. The government's role in facilitating the provision of insurance is not a salient feature of the exchanges' design. The exchanges are simply a marketplace: once customers are enrolled, their primary interactions are not with government but a private insurer.

Other aspects of the exchanges' operation are even more opaque, further reducing citizens' capacity to *trace* elements of their insurance back to government action (see also Kogan and Wood 2018; on other policies, see Mettler 2011; Morgan and Campbell 2011; Campbell 2012). For example, citizens may not be aware of the ways that government regulations structure the marketplaces, especially by preventing discrimination based on pre-existing conditions. By contrast, the ACA's design might have inadvertently heightened the salience and traceability of one of the law's least popular elements, its individual mandate (see also Jacobs and Mettler, 2018). While citizens may be very aware of the personal costs from the individual mandate, they may not recognize that the mandate was intended to prevent adverse selection on the exchanges and so keep prices down while protecting people with pre-existing conditions.

Second, the exchanges provide much more substantial benefits to some than others. The experiences of those who use the exchanges are likely to differ dramatically depending on their health care utilization, their eligibility for subsidies, the quality of their state's exchange, the options available in their market, and other factors (Chattopadhyay, 2018). A person forced to buy comprehensive coverage or, conversely, who pays a lot for insurance that leaves her exposed to significant costs may sour on the ACA; someone with a serious pre-existing condition who can get heavily subsidized insurance may feel quite differently. Prices on the exchanges also vary dramatically across the country, meaning that similar people can pay very different premiums depending on where they live (Kamal et al., 2018). This hypothesis yields an observationally distinctive prediction from that above, as it leads us to expect that exchange users with different experiences will differ in their ACA attitudes, too. The exchanges' impacts will not be null but heterogeneous.

Third, even if the people on the exchanges both benefited from the exchanges and attributed

the benefits to the ACA, *the population of Americans using the exchanges was relatively small*. As the percentage of an overall population which uses a given policy declines, that policy becomes “distant” to a large fraction of the electorate, meaning that relatively few people have direct experience with it (Soss and Schram, 2007). That, in turn, makes citizens more reliant on media portrayals and partisan cues in generating attitudes (Jacobs and Mettler, 2018).

Moreover, the ACA’s exchanges were one element in a complex package of reforms. Campbell (2003), Soss and Schram (2007), and Campbell (2012) note that policy feedbacks are more likely when the policy’s beneficiaries are concentrated in ways which encourage them to identify as a coherent group and act on their shared interests. But the ACA’s complex, multi-faceted design means that key beneficiaries have a stake only in specific, often disparate provisions (see also Béland, Rocco and Waddan, 2018; Chattopadhyay, 2018). This complexity may fragment the beneficiary population, reducing its capacity to generate a cohesive identity.

Research Design and Data Sets

Prior research leads us to expect substantial selection bias, as those who anticipate higher health care bills or are not Republican are more likely to enroll via the exchanges (Lerman, Sadin and Trachtman, 2017). As a result, any straightforward comparison of people who did or did not use the exchanges is almost certainly biased.

How, then, to evaluate the exchanges’ impact? Descriptive analyses serve as a useful starting point, since we should expect causal effects to generate associations at a minimum. Cross-sectionally, it is valuable to know whether those who use the exchanges or receive subsidies think differently about the ACA than those without insurance or with insurance from elsewhere. Knowing which groups of Americans are more likely to use the exchanges also helps identify promising discontinuities or other strategies for causal inference. In addition, our descriptive analyses allow us to consider temporal variation: did attitudes toward the ACA among different sub-groups shift at the time of its implementation? In places, we are able to use panel data to track within-respondent changes before and after the ACA’s implementation.

Armed with a knowledge of various statistical associations, we then proceed to estimates that are credibly causally identified. Instead of providing a single, decisive test, we employ multiple tests with different samples, estimands, outcomes, strengths, and weaknesses. Below, we discuss

each test in detail; here, we motivate the tests collectively. To help readers keep track of the various tests, both descriptive and causal, Table 1 lists them.

Question	Data: Variables and Subgroups
<p>1. Cross-sectional association Is there an association between ACA favorability, purchasing health insurance on exchanges?</p>	<p>Favorability ratings among: 1) HTS respondents 2) KFF non-group respondents 3) KFF Kentucky respondents</p>
<p>2. Temporal correspondence Did attitudes change with ACA’s implementation?</p> <p>Was <i>change</i> in insurance status in the post-implementation period related to ACA favorability?</p>	<p>ACA favorability in HTS</p> <p>Panel-based estimates from [name redacted]</p> <p>Association between ACA favorability, insurance source in panel data from [name redacted]</p>
<p>3. Dose-dependence Is association between ACA favorability, exchange use related to level of personal benefit?</p>	<p>Associations between favorability ratings and subsidies: KFF non-group, KFF Kentucky</p>
<p>4. Within-group change (causal tests) Was <i>likely</i> benefit from the exchanges related to ACA favorability?</p> <p>Did ACA’s implementation reduce ACA favorability gap at age 65?</p>	<p>Association between ACA favorability, insurance source scores over time among HTS respondents under 65</p> <p>Regression discontinuity using older HTS respondents before, after Jan. 2014</p>
<p>5. Partisan integration Were those using exchanges favorable toward the ACA before its implementation?</p> <p>Was <i>likely</i> benefit from the exchanges related to party ID?</p>	<p>HTS</p> <p>Causal tests from above with party ID as outcome</p>
<p>6. Heterogeneity in exchange effects Did price changes on local exchanges influence ACA favorability?</p>	<p>HTS merged with exchange pricing, plan data</p>

Table 1: *Research questions, tests, and data sources.*

Data Sources

For these tests, we use a range of available data. One primary source is the KFF's HTS, a rolling cross-sectional telephone survey of adults' attitudes about the ACA and health policy. This survey has been conducted in most months between 2009 and the present; here, we report results for 117,234 respondents surveyed between February 2009 and September 2017. On account of its extraordinary sample size, these surveys provide a unique opportunity to observe the evolution of Americans' ACA attitudes, as well as the possibly differing trajectories of key sub-groups.

Our descriptive analyses supplement the HTS with separate KFF surveys of the non-group insured in 2014, 2015, and 2016 as well as a 2015 survey of Kentucky residents, all of which are presented in the Supplemental Information (SI). For our causal estimates, we return to the HTS when using the difference-in-difference estimation strategy based on insurance source scores as well as that employing a regression discontinuity design for those near the Medicare age threshold. We also merge a geo-coded version of the HTS with administrative data on exchange pricing to estimate the effects of local price changes on ACA favorability.⁴

Last, we draw on the [name redacted] panel survey administered via Knowledge Networks/GfK. This panel is unique, as it tracks a nationally representative, probability-based sample which was recruited through address-based sampling and random-digit dialing over ten years. Specifically, the panel has 13 separate waves: 5 waves in 2007-2008, 2 waves in 2012 (which included the first questions about the ACA specifically), 2 waves in 2014, 3 waves in 2016, and 1 wave in 2018. Table 1 summarizes the data sets employed for each test. See the SI for details.

Descriptive Statistics on ACA Exchanges & Attitudes

To understand the exchanges' impacts on public opinion, we first need to identify who actually uses them. In a given year, only a small fraction of American adults enroll through the exchanges, and those who do have distinctive demographic profiles. Through these descriptive analyses, this section lays the groundwork for subsequent analyses, as we will be better positioned to identify confounding variables and promising strategies for causal estimation. In later analyses, we will

⁴Specifically, we obtained the administrative data on plan pricing and offerings by geography for 2014-2017 via the Robert Wood Johnson Foundation's National Narrow Network Project.

examine attitudes among people who purchased on the exchanges, as well as the much broader set of people with demographics similar to those who purchased on the exchanges, compared to individuals with other insurance sources and their demographically similar cohorts.

Demographic Differences by Insurance Type

SI Figure A1 uses the HTS to illustrate the distribution of types of health insurance before and after the ACA's primary provisions came into effect. It shows the decline in the uninsured rate from 14.4% before January 2014 to 9.7% after. It also shows the uptick in Medicaid receipt that followed the expansion of Medicaid in some states, from 3.7% to 6.6%.

But overall, these changes are marginal: the basic contours of health care provision in the U.S. remain much as they were before the ACA. What's more, the share of U.S. adults who reported using the exchanges is on average only 2% of the surveyed population after their creation. Categories of insurance that were less overtly affected by the ACA—employer-provided insurance and Medicare—remain far and away the largest sources of insurance. These patterns alone place important limits on the ACA's direct policy feedback effects: people in the parts of the insurance market most influenced by the ACA represent a small minority of citizens.

How does the small share of Americans who purchase insurance via the exchanges compare to other groups? SI Table A1 summarizes the means of several key variables by source of insurance for post-January 2014 respondents. Consistent with Lerman, Sadin and Trachtman (2017) and Sances and Clinton (2019), those who bought insurance through the exchanges were more Democratic than the population overall. Medicaid recipients and those who used the exchanges had similar partisanship, even though those using the exchanges had higher incomes and were more likely to be White. This indicates that any causal analyses will have to be wary about selection biases, especially given the general stability of partisan identification. Jointly, the table and figure also show that exchange enrollees had lower incomes and were younger than the population overall.

Trends in ACA Attitudes by Insurance Type

Having provided descriptive statistics for our key independent variables, we now do the same for our dependent variable. Descriptively, is there evidence that ACA favorability differs depending on Americans' sources of health insurance?

Figure 1 summarizes the trends in a dichotomized measure of ACA favorability by insurance

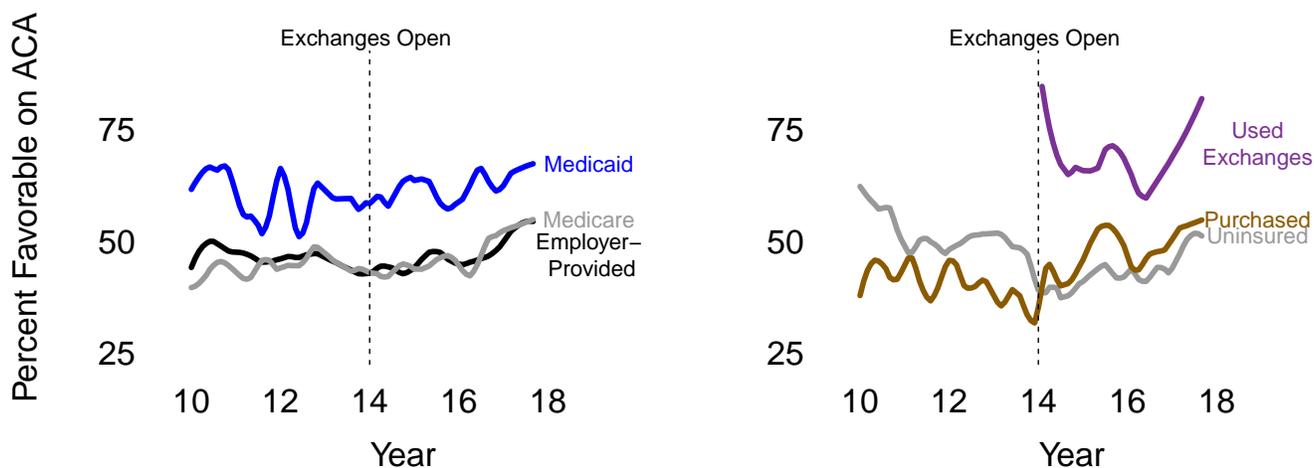


Figure 1: *Dichotomized ACA favorability by health insurance status and source over time.* The survey question for “used exchanges” was introduced in 2014, and it is a subset of “self-purchased”. It measures whether an individual *purchased* insurance through the exchanges.

type. The left side illustrates that respondents insured through Medicaid are always more favorable toward the ACA than those insured through Medicare or employers, with 61% of Medicaid recipients reporting favorable attitudes versus 47% for those with employer-provided insurance and 45% for those on Medicare. The over-time patterns for these groups are roughly similar, with growing favorability between 2016 and 2018, although the smallest group (Medicaid recipients) is understandably more variable.

On the right, we see the trends for the insurance categories for whom the ACA’s exchanges and mandate are especially impactful: those who lack insurance, buy insurance themselves, and buy insurance via the exchanges. In the period before ACA implementation, those without insurance are consistently more favorable toward the ACA than those purchasing insurance on the individual market. The average ACA favorability for those without insurance is 52% in the pre-implementation period, as compared to 41% who are self-insured.

But the groups switch places almost immediately after implementation. Favorability among the uninsured drops sharply in the run-up to implementation and averages just 42% in post-implementation surveys. By contrast, favorability among those on the individual market rises, and averages 47% after implementation (see also McCabe, 2015). Such changes could reflect people’s experiences, as those without insurance faced a newly implemented tax penalty while those needing to self-insure instead enjoyed a range of new options and protections, alongside

subsidies in some cases. These sudden attitudinal changes thus provide a valuable motivation for subsequent analyses. However, such changes might also reflect selection effects.

Finally, we show in the right panel that those who purchased on the exchanges felt particularly favorably toward the ACA. Like Medicaid recipients, this group is small and so favorability over time is noisy. In the SI, we use a regression model of respondents' attitudes pre-implementation to show that exchange users are markedly more favorable—and the uninsured markedly less favorable—than we would expect based on demographics alone.

Attitudes among the Non-Group Insured

Within the broad categories of insurance status identified above, there are sure to be substantial variations in people's experiences. After all, enrollees' experiences with government-funded programs like Medicaid can vary substantially (Michener, 2018; Tallevi, 2018), and the variation in experiences with private insurance is potentially wider still. As a next step, in SI section A.2.1 and in SI Table A3, we consider the predictors of ACA support among those who were insured through the non-group insurance market post-implementation. The core descriptive result of those analyses is that while exchange users generally are no more favorable toward the ACA than others without group-based insurance like Medicare or employer-provided insurance, those who received subsidies are a bit more favorable toward the ACA, especially in its first year of full implementation.

Panel-Based Estimates of Trends in ACA Attitudes by Insurance

Figure 1 shows that around the time of the ACA's implementation, the relationship between people's insurance status and their ACA attitudes shifted: people who purchased their own insurance became more favorable while the uninsured became less so. But it is quite possible that individuals who already felt favorably toward the ACA shifted from one insurance source to another.

Panel data provides one way to avoid this challenge, as the same individuals report their attitudes in multiple waves conducted over time. Here, we draw on the 2012-2018 [name redacted] panel to evaluate Americans' attitudes toward the ACA over time.

To be sure, there simply aren't many respondents who fall into the self-purchased or uninsured categories. Of the 589 panelists who completed both the November/December 2016 wave and the October 2018 wave, 38—or 6%—reported in January 2016 that they purchased their insurance

themselves, and not all used the exchanges. Another 35 panelists—also 6%—reported being uninsured at that time, an estimate that is lower than the 2016 national benchmark of 11% (Witters, 2019). But the panel nonetheless provides a critical advantage, as it allows us to observe how those respondents’ attitudes shifted relative to the population overall. We also get increased statistical precision from making comparisons within individuals. In these analyses, our outcome comes from a question the panel has asked since 2012: “Some people think the health care reform law should be kept as it is. Others want to repeal the entire health care law. Still others are somewhere in between.” Respondents could reply on a 1-7 scale, with 1 indicating the “health care reform law should be kept as it is” and 7 indicating that “entire health care law should be repealed.”

As Figure 2 illustrates, the groups defined by insurance status in 2016 differ at baseline but largely move in parallel. In particular, there is no evidence that those who purchased insurance themselves differ markedly in the trajectory of their ACA attitudes compared to other insured groups, though there is a significant uptick in support for repealing the ACA just after the 2012 election among those without insurance in 2016. Overall, there is little evidence that respondents who purchased their own insurance in 2016 had a distinctive trajectory. Note, however, that for those without insurance, there is somewhat less evidence of an improvement in ACA attitudes between 2016 and 2018.⁵

We next estimate OLS models which more fully leverage the panel data set by analyzing shifts in insurance status as a predictor of shifts in ACA attitudes. This approach tests whether changes in insurance status are associated with changes in ACA attitudes generally in the post-implementation era, rather than focusing on changes immediately after implementation.

Specifically, we model 2018 respondents’ views of the ACA as a function of their fall 2018 insurance status, their January 2016 insurance status, attitudes toward the ACA in 2012 and 2016, 2012 partisan identification, and a series of basic demographics such as gender, education and income. Such models are well suited to isolate the extent to which *changes* in insurance status predict subsequent shifts in ACA attitudes. However, they are a conservative test, as we are isolating people who became uninsured only after 2016, nearly three years after the ACA’s

⁵Specifically, those without insurance in January 2016 shifted toward the ACA by 0.30 while for those with insurance the pro-ACA shift was 0.53.

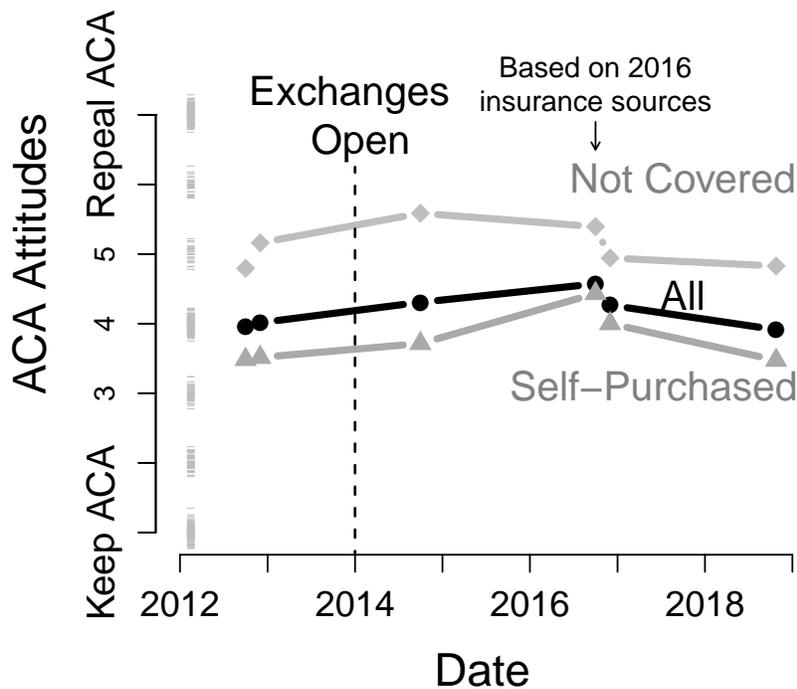


Figure 2: This figure reports estimates for respondents to the 2012-2018 [name redacted] panel who completed the November/December 2016 and October 2018 waves conducted through GfK’s Knowledge Panel. N=589. Insurance status is as of January 2016. The gray horizontal lines at left report the distribution of the outcome variable for the October 2012 survey.

implementation. As a result, those who were uninsured in the first few years after the ACA’s 2014 implementation will not influence our estimate.

The core results are presented graphically in Figure 3, while SI Table A8 provides the full, fitted model. Most of the indicators of different insurance statuses are not predictive of 2018 ACA attitudes. Importantly, those who purchased insurance plans themselves in 2018 are not much more or less supportive of the ACA. However, respondents who were uninsured in 2018 are dramatically less supportive of the ACA, even accounting for 2016 insurance status. The coefficient is 1.23 (SE=0.43), meaning that those who became uninsured between 2016 and 2018 also downgraded their opinion of the ACA by a whopping 1.23 on a 1-7 scale. That effect is 61% of a standard deviation. It is strongly suggestive of a causal relationship between becoming uninsured and thinking more negatively toward the ACA.⁶

⁶As Table A9 illustrates, we reach very similar conclusions with an alternative model specification that examines the change in attitudes among those uninsured in January 2016, indicating

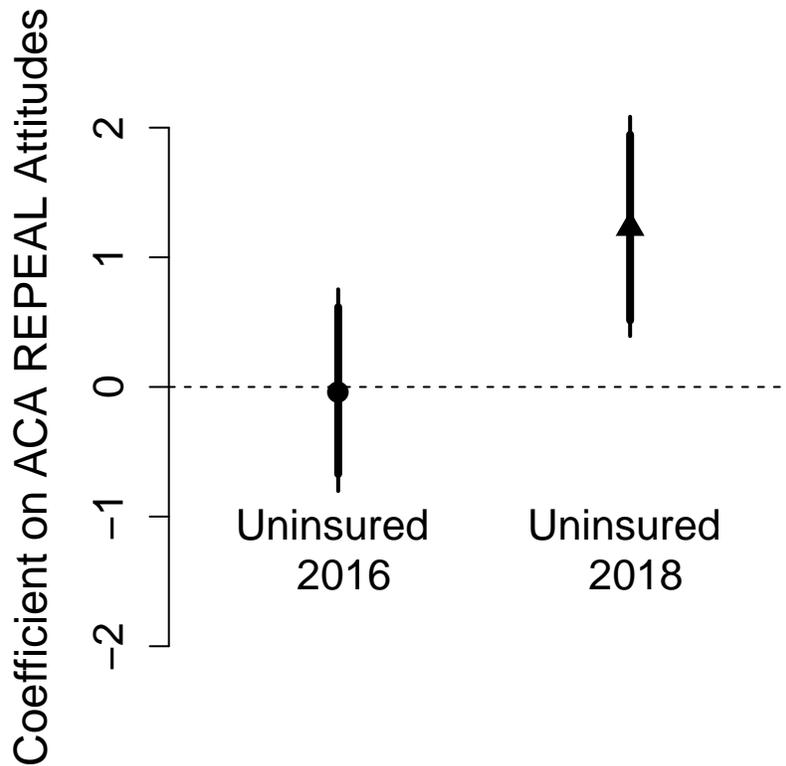


Figure 3: This figure reports coefficients for being uninsured in 2016 and 2018 (included in the same model) when predicting 2018 support for ACA *repeal* using the [name redacted] panel.

Causal Estimates

The following analyses are designed to remove the effects of selection from our estimates to identify the causal effects of changes in insurance status.

Over-time Changes by Predicted Insurance Source

One common research design for assessing a policy introduced at a specific moment in time is difference-in-difference estimation. Such estimators isolate whether there were disproportionate changes in a treated group after the treatment's administration. Yet when analyzing the HTS, which was administered as a rolling cross section, we are barred from employing classical difference-in-difference estimation because there is no fixed control group. There were no federal exchanges or people using them prior to the ACA's implementation. What's more, the exchanges are likely that this result is robust to alternative specifications.

to have had system-wide effects, as they may have influenced the insurance status and experiences of people who were previously uninsured as well as those who previously purchased insurance for themselves and others. Most importantly, we should avoid using actual insurance status (as we did in Figure 1) because the people who chose to use the exchanges may have already felt favorably toward the law.

We attempt to overcome selection concerns by using predicted insurance scores coupled with temporal variation to estimate the causal impact of using the exchanges (and later of being uninsured). In essence, we model who in the data set is more or less likely to use the exchanges (rather than who *chose* to), construct synthetic cohorts of likely users, and then evaluate whether over-time shifts in ACA attitudes were more pronounced among those whom the model flags as especially likely to be exchange users. In doing so, we are guided by prior work on causal inference including Athey and Imbens (2006), Heckman and Vytlacil (2007), and Frangakis and Rubin (2002).

Note that this analysis relies on the timing of implementation, and mimics a difference in difference design – one where we consider changes in attitudes among those with demographics similar to those with given insurance sources. In the language of an encouragement design, we assign intention-to-treat status for all individuals in time periods after 2013 and a probability of complying – taking up the treatment – based on their demographics, whether or not they actually purchase a given insurance source. Like in difference-in-difference analyses, we look for parallel trends (or, more precisely, parallel slopes for likely vs unlikely to purchase) before the implementation and examine relative changes in attitudes after implementation. Those with predicted insurance scores close to zero we expect to never take advantage of the exchanges, and those with lower scores will be more likely to do so. We do not construct matched cohorts that have very similar attitudes before implementation. We also do not predict insurance source using variables that could plausibly be changed by the introduction of the exchanges, including partisanship (we will show changes in partisanship later in the paper).

As an initial step, we isolate respondents who were at all likely to use the exchanges or be affected by the individual mandate. In practice, this means excluding respondents over 64 (who are likely on Medicare) or with incomes under \$40,000 per year (who are more likely to be on

Medicaid, for which there is a competing effect).⁷⁸

Next, we estimate three separate insurance source scores to identify the types of respondents who are especially likely to 1) use the exchanges, 2) purchase their own insurance, or 3) be uninsured in the post-2014 period. Estimating the probability of self-purchasing allows us to consider the effects of changing conditions among self-purchasers on ACA attitudes without requiring that survey respondents link their insurance to government-operated markets. The uninsured score analysis is important because it considers the possibility of negative feedback effects from the individual mandate or other sources.⁹

These insurance source scores provide one-number summaries of respondents' probabilities of falling into specific ACA-related categories given their demographic characteristics.

Health insurance was not available via the exchanges until January 2014, so we use logistic regression to model the predictors of enrolling via the exchanges in that period alone.¹⁰¹¹ Americans' experiences with the exchanges varied substantially across states (Béland, Rocco and Waddan, 2016; Jones, 2017), so the insurance source score model includes indicators for respondents' state of residence alongside 5 polynomials of respondents' age in years, indicators for years of educational attainment, and identification as Asian American, Black, Hispanic, female, or retired. Subsidies and exchange usage are closely connected to respondents' incomes, so the insurance source score

⁷The minimum income cutoff to remove likely Medicaid recipients varies by state. For simplicity, we set this cutoff at the Medicaid expansion level for all states. Reducing this cutoff to include lower-income people in non-expansion states does not meaningfully alter our results.

⁸Very few people who previously received Medicare or Medicaid transferred into the individual market post-ACA. In fact, 53% of all respondents who used the exchanges in 2014 had previously lacked health insurance, and another 19% had previously purchased a plan themselves. See especially SI Table A2, which reports the shares of people with non-group insurance by their previous insurance status.

⁹Those with higher uninsured scores might also be those who are especially vulnerable to the erosion of insurance offerings through their employers.

¹⁰To be coded as using the ACA exchanges, respondents answered that they purchased insurance "From healthcare.gov or [state marketplace name]".

¹¹Using linear regression to produce the insurance source scores does not alter our results.

model also includes measures of respondents' income, income squared, and income cubed. From this model, we can then estimate any respondent's probability of using the exchanges based on her background characteristics, regardless of whether the exchanges were actually operating when she was surveyed.¹²

SI Table A11 shows the correlations for the three insurance source scores. As expected, given that those using the exchanges are a subset of those purchasing their own insurance, the scores for self-purchasing and using the exchanges are correlated at 0.5. Correlations between using the exchanges and being uninsured (0.14) and self-purchasing and being uninsured (-0.08) are substantially smaller. Critically, our second-stage estimates depend on the first-stage predictions—the better we predict in the first stage, the more effectively we isolate those likely to use the exchanges. SI Table A12 shows the first stage of each model.

Results from Predicted Insurance Score Analyses

To distinguish ACA experiences from long-term shifts, we examine over time associations between likely insurance use and ACA favorability. Effects driven by the ACA should be apparent discontinuously in 2013 (when the exchanges first opened) and 2014.

We include the insurance source scores in OLS models predicting dichotomized ACA favorability in each of the six-month increments between early 2010 (when the ACA was passed) and late 2017. Accordingly, Figure 4 shows 16 separate correlations between ACA favorability and respondents' estimated probability of using the exchanges.¹³ If the experience of buying insurance via the exchanges shifted ACA attitudes, the insurance source scores should become stronger predictors of ACA support after the exchanges opened.

Figure 4 illustrates the correlations for a 10 percentage-point change in probability of using the exchanges, self-purchasing, or going without insurance for each six-month window of time.¹⁴ On the right side of each graphic, we report the mean change in the effect for estimates after January

¹²Because partisan identification is potentially endogenous, and also because it is strongly associated with health care attitudes (Kriner and Reeves, 2014), we repeat these analyses with partisanship as the dependent variable below.

¹³The minimum sample size for these six-month periods is 1,739 while the maximum is 6,227.

¹⁴One empirical challenge is that even among those under 65, most people making more than \$40,000 per year have access to employer-provided insurance, so the variation in the estimated

2014. Neither the scores for using the exchanges nor those for self-purchasing are associated with meaningful changes in ACA favorability after implementation. On balance, people more likely to use the exchanges or self-insure don't show different over-time trends.

However, the scores for being uninsured are associated with markedly more *negative* ACA attitudes after implementation. This uninsured score is a catch-all. It identifies both respondents who actually chose to forego insurance and demographically similar respondents who *would have* foregone purchasing on the individual markets had they not had some other source of insurance. Regardless of the specific mechanisms at work, these results reinforce the claim that the ACA induced negative experiences for at least some Americans.

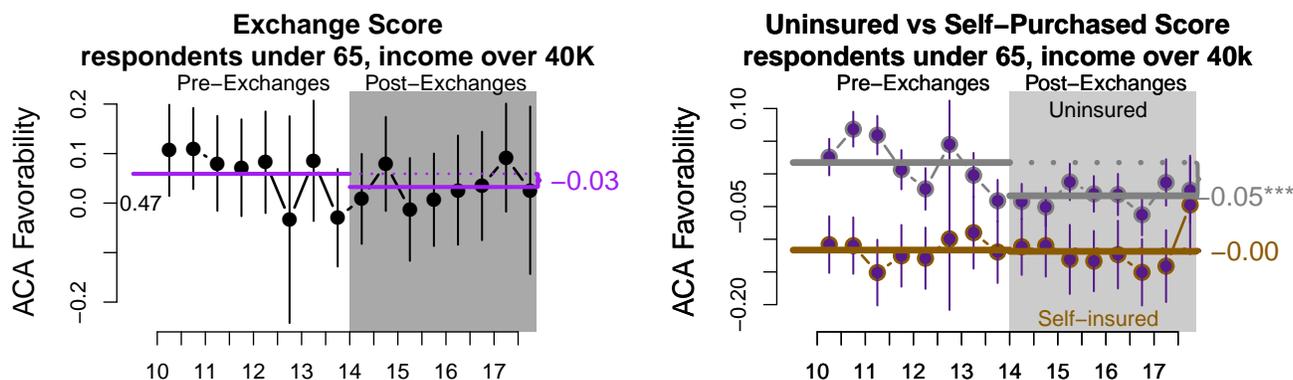


Figure 4: *Over-time associations between insurance source scores, ACA attitudes.* These figures show the effect of increasing the average insurance source score by 10 percentage points.

To formally test the impact of the exchanges' introduction, we estimate a multi-level model predicting ACA favorability using the same insurance source scores and timeframe shown in Figure 4, with respondents clustered by the six-month period of their response. We also include an interaction between the insurance source score and whether the response comes after January 1st, insurance source scores is relatively small, as illustrated in SI Table A10. Accordingly, we generate multiple estimates from the second stage over-time models. In some cases, we show the effect of a 10 percentage point increase in the likelihood of using the exchanges (or of self-purchasing or going uninsured); in others, we report the effect of a full change from 0 to 1 in the insurance source score (the typical interpretation of the coefficients), even though this is an extrapolation outside the range of the observed scores.

2014, which provides a formal measure of any change in the mapping between the probability of using the exchanges and ACA favorability. In this test, we use the typical scaling for such models in which the estimated probability shifts from 0 to 1. The estimates in SI Table A13 otherwise mirror those presented graphically in Figure 4.

Beyond reporting the same results as in Figure 4, Table A13 provides an additional, if more peripheral, take-away: the attitudes among those purchasing on the exchanges appeared to be extremely positive before the ACA's implementation. This indicates possible ceiling effects that limit the potential for further increases in favorability.

Expected Insurance Source & Partisanship

We now replicate the insurance source score tests from above using partisan identification as the outcome. ACA attitudes are strongly associated with partisan identification (e.g. Kriner and Reeves, 2014), and since partisan identification is strongly resistant to change (Green, Palmquist and Schickler, 2002), partisanship could limit the extent to which improvements in personal circumstances can increase favorability toward a politicized policy (Jacobs and Mettler, 2018). At the same time, any changes in partisan identification induced by the ACA might have far-reaching effects, since partisan identification can influence a range of political opinions and behaviors (Achen and Bartels, 2016).

In this, we follow extensive research by considering partisan identification to be causally prior to ACA attitudes (Green, Palmquist and Schickler, 2002). Democrats are more likely to favor the ACA and Republicans the reverse. Here, we consider whether *changes* in exposure to the ACA and the resulting changes in ACA attitudes correspond with *changes* in partisanship.

As in some tests above, we use a linear probability model with an indicator for Democratic identification as the outcome when analyzing the insurance source scores. In select models, we also show the effect of controlling for ACA favorability on any associations between insurance source and partisanship. This is an informal mediation analysis. If the standard assumptions hold¹⁵, the

¹⁵The assumptions underpinning causal mediation analysis are similar to those underpinning linear regressions, but there are additional assumptions about lack of unmeasured confounding (or, at minimum, “deconfounding”) along the expanded causal pathway (Pearl, 2013). In our analysis, for example, a joint effect on ACA favorability and partisanship in the year or so after

decrease in the size of the partisan identification coefficient after controlling for ACA favorability roughly corresponds to the size of the mediated effect.¹⁶

In Figure 5, we show that respondents who are more likely to use the exchanges or purchase insurance do not differ disproportionately in their partisanship after January 2014. However, those who were more likely to be uninsured shifted disproportionately away from the Democrats and toward identification as independents. A 10 percentage-point increase in the likelihood of going without insurance was associated with 3 percentage point decrease (SE=0.7 percentage points) in the probability of being a Democrat after 2014 compared to before 2014.¹⁷

Not identifying with either major party was also positively associated with the uninsured scores. Put differently, people whose demographics made them more likely to be uninsured were also more likely to identify as independents after 2014. At the same time, those predicted to purchase their own insurance were more likely to identify with a party.

Overall, these results mirror the earlier findings. We find that the people who were relatively likely to go without insurance shifted away from the Democrats, and those likely to purchase their own insurance were *polarized* rather than shifting uniformly toward either party. In other words, the ACA's implementation resulted in both positive and negative policy feedback effects, and these heterogeneous experiences partly cancelled out.

the 2012 election that is not due to the ACA would be such a confounder, and would be especially concerning if it were concentrated among individuals with high likelihoods of having the insurance sources considered here. Such a confounder would prove problematic not only for our analyses but for much of the extant research on the effects of the ACA on political attitudes.

¹⁶If the ACA's rollout did change partisan identification with the Republicans or Democrats, the *absence* of mediation would perhaps be more surprising than the presence of mediation. After all, such a finding would imply that the ACA had shifted partisanship without influencing attitudes toward the ACA itself.

¹⁷Note that for the estimates in Figure 5, we are using the full 0 to 1 probability as our independent variable. Dividing the y axis by 10 keeps the estimates closer to the observed variability in our data. For simplicity, the 3 percentage-point decrease above corresponds to a dependent variable coded 1 for Democrats and 0 otherwise.

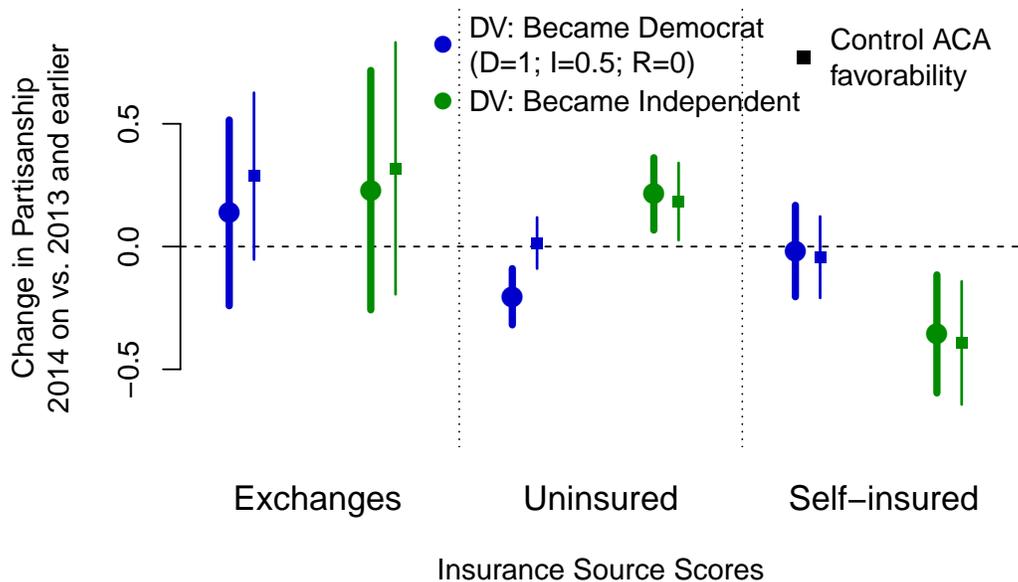


Figure 5: *Partisanship & Insurance Status Scores*. This figure shows estimates for pre- vs. post-implementation associations between partisanship and the insurance status scores. In each case, the first estimate is the change in the partisanship-insurance source score association after implementation while the second reports the same association conditioning on ACA attitudes. Over-time estimates are shown in the supporting information.

The Discontinuity in Medicare Eligibility

The ACA used various policy levers to modify the availability of health insurance, so we continue assessing its causal impacts by analyzing its overall impact on older Americans. Specifically, the ACA provided valuable new opportunities and protections for those in their early 60s: it created exchanges on which non-elderly adults could purchase insurance—sometimes with subsidies—while also mandating that insurers not discriminate on the basis of pre-existing conditions. It also limited premiums for older customers to no more than 3 times those of younger customers. Some in their early 60s became newly eligible for Medicaid. At the same time, the ACA did not entail substantial consumer-facing changes to Medicare, the program which insures 68% of adults 65 and older according to the HTS. To the extent that the ACA affected Americans directly, it was substantially more likely to affect those under 65.¹⁸

¹⁸Policies targeting those 65 and older included free preventative care (for certain services), and the gradual, partial closing of the “donut hole” in prescription drug coverage between 2010 and 2020.

We thus assess the ACA's impacts on those in their early 60s relative to those 65 or slightly older. There is a precedent for this research design. Studying the period before the ACA's full implementation, Lerman and McCabe (2017) sought to understand whether Americans' experiences with publicly provided insurance through Medicare changed their health policy attitudes.

Our goal is to use regression discontinuity designs to estimate the difference in ACA attitudes associated with Medicare eligibility both before and after the ACA's full implementation and then to compare those estimates. To do so, we turn back to the HTS. We first calculate the Imbens-Kalyanaraman optimal bandwidth to be 2.42 (Imbens and Kalyanaraman, 2012), so our initial analyses include respondents ages 62-68. After confirming that key variables do not differ across the discontinuity¹⁹ we then estimate reduced-form equations in which we regress the four-category measure of ACA favorability on several variables, including measures picking up time trends as well as respondents' age in years; Medicare eligibility via age; levels of education; self-identification as male, Black, Hispanic, or Asian American; income; and a five-category measure of partisan identification. In the pre-implementation surveys, we find that being under 65 and not yet eligible for Medicare produces a coefficient of -0.12 (SE=0.06),²⁰ meaning that Medicare eligibility has a nearly significant positive effect on ACA attitudes that averages 10% of the dependent variable's standard deviation, consistent with Lerman and McCabe (2017).

A majority of Americans shift their insurance status at 65 with Medicare eligibility, so the key estimate is not the change at 65 but the extent to which the ACA's implementation modified that change. We thus estimate a parallel OLS model for respondents after the ACA's January 1 2014 implementation, finding that the reduced-form impact of being under 65 became 0.01 (SE=0.06). Post-implementation, those who are under 65 and so more directly affected by the ACA are no more supportive than people just slightly older.

For our purposes, the key question is the extent to which the effect *declined* after the ACA was implemented, and Figure 6 presents that differenced estimate using dots. Using triangles, Figure 6 shows the comparable estimates via an estimator that does not condition on any variables other than age. In both cases, the estimated difference is positive, as it varies between 0.13 (SE=0.087) with controls to 0.183 (SE=0.095) without. Given the four-category dependent variable, *the ACA's*

¹⁹See SI Figure A16.

²⁰See SI Table A15 for the full fitted model.

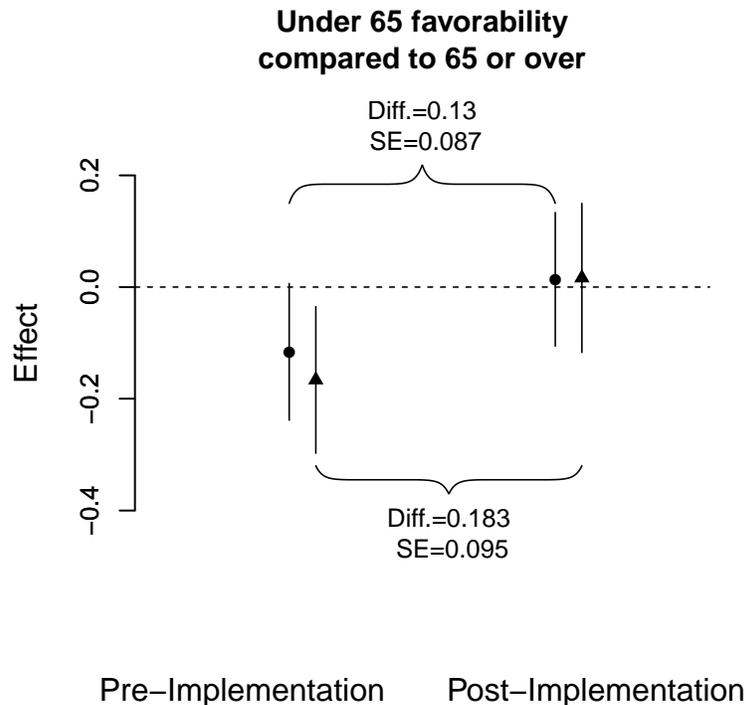


Figure 6: This figure illustrates the change in ACA attitudes associated with turning 65 and becoming Medicare eligible for HTS respondents before and after the ACA was implemented. Estimates from models conditioning on various potential confounders are shown using dots; estimates from models without such controls are depicted using triangles.

*implementation had a detectable but substantively small overall impact closing the attitudinal gap between those who were and were not Medicare eligible.*²¹ Given the ACA's regulations about age and insurance pricing, this group is likely to have been more positively affected by the exchanges than others.

Geographic Variability in Exchange-based Insurance Prices

Do those who use the exchanges largely ignore their personal experience when assessing the ACA, or they do use personal experience but in different ways depending on their experiences? The results above show that the ACA's implementation had heterogeneous effects, suggesting that limited feedback effects were not due to citizens' difficulty attributing their experiences to government policy. Here, we use price shocks on the ACA exchanges to further differentiate hypotheses

²¹This result is of course compatible with a larger effect on the small number of people most directly affected.

about the difficulties of political attribution from those emphasizing heterogeneous personal experiences.

Here, we use geographic variation in exchange pricing as leverage to estimate the attitudinal effects of price changes on local markets. Specifically, customers on the exchanges are divided into geographic rating areas; in 2017, there were 504 such rating areas across the country. We acquired data on the insurance options and their prices for rating area and year between 2014 and 2017. We then developed a crosswalk which enables us to match KFF respondents based on their ZIP codes and/or counties of residence to their rating area and the corresponding options and prices available to them.²² In addition, we also used respondents' ZIP codes or county of residence to merge in basic county-level Census demographic information.

Given how few people in the survey use the exchanges—the sample sizes by year for exchange users are 157, 134, and 151—we fit a multi-level model to borrow strength across the three years and estimate the effects of price shocks jointly. This multilevel model conditions on the total number of plans available on each market as well as the mean premium for that year and the mean change in premiums from the prior year.²³ It also includes random effects for each of the three years as well as various individual- and county-level covariates. Shown in Table 2, and in full in the SI, the model suggests that rising average premiums on the ACA exchanges are associated with sizable declines in ACA favorability for respondents who actually use the exchanges. The estimated coefficient is -0.19 ($SE=0.07$), meaning that an increase of one standard deviation in the mean monthly premium change ($\$33$) is associated with a -0.19 drop on the ACA favorability scale. That shift is substantively meaningful, as it is a change of -7.6 percentage points in the dichotomized ACA favorability scale. However, this relationship does not hold for other respondents or for the uninsured, as the other columns of Table 2 demonstrate. These findings bolster the claim that we are detecting the effects of experience with the markets, and not simply

²²Data are missing for 2014 in some rating areas, but overall, we are able to identify the 2014 market conditions for 64% of respondents, 2015 conditions for 91% of respondents, and 2016 and 2017 conditions for 96% of respondents.

²³Note that our interest in the change in premiums from the prior year requires us to exclude 2014.

	ACA favorability		
	Non-Market	Market	Uninsured
Number of plans (logged, in sd's)	0.00 (0.01)	-0.07 (0.06)	-0.00 (0.03)
Mean Premium (in sd's)	0.00 (0.01)	0.04 (0.07)	0.05 (0.03)
Mean Change in Premium (in sd's)	0.00 (0.01)	-0.19** (0.07)	-0.02 (0.04)
Observations	15,987	442	1,772
Month FEs	Y	Y	Y
County-Level Demographics	Y	Y	Y
Clusters (Year)	3	3	3
*p<0.05; **p<0.01; ***p<0.001			

Table 2: This table presents the results of multi-level models fit to KFF respondents from 2015, 2016, and 2017 in which certain insurance market conditions (and various other independent variables) predict ACA favorability, measured on a 1 to 4 scale.

spurious county-level associations.²⁴

Conclusion

Which features of policy designs make feedbacks on public opinion more or less likely? Prior research on other market-based health policies (Morgan and Campbell, 2011)—and on the more general class of policies administered through back-door mechanisms including private markets and the tax code (Howard, 1999; Mettler, 2011, 2018)—indicates that they typically do not produce strong imprints on public opinion. Such policies can be relatively invisible, making them difficult to trace back to government action.

Key elements of the ACA were administered via market-based mechanisms, making it possible that the ACA's exchanges and its individual mandate might leave a similarly faint imprint on public opinion. Yet the ACA was also highly salient and contentious, so much so that citizens' partisan predispositions shaped their willingness to enroll via the exchanges. In this paper, we

²⁴Our respondents are distributed across many counties—in 2015, for example, the 359 respondents using the markets were in 208 counties. As a consequence, we cannot employ county fixed effects.

employed a wide range of data sets and analyses to test these competing influences on possible policy feedback effects.

Overall, we find that while the net effect of the ACA's core, market-based features was muted, the ACA's exchanges and its individual mandate did influence the opinions of those most likely to be affected. Specifically, differenced regression discontinuity estimates indicate that the ACA reduced the gap in ACA favorability between those who are 64 and 65. Yet those in their early 60s were among those who benefited most from the ACA, and separate analyses using predicted insurance sources indicate that the types of people most likely to be uninsured became *less* favorable toward the ACA around the time of its introduction. A conservative test using panel data demonstrates that those who became uninsured between 2016 and 2018 also became substantially less favorable toward the ACA. In addition, local price spikes on the ACA exchanges reduced support for the law, another indication that Americans were able to link their experiences with their ACA attitudes. Different groups in different places had quite different experiences with the ACA's exchanges and its mandate. One key factor limiting the ACA's mass-level policy feedbacks was heterogeneity, not invisibility.

The politics of the ACA abound in puzzles and paradoxes. One such puzzle is especially relevant here: why were Republican repeal efforts in 2017 more successful when it came to ending the market-based individual mandate than scaling back the Medicaid expansion? Public opinion is certainly not the sole answer, or even necessarily the most important one (Hertel-Fernandez, Skocpol and Lynch, 2016; Jones, 2017; Hacker and Pierson, 2018; Patashnik and Oberlander, 2018). Still, the results uncovered here do suggest a partial explanation: while the Medicaid expansion induced heightened support for the ACA (Sances and Clinton, 2019; Hopkins and Parish, 2019), the overall impacts of the ACA's exchanges and its individual mandate were ambiguous. Those findings, in turn, may help explain why some of the ACA's more conservative, market-based elements were also among the most politically vulnerable.

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**Supplemental Information for “Offsetting Policy Feedback Effects:
Evidence from the Affordable Care Act”**

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A Supplemental Information

A.1 Health Insurance Sources

A.2 KFF Non-Group Surveys

A.2.1 Attitudes among the Non-Group Insured

Those with non-group insurance include people who purchase insurance on the exchanges or otherwise purchase their insurance directly, but exclude those insured through Medicare, Medicaid, or an employer as well as the uninsured. This group's members were among those most directly affected by the ACA. Below, we present similar estimates for another group that saw disproportionate impacts from the ACA: Kentucky residents, whose state saw the most dramatic decline in its post-ACA uninsured rate.

To be sure, the non-group insured is a select population, as its members differ from those with other sources of insurance in terms of their age, income, employment, and various other

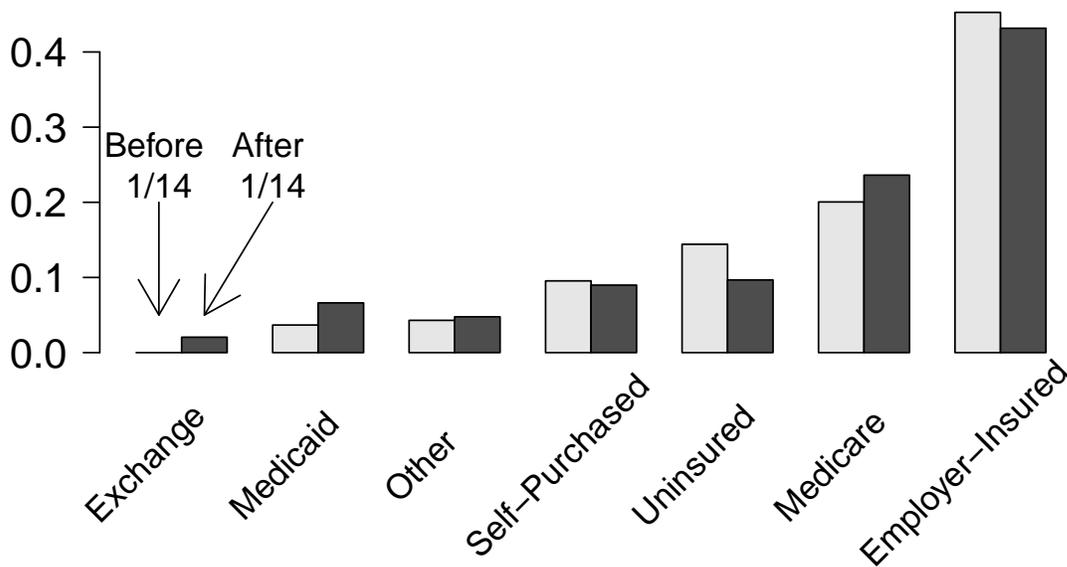


Figure A1: *Distribution of HTS respondents by health insurance source before and after the ACA's January 2014 implementation.*

	<i>Income</i>	<i>Education</i>	<i>Age</i>	<i>Male</i>	<i>Black</i>	<i>Hispanic</i>	<i>Party ID</i>	<i>N</i>
Used Exchanges	59.8	14.7	47	0.50	0.12	0.13	2.5	1,074
Medicaid	26.8	12.7	46	0.42	0.23	0.21	2.5	3,469
Other	57.5	13.4	44	0.54	0.00	0.45	2.6	2,499
Uninsured	37.1	12.5	41	0.58	0.17	0.33	2.8	5,069
Medicare	58.5	14.3	70	0.44	0.09	0.05	2.9	12,379
Employer-Insured	106.6	15.3	48	0.53	0.10	0.09	2.9	22,624
Self-Purchased	76.6	14.6	53	0.51	0.09	0.09	3.0	4,708
All	77.9	14.4	52	0.51	0.11	0.12	2.9	52,424

Table A1: *Key covariates and their associations with insurance sources, respondents to surveys after December 2013.* “Used exchanges” is a subset of “self-purchased.” Party ID is a 5 point scale. All summary values are means.

Prior Coverage	Non-group, All	Non-group, Exchange only
Covered by a different plan you purchased yourself	0.24	0.19
Covered by an employer	0.21	0.14
Covered by COBRA	0.05	0.04
Had Medicaid or other public coverage	0.06	0.06
Was uninsured	0.38	0.53
Covered by parents	0.01	0.01
Covered by family member (non-specific)	0.00	0.00
Military/VA	0.00	0.00
Same coverage/no change (non-specific)	0.01	0.00
Had coverage from some other source	0.01	0.01
Don't know	0.01	0.01
Refused	0.01	0.01

Table A2: This table shows the prior insurance sources for respondents in the KFF 2014 non-group survey.

factors. It is also a moving target. But such an examination can nonetheless identify variation in ACA attitudes that might plausibly be connected to variation in Americans’ experiences with the exchanges.²⁵

²⁵Among those who have non-group insurance, there are a few key distinctions that may appear in respondents’ ACA attitudes. For one, while many of these respondents purchased health insurance through the exchanges, others used a broker or purchased a policy directly from an insurance company. Some of these respondents have plans which meet the requirements for sale on the exchanges, while others do not.

In the years immediately following the ACA’s implementation, KFF surveyed Americans who got their insurance on the non-group market, whether using the exchanges or not.²⁶ The non-group data includes several variables of interest: whether respondents purchased insurance through the exchanges, whether their plan qualifies for sale on the exchanges, and whether they received a subsidy to pay for the plan.

Using the four-category measure of ACA favorability as our outcome, we estimate linear models which predict ACA attitudes as a function of basic demographics including age, education, five-category partisan identification, and identification as male, Asian American, Black, and Hispanic/Latinx. The models also include indicator variables for being surveyed on a landline as well as the respondent’s stratum if surveyed in 2014.²⁷ The models also include a five-category measure of respondents’ income as a percentage of the Federal Poverty Line. Determinations about enrollees’ eligibility for subsidies as well as their size are made based on this coarse measure.

SI Table A3 uses OLS models fit separately for each year to examine the correlates of four-category ACA favorability when looking at those on the non-group insurance market. Simply buying insurance that meets the regulations for exchange-traded plans has no strong association on its own, as none of the corresponding coefficients even reach the size of their standard errors. The coefficient for people who themselves bought insurance on the exchanges is consistently positive and in the range of 0.16 to 0.21; the precision-weighted average is 0.18. While the signal isn’t overwhelming, this result is consistent with the claim that exchange users were somewhat more supportive of the ACA.

One of the ACA’s key policy levers was to subsidize lower-income Americans using the exchanges.²⁸ Receiving a subsidy is associated with more favorable ACA attitudes in all three surveys, with coefficients of 0.49 (SE=0.11), 0.14 (SE=0.11), and 0.09 (SE=0.11). In 2014, people

²⁶The 2014, 2015, and 2016 surveys of individuals on the non-group market included 742, 804, and 786 respondents between ages 18 and 64, respectively.

²⁷The 2014 survey included over-samples of those who are “very high poverty” and “high poverty”; our models include indicator variables for being in these strata.

²⁸To validate the self-reported measure of subsidy receipt, we confirmed that subsidy receipt is concentrated among lower-income respondents. While just 7% of 2014 non-group respondents with incomes over 400% of FPL report receiving subsidies, 59% of those with lower incomes do.

who reported receiving subsidies were markedly more favorable toward the ACA, a relationship that has attenuated since. Substantively, the association between receiving a subsidy and evaluating the ACA favorably declined markedly between 2014 and 2016 ($p=0.01$, two-sided). It's possible that the people opting into the ACA markets have changed over time, and that explains the declining relationship. But it's also possible that initial impact of the subsidies waned as their recipients adjusted to them or as conditions in the marketplaces changed.

A.2.2 Attitudes in Kentucky, 2015

American adults who don't have group-based insurance are one especially instructive group; the attitudes of residents of specific states with unique post-ACA trajectories are another. Kentucky was one of the few Republican-leaning states to create its own health insurance exchange, known as "Kynect," as well as expand its Medicaid program (Hertel-Fernandez, Skocpol and Lynch, 2016). It saw among the largest increases in its percentage of insured residents—and in late 2015, its voters elected Republican Matt Bevin Governor after he had promised to shut down Kynect.

At the time of Bevin's election, the Kaiser Family Foundation conducted a poll of Kentucky residents (which included 692 respondents under 65) and asked various questions about respondents' attitudes toward the ACA and health insurance. Here, we estimate models similar to those just above in which we regress the four-category measure of ACA favorability on a variety of standard demographic questions as well as respondents' sources of insurance (if any). Even in Kentucky, the survey data indicate that only 10% of adults under 65 had insurance they purchased themselves, and just 21% of those respondents reported having gotten their insurance through Kynect. What's more, only 13% of adults who purchased their own insurance—and just 2% of adults under 65 overall—report receiving a subsidy to help with their insurance premiums.²⁹ As with some other analyses, the coefficients are not meant to provide causal estimates, but they do provide conditional correlations which will guide other inquiries.

Table A4 presents the fitted models. In some, we include a measure for whether the respondent used Kynect while in the others we do not. People who purchase their own insurance are not noticeably different from other Kentucky respondents, with coefficients that are substantively

²⁹Again, we restrict our analyses to those under 65 given that the exchanges target that population.

small and inconsistent in their signs. Nor is there evidence that people who use the “Kynect” exchange were more favorable—there, the coefficients are: -0.20 (SE=0.15) or -0.24 (SE=0.15). But conditional on income, partisanship, and other background characteristics, respondents who receive Medicaid are somewhat more favorable toward the ACA than others, with coefficients varying from 0.21 (SE=0.15) to 0.35 (SE=0.15) depending on the specification.

We observe too few people receiving a subsidy to estimate its correlation with any precision, but it is noteworthy that the coefficient corresponding to receiving a subsidy is substantively very large, at 0.32-0.46. Still, this estimate for Kentucky residents with subsidies matches the estimate from the analyses of all adults with non-group insurance in 2014 nationwide closely. As best we can tell, receiving a subsidy is associated with a substantial boost in ACA favorability, even conditional on income, education, and indicators for respondents’ racial/ethnic backgrounds.

ACA favorability			
	2014	2015	2016
Age	-0.01 (0.003)	-0.001 (0.003)	-0.004 (0.003)
Education - 12 years	-0.12 (0.17)	-0.25 (0.17)	0.08 (0.23)
Education - 13	-0.02 (0.28)	-0.14 (0.29)	-0.60 (0.49)
Education - 14	0.10 (0.17)	-0.38* (0.17)	0.29 (0.24)
Education - 16	0.24 (0.18)	-0.34 (0.18)	0.17 (0.24)
Education - 19	0.45* (0.20)	-0.10 (0.20)	0.37 (0.26)
Income (percentage of Federal Poverty Line)	0.01 (0.03)	-0.02 (0.02)	-0.04 (0.03)
Male	-0.09 (0.09)	0.04 (0.08)	0.02 (0.09)
Hispanic	0.55*** (0.14)	-0.004 (0.13)	-0.48 (0.27)
Black	0.47*** (0.14)	0.08 (0.13)	0.55*** (0.16)
Asian	0.66** (0.25)	0.13 (0.22)	-0.29 (0.24)
Weak Democrat	-0.28* (0.13)	0.07 (0.12)	-0.36** (0.13)
Independent	-0.63*** (0.12)	-0.83*** (0.12)	-0.63*** (0.13)
Weak Republican	-0.98*** (0.15)	-1.13*** (0.14)	-1.16*** (0.16)
Strong Republican	-1.27*** (0.13)	-1.18*** (0.11)	-1.23*** (0.13)
Health Plan Meets Exchange Regulations	0.06 (0.19)	-0.05 (0.14)	0.16 (0.14)
Used exchange	0.21 (0.19)	0.19 (0.13)	0.16 (0.13)
Insurance subsidized	0.49*** (0.11)	0.14 (0.11)	0.09 (0.11)
Intercept	2.69*** (0.28)	3.42*** (0.26)	2.89*** (0.30)
Observations	577	687	538

*p<0.05; **p<0.01; ***p<0.001

Table A3: *Models of ACA favorability among the non-group insured, surveyed by KFF in 2014-2016 (full models)*. Note that measures of the respondents' strata for the 2014 survey and use of a cell phone and family size for all three surveys were also included.

	ACA favorability			
	Model 1	Model 2	Model 3	Model 4
Received Medicaid	0.30*	0.21	0.35*	0.27
	(0.14)	(0.15)	(0.15)	(0.15)
Purchased Own Insurance	0.11	-0.06	0.12	-0.07
	(0.18)	(0.20)	(0.18)	(0.20)
Employer-Provided Insurance	0.03	-0.09	0.005	-0.12
	(0.13)	(0.15)	(0.14)	(0.15)
Received Subsidy		0.32		0.46
		(0.40)		(0.41)
Insured		0.41*		0.41*
		(0.20)		(0.20)
Used Kynect			-0.20	-0.24
			(0.15)	(0.15)
Employed	-0.17	-0.14	-0.15	-0.12
	(0.11)	(0.11)	(0.11)	(0.11)
Male	-0.10	-0.07	-0.10	-0.08
	(0.10)	(0.10)	(0.10)	(0.10)
Income - 25	-0.42**	-0.38*	-0.44**	-0.40*
	(0.16)	(0.16)	(0.16)	(0.16)
Income - 35	-0.24	-0.23	-0.26	-0.25
	(0.18)	(0.18)	(0.18)	(0.18)
Income - 45	-0.53**	-0.49**	-0.53**	-0.50**
	(0.18)	(0.18)	(0.18)	(0.18)
Income - 62.5	-0.50**	-0.46*	-0.53**	-0.48**
	(0.18)	(0.18)	(0.18)	(0.18)
Income - 82.5	-0.39	-0.35	-0.40	-0.36
	(0.23)	(0.23)	(0.23)	(0.23)
Income - 95	-0.46	-0.40	-0.47	-0.41
	(0.29)	(0.29)	(0.29)	(0.29)
Income - 200	-0.64**	-0.60**	-0.65**	-0.62**
	(0.20)	(0.20)	(0.20)	(0.20)
Age	-0.01	-0.01	-0.01	-0.01
	(0.004)	(0.004)	(0.004)	(0.004)
Black	0.45*	0.46*	0.45*	0.45*
	(0.19)	(0.19)	(0.19)	(0.19)
Hispanic	-0.18	-0.08	-0.19	-0.08
	(0.27)	(0.27)	(0.27)	(0.27)
Weak Democrat	-0.18	-0.20	-0.18	-0.19
	(0.16)	(0.16)	(0.16)	(0.16)
Independent	-0.85***	-0.85***	-0.84***	-0.84***
	(0.15)	(0.15)	(0.15)	(0.15)
Weak Republican	-0.78***	-0.78***	-0.78***	-0.79***
	(0.16)	(0.16)	(0.16)	(0.16)
Strong Republican	-1.06***	-1.06***	-1.07***	-1.07***
	(0.12)	(0.12)	(0.12)	(0.12)
Education - 10	-0.04	-0.14	0.01	-0.09
	(0.30)	(0.30)	(0.30)	(0.30)
Education - 12	-0.12	-0.18	-0.09	-0.14
	(0.27)	(0.27)	(0.28)	(0.28)
Education - 13	-0.20	-0.26	-0.15	-0.21
	(0.28)	(0.29)	(0.29)	(0.29)
Education - 14	0.09	0.02	0.13	0.06
	(0.29)	(0.29)	(0.29)	(0.29)
Education - 16	0.06	-0.01	0.10	0.03
	(0.30)	(0.30)	(0.30)	(0.30)
Education - 17	-0.08	-0.15	-0.05	-0.12
	(0.58)	(0.58)	(0.58)	(0.58)
Education - 19	0.17	0.12	0.21	0.16
	(0.31)	(0.31)	(0.31)	(0.31)
Intercept	3.60***	3.32***	3.58***	3.31***
	(0.33)	(0.36)	(0.33)	(0.36)
Observations	567	567	567	567

*p<0.05; **p<0.01; ***p<0.001

Table A4: OLS models of respondents to the 2015 Kentucky KFF survey (full model).

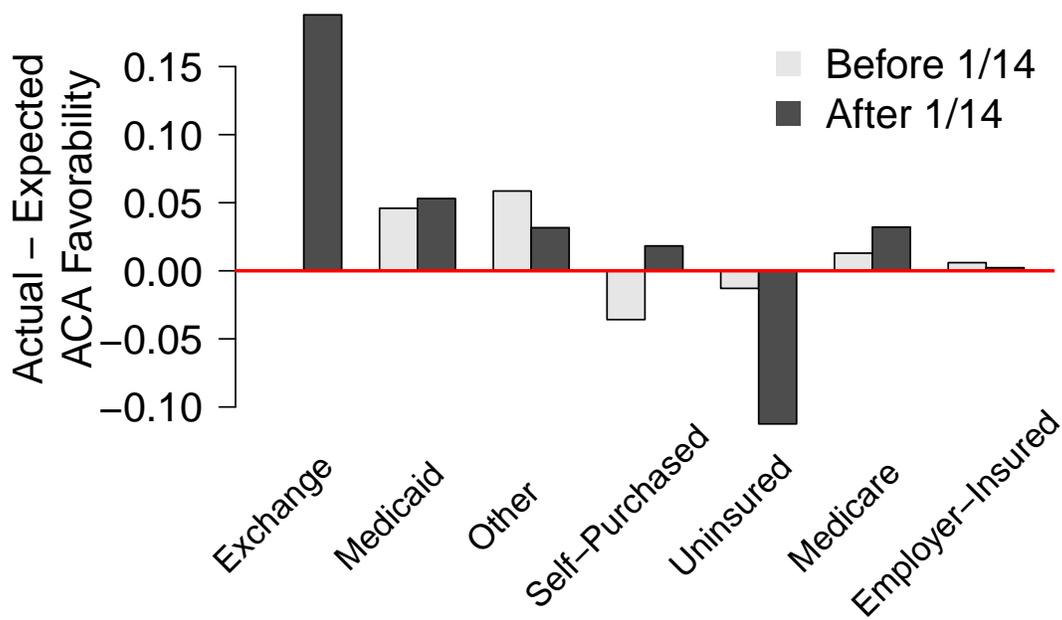


Figure A2: *Actual versus expected favorability by insurance source.* Expected favorability is estimated using a linear regression on all respondents in the pre-2014 period. The actual versus expected estimates are the observed ACA attitudes minus the predicted ACA attitudes.

A.3 Benchmarking with a Pre-Implementation Model

To contextualize the results in the main manuscript on favorability by insurance source, we here consider expected ACA favorability by insurance source. Using the same basic demographic variables as in the insurance source score analysis below, we predict favorability toward the ACA in the pre-implementation period. We then calculate the difference between actual ACA favorability and predicted favorability for all insurance types. Figure A2 shows these differences, with exchange users significantly more favorable toward the ACA than their demographics would suggest while those without insurance were significantly more negative.³⁰ SI Table A5 shows the model predicting favorability by demographics from which expected favorability is estimated.

Like the earlier results, this analysis is subject to selection concerns. Even so, it provides a sense of the magnitude of the potential selection effect. Also, SI Figure A3 shows these same

³⁰Note that the pre-implementation differences are model residuals—the key point here is the difference pre- and post-implementation.

effects multiplied by the population size within each insurance category. Even subtle changes in attitudes among those with employer-provided insurance or Medicare—much larger groups than exchange users—can have an aggregate impact on ACA attitudes equal to that generated by much more dramatic shifts among exchange users themselves.

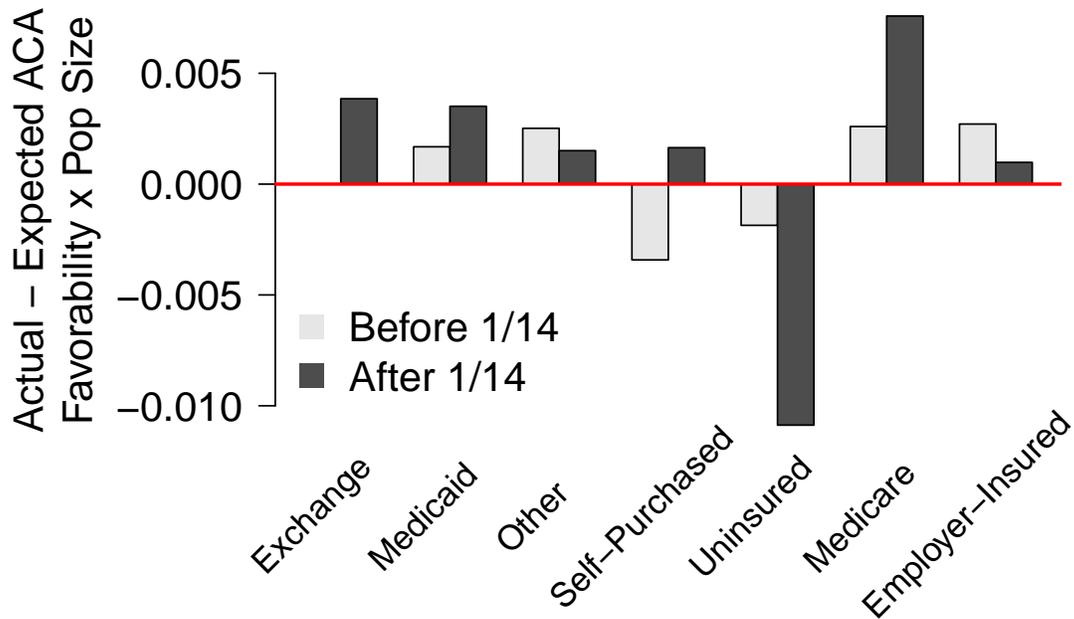


Figure A3: This figure presents the difference between each group’s actual ACA favorability and its expected ACA favorability given its demographics multiplied by its overall population size. When summed, very small declines in attitudes among those with employer-provided insurance are similar in size to the total increase in ACA favorability among Medicaid recipients.

	ACA Favorability
Education - 10	-0.02 (0.02)
Education - 12	-0.07*** (0.02)
Education - 13	-0.05** (0.02)
Education - 14	-0.04* (0.02)
Education - 16	0.004 (0.02)
Education - 17	0.06** (0.02)
Education - 19	0.11*** (0.02)
Income	-0.06*** (0.01)
Income ²	-0.01 (0.01)
Income ³	0.02** (0.01)
Age	0.003 (0.01)
Age ²	0.02* (0.01)
Age ³	-0.04*** (0.01)
Age ⁴	-0.003 (0.002)
Age ⁵	0.01*** (0.002)
Black	0.37*** (0.01)
Hispanic	0.22*** (0.01)
Asian	0.20*** (0.02)
Male	-0.05*** (0.01)
Retired	0.03*** (0.01)
Constant	0.32*** (0.03)
Observations	34,169
F Statistic	62.42

Table A5: *ACA favorability model using demographics only, HTS data pre-implementation.* State fixed effects not shown.

A.4 Panel Details

Wave	Start Date	End Date	N
Wave 1	October 2nd, 2007	December 31st, 2007	19,190
Wave 2	January 1st, 2008	March 31st, 2008	17,747
Wave 3	April 2nd, 2008	August 28th, 2008	20,052
Wave 4	August 29th, 2008	November 4th, 2008	19,241
Wave 5	November 5th, 2008	January 20th, 2009	19,234
Wave 6	October 19th, 2012	October 29th, 2012	2,606
Wave 7	November 14th, 2012	January 29th, 2013	2,471
Wave 8	October 17th, 2014	October 31st, 2014	1,693
Wave 9	November 19th, 2014	January 14th, 2015	1,493
Wave 10	January 22nd, 2016	February 8th, 2016	1,562
Wave 11	October 14th, 2016	October 24th, 2016	1,227
Wave 12	November 28th, 2016	December 7th, 2016	1,075
Wave 13	October 23rd, 2018	November 5th, 2018	1,024

Table A6: This table summarizes the panel waves and sample sizes.

Here, we detail the demographics of our GfK-based panel. Our respondents are older than the U.S. population, but that is to be expected: they had to be 18 in late 2007 to participate. On a variety of other metrics, even the sample which participated in the final waves are a reasonable approximation of the target population of U.S. adults over 25 (see Appendix Table A7). For example, our sample's mean income in 2016 was \$58.4K, which is not far from the 2015 U.S. median household income of \$54.9K. Of particular importance is the fact that there is no evidence of heightened attrition rates among those who are less politically engaged. We merged our data with validated vote histories provided by the data vendor Catalist and found that voter turnout was essentially indistinguishable among those who did and did not remain in the panel between 2012 and 2016. Specifically, 2008 turnout was 69.4% among the 2,471 respondents to the post-election 2012 wave and 69.0% among the 1,075 respondents who participated in post-election 2016 wave.

	Min 2008	Max 2008	Mean 2008	Miss. 2008	Mean 2012	Miss. 2012	Mean 2016	Miss. 2016	ACS 2015
Income '08*	2.50	250.00	61.38	0.31	57.72	0.07	58.40	0.00	54.89
Years of Ed. '08	4.00	19.00	14.33	0.00	13.76	0.00	13.68	0.00	
HS Degree '08	0.00	1.00	0.96	0.00	0.94	0.00	0.94	0.00	0.87
Has BA '08	0.00	1.00	0.40	0.00	0.31	0.00	0.30	0.00	0.30
Party ID '08	1.00	7.00	3.87	0.15	3.82	0.17	3.87	0.14	
Union Hsh. '08	0.00	1.00	0.09	0.00	0.12	0.00	0.13	0.00	
Catholic '08	0.00	1.00	0.16	0.00	0.21	0.00	0.20	0.00	
Protestant '08	0.00	1.00	0.27	0.00	0.31	0.00	0.33	0.00	
Female '08	0.00	1.00	0.56	0.00	0.53	0.00	0.50	0.00	0.51
Age '08	18.00	110.00	50.13	0.00	47.12	0.00	48.84	0.00	
Over 65 '08	0.00	1.00	0.17	0.00	0.14	0.00	0.15	0.00	0.15
Black '08	0.00	1.00	0.09	0.00	0.13	0.00	0.12	0.00	0.13
Hispanic '08	0.00	1.00	0.06	0.00	0.10	0.00	0.10	0.00	0.17
White '08	0.00	1.00	0.80	0.00	0.71	0.00	0.71	0.00	0.77
Voted '12					0.69	.21	0.69	0.21	

Table A7: Demographics for the (1) 19,241 respondents to panel wave 4 in 2008; (2) 2,471 respondents to panel wave 7 in 2012; (3) 1,075 respondents to panel wave 16 in November-December 2016. “Miss.” refers to the share of that variable which is missing for respondents to the designated panel wave. The American Community Survey benchmarks come from July 1, 2015 estimates for the full U.S. population. The asterisk (*) denotes that the U.S. Census reports median household income, not mean income.

	Repeal ACA (1-7)
Intercept	3.56*
	(0.35)
Uninsured, Fall '18	1.23*
	(0.43)
Uninsured, Jan. '16	-0.03
	(0.40)
Jan. '16: Medicare	-0.13
	(0.33)
Jan. '16: Plan through parent	-0.91
	(1.10)
Jan. '16: Plan through employer	-0.15
	(0.34)
Jan. '16: Self-purchased plan	-0.62
	(0.40)
Jan. '16 Insurance: Refused	-1.04
	(0.99)
Jan. '16: Other government insurance	-0.30
	(0.47)
Jan. '16: Somewhere else	0.38
	(0.45)
Jan. '18: Medicare	0.07
	(0.37)
Jan. '18: Plan through parent	2.80
	(1.66)
Jan. '18: Plan through employer	0.45
	(0.39)
Jan. '18: Self-purchased plan	0.56
	(0.46)
Jan. '18 Insurance: Refused	1.03
	(1.76)
Jan. '18: Other government insurance	0.75
	(0.53)
Jan. '18: Somewhere else	-0.68
	(0.51)
Oct. '12: Repeal ACA	0.60*
	(0.11)
Nov. '12-Jan. '13: Repeal ACA	-0.02
	(0.10)
Oct. '16: Repeal ACA	0.76*
	(0.08)
Oct. '12: GOP Partisan ID	0.34*
	(0.07)
Oct. '12: Years of Education	0.10
	(0.06)
Oct. '12: White	0.19
	(0.14)
Oct. '12: Black	-0.27
	(0.22)
Fall '07: Income	0.01
	(0.05)
Oct. '12: Female	0.02
	(0.11)
Oct. '12: Age	0.24*
	(0.08)
Fall '07: Union household	0.13
	(0.17)
Fall '07: Catholic	-0.05
	(0.15)
Fall '07: Protestant	0.21
	(0.13)
R ²	0.65
Num. obs.	567

* $p < 0.05$

Table A8: This table presents an OLS model of fall 2018 anti-ACA attitudes, measured on a 1-7 scale, as a function of various variables.

	Repeal ACA '18 Model 1	Repeal ACA '18 Model 2
(Intercept)	3.83* (0.26)	3.78* (0.29)
Jan. '16: Medicare	0.04 (0.28)	-0.14 (0.29)
Jan. '16: Plan through parent	-0.78 (1.09)	-0.73 (1.10)
Jan. '16: Plan through employer	0.08 (0.27)	0.04 (0.27)
Jan. '16: Self-purchased plan	-0.31 (0.33)	-0.33 (0.33)
Jan. '16 Insurance: Refused	-0.98 (0.99)	-1.02 (0.99)
Jan. '16: Other government program	0.19 (0.41)	0.10 (0.41)
Jan. '16: Other insurance	0.34 (0.42)	0.19 (0.43)
Jan. '16: Uninsured	0.53 (0.35)	0.54 (0.35)
Oct. '12: ACA attitudes	0.61* (0.10)	0.59* (0.11)
Nov. '12-Jan. '13: ACA attitudes	0.03 (0.10)	0.03 (0.10)
Oct. '16: ACA attitudes	0.75* (0.08)	0.76* (0.08)
Oct. '12: GOP partisan ID	0.35* (0.07)	0.32* (0.07)
Oct. '12: Years of Ed.		0.09 (0.06)
Oct. '12: White		0.17 (0.14)
Oct. '12: Black		-0.20 (0.22)
Fall '07: Income		0.00 (0.05)
Oct. '12: Female		0.04 (0.11)
Oct. '12: Age		0.18* (0.08)
Fall '07: Union Household		0.12 (0.17)
Fall '07: Catholic		-0.09 (0.15)
Fall '07: Protestant		0.19 (0.13)
R ²	0.62	0.63
Num. obs.	573	573

* $p < 0.05$

Table A9: This table presents an OLS model of fall 2018 anti-ACA attitudes, measured on a 1-7 scale, as a function of various variables.

A.5 Insurance Source Score Tests

	Market score	Uninsured score	Self-purchased score
Mean	0.03	0.06	0.10
SD	0.02	0.06	0.04
Min	0.00	0.00	0.01
Max	0.34	0.75	0.54

Table A10: *Summary statistics for insurance source scores.* Given that we observe a relatively small amount of variability in these scores, we limit our interpretation of coefficient sizes in several analyses to a change of 0.1 in the scores. Although the models themselves are no different depending on this scaling, this limited interpretation is perhaps more clear and faithful to the amount of variation in the data we have available.

	Predicted Insurance Source		
	Exchanges	Uninsured	Self-purchased
Exchanges	1.00	0.14	0.53
Uninsured	0.14	1.00	-0.08
Self-purchased	0.53	-0.08	1.00

Table A11: Exchange scores are highly correlated with self-purchased scores because exchanges are a subset of the self-purchased in the data.

	Exchanges	Uninsured	Self-Purchased
Education - 10	-0.01 (0.02)	-0.18*** (0.03)	-0.05* (0.03)
Education - 12	0.001 (0.01)	-0.26*** (0.02)	-0.04 (0.03)
Education - 13	0.01 (0.01)	-0.29*** (0.02)	-0.04* (0.03)
Education - 14	-0.001 (0.01)	-0.30*** (0.02)	-0.04 (0.03)
Education - 16	0.02 (0.01)	-0.31*** (0.02)	-0.01 (0.03)
Education - 17	0.002 (0.02)	-0.31*** (0.03)	-0.02 (0.03)
Education - 19	0.01 (0.01)	-0.33*** (0.02)	-0.03 (0.03)
Income	-0.01 (0.02)	-0.01 (0.03)	-0.004 (0.03)
Income ²	0.02*** (0.005)	0.03*** (0.01)	0.02*** (0.01)
Income ³	-0.01 (0.01)	-0.02 (0.02)	-0.01 (0.02)
Age	0.78*** (0.18)	-2.04*** (0.28)	2.55*** (0.31)
Age ²	0.04 (0.18)	-1.03*** (0.27)	0.51* (0.30)
Age ³	0.30* (0.17)	1.52*** (0.26)	0.57* (0.30)
Age ⁴	-0.04 (0.17)	0.05 (0.26)	0.26 (0.29)
Age ⁵	0.24 (0.17)	0.41 (0.25)	0.57** (0.29)
Black	-0.002 (0.004)	0.01 (0.01)	-0.03*** (0.01)
Hispanic	0.003 (0.004)	0.03*** (0.01)	-0.01 (0.01)
Asian	-0.01*** (0.01)	-0.02** (0.01)	-0.01 (0.01)
Male	0.002 (0.002)	0.001 (0.004)	0.02*** (0.004)
Retired	0.0001 (0.01)	-0.02** (0.01)	0.05*** (0.01)
Constant	0.01 (0.02)	0.31*** (0.03)	0.10*** (0.03)
Observations	17,018	19,688	19,688
F Statistic	3.38	17.82	5.33

Table A12: *Insurance source score models*. State fixed effects not shown.

	ACA favorability		
	Purchased on Exchanges	Uninsured	Self-Insured
Intercept	0.45*** (0.01)	0.45*** (0.01)	0.58*** (0.01)
Insurance Source Score	0.71*** _o (0.19)	0.18*** (0.06)	-1.18*** _o (0.09)
Post-Implementation	0.01 (0.02)	0.03* (0.02)	-0.01 (0.02)
Post-Implementation × Insurance Source Score	-0.39 (0.26)	-0.52*** (0.08)	-0.04 (0.13)
Observations	35,474	35,474	35,474
Number of period random-effects	10	10	10

*p<0.05; **p<0.01; ***p<0.001; _o leads to prediction outside 0-1

Table A13: This table presents the results of a multi-level model predicting four-category ACA favorability. Unlike the figure above, these scores estimate moving from 0% to 100% probability. This extrapolates from the observed data, since the demographic variables do not strongly predict purchasing on the exchanges and being uninsured. The insurance source score result for purchasing on the exchanges suggests that these groups were already extremely favorable toward the law, and perhaps could only maintain that high opinion or lower it.

A.6 Integration of ACA Attitudes & Partisanship

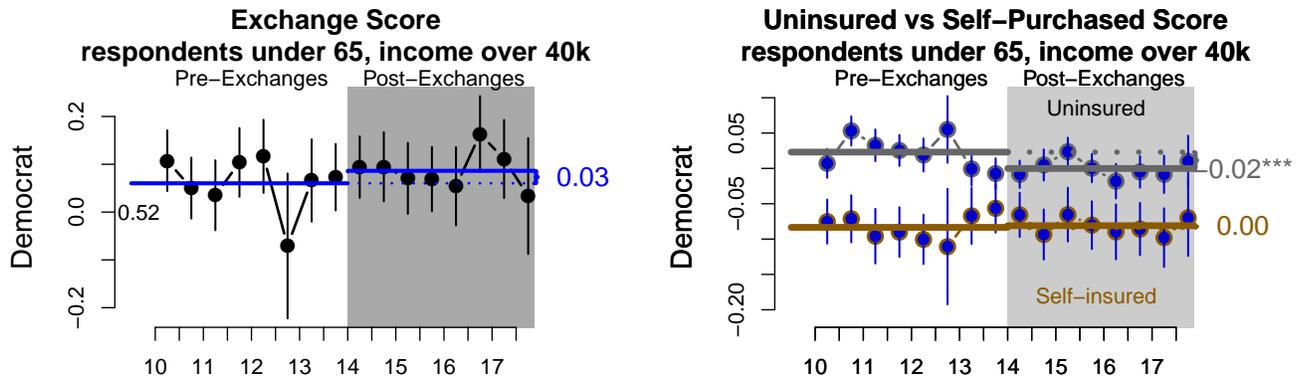


Figure A4: *Over-time associations between insurance source scores, partisan identification.* This figure shows the effect of a 10-percentage-point increase over the mean score. The uninsured score is meant to capture the effect of the individual mandate on ACA attitudes. However, the uninsured score is a catch-all score—it identifies both those groups who chose to forego insurance rather than purchase insurance on the individual market and those groups who *would have* foregone purchasing on the individual markets had they not had some other source of insurance, such as an employer-based health plan. Partisanship here is measured via 1=Democrat, 0.5=Independent, and 0=Republican.

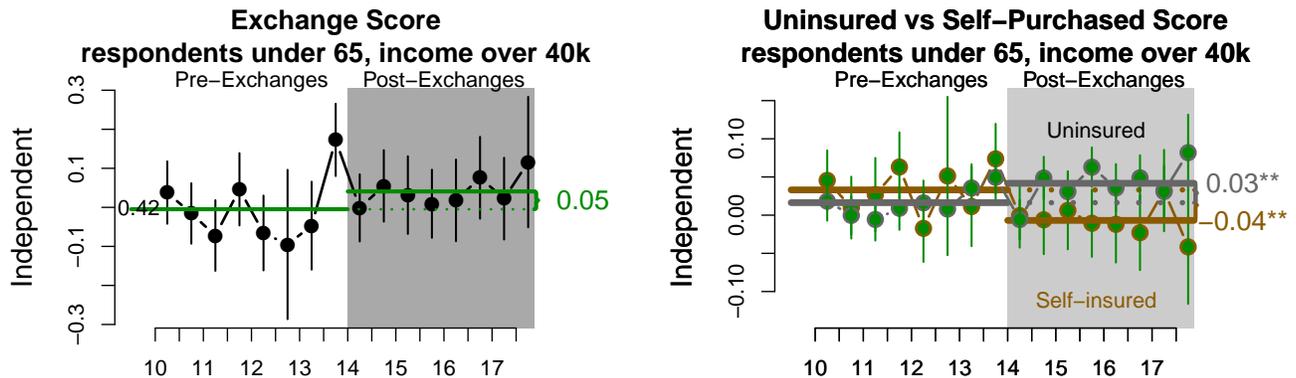


Figure A5: *Over-time associations between insurance source scores, identification as independent.* This figure shows the effect of a 10-percentage-point increase over the mean score.

	Party ID		Identify as Independent	
	(Identify as Democrat)			
	(control ACA)		(control ACA)	
Under 65	-0.02 (0.02)	-0.01 (0.02)	-0.01 (0.02)	-0.04* (0.02)
Post-Implementation	-0.05*** (0.02)	-0.04*** (0.01)	-0.02 (0.02)	-0.05** (0.02)
Post-Implementation x Under 65	0.02* (0.01)	0.01 (0.01)	0.01 (0.02)	0.04** (0.02)
ACA favorability		0.16*** (0.003)		-0.04*** (0.004)
Observations	11,276	9,499	11,276	9,499

*p<0.05; **p<0.01; ***p<0.001

Table A14: “Democrat” is measured via 3-category partisan identification, with Democrat coded 1, Republican coded 0, and Independent coded 0.5.

	Pre-ACA	Post-ACA
Intercept	3.57*	1.91
	(1.24)	(1.37)
Age > 64	-0.12	0.01
	(0.06)	(0.06)
Age	-0.02	-0.00
	(0.02)	(0.02)
Survey Month	1.92*	1.43
	(0.92)	(1.65)
Survey Month Sq.	-2.18	-0.60
	(1.14)	(1.00)
Education	0.05*	0.03*
	(0.01)	(0.01)
Black	0.49*	0.44*
	(0.06)	(0.05)
Hispanic	0.37*	0.40*
	(0.07)	(0.06)
Asian Am.	0.29*	0.28*
	(0.13)	(0.13)
Male	-0.04	-0.01
	(0.03)	(0.03)
Income	0.20	0.09
	(0.29)	(0.26)
Weak Dem	-0.24*	-0.85*
	(0.05)	(0.05)
Pure Independent	-0.96*	-0.82*
	(0.06)	(0.05)
Weak GOP	-1.60*	-0.94*
	(0.05)	(0.05)
Strong GOP	-1.60*	-1.56*
	(0.04)	(0.04)
R ²	0.42	0.32
N	3843	4791

* $p < 0.05$

Table A15: These tables present the RDD models fit to respondents between 62 and 68 before (left) and after (right) the ACA's implementation.

	Beta	SE	t-value
Education	-0.108	0.133	-0.813
Black	0.015	0.013	1.158
Hispanic	0.008	0.011	0.698
Asian American	0.012	0.005	2.285
Male	-0.000	0.022	-0.020
Republican	-0.024	0.080	-0.296
Income	-1.250	2.912	-0.429

Table A16: This table presents checks of the key assumption underpinning the RDD analyses, which is that potentially confounding variables are distributed smoothly at the point of the discontinuity. Using the same specification as in the RDD models in Table A15, it presents the coefficient, standard error, and t-value from models in which the listed variable is the dependent variable predicted by being 65 or older. As the table makes clear, none of these variables are significantly different for those just over 65 years old save for being Asian American, which is slightly more common just over 65.

A.7 Geographic Variability in Exchange-based Insurance Prices

	ACA favorability		
	Non-Market	Market	Uninsured
Weak Democrat	-0.63*** (0.02)	-0.19 (0.15)	-0.42*** (0.07)
Independent	-0.87*** (0.03)	-0.60*** (0.16)	-0.63*** (0.07)
Weak Republican	-1.01*** (0.03)	-0.94*** (0.15)	-0.50*** (0.07)
Strong Republican	-1.51*** (0.02)	-1.27*** (0.14)	-0.91*** (0.08)
Education	0.03*** (0.003)	-0.01 (0.02)	-0.03** (0.01)
Income	-0.0003* (0.0001)	-0.001 (0.001)	-0.002** (0.001)
Black	0.37*** (0.03)	0.40* (0.16)	0.36*** (0.07)
Hispanic	0.28*** (0.03)	0.01 (0.16)	0.46*** (0.07)
Asian	0.20*** (0.05)	0.74* (0.35)	0.19 (0.19)
Age	-0.001 (0.0005)	0.001 (0.004)	-0.01*** (0.002)
County % Black 2010	-0.20* (0.08)	0.54 (0.53)	0.38 (0.23)
County % Hispanic 2010	0.03 (0.06)	0.23 (0.36)	0.38* (0.17)
County Med. Hsh. Inc. 2010	-0.0000 (0.0000)	0.0000 (0.0000)	-0.0000 (0.0000)
County % Poor 2010	0.003 (0.30)	1.01 (1.81)	-0.16 (0.92)
County % Unemployed 2010	5.15*** (0.90)	-6.40 (6.14)	2.24 (2.62)
County Δ % Unemployed 2010-16	3.56** (1.15)	-2.43 (7.88)	4.30 (3.51)
County % with BA 2010	0.97*** (0.13)	1.24 (0.79)	0.48 (0.40)
Number of plans (logged, in sd's)	0.001 (0.01)	-0.07 (0.06)	-0.004 (0.03)
Mean Premium (in sd's)	0.004 (0.01)	0.04 (0.07)	0.05 (0.03)
Mean Change in Premium (in sd's)	0.001 (0.01)	-0.19** (0.07)	-0.02 (0.04)
Constant	2.17*** (0.14)	2.87*** (0.70)	2.65*** (0.34)
Month FEs	Y	Y	Y
County-Level Demographics	Y	Y	Y
Clusters (Year)	3	3	3
Note:	*p<0.05; **p<0.01; ***p<0.001		

Table A17: This table presents the results of the full multi-level models fit to KFF respondents from 2015, 2016, and 2017 in which certain insurance market conditions (and various other independent variables) predict ACA favorability, measured on a 1 to 4 scale.