The Hostile Audience: The Effect of Access to Broadband Internet on Partisan Affect*

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Abstract

Over the last two decades, as the number of media choices available to consumers has exploded, so too have worries over self-selection into media audiences. Some fear greater apathy, others heightened polarization. In this paper, we shed light on the latter possibility. We identify the impact of access to broadband Internet on affective polarization by exploiting differences in broadband availability brought about by variation in state right-of-way regulations (ROW). We merge state-level regulation data with county-level broadband penetration data and a large-N sample of survey data from 2004 to 2008 and find that access to broadband Internet increases partisan hostility. The effect occurs in both years and is stable across levels of political interest. We also find that access to broadband Internet boosts partisans' consumption of partisan media, a likely cause of increased polarization.

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Over the past fifty years, partisans have come to increasingly dislike each other (Iyengar, Sood and Lelkes 2012), so much so, that today implicit partisan prejudice exceeds implicit racial prejudice (Iyengar and Westwood 2014; see also Chambers, Schlenker and Collisson 2013). Party cues now constrain social and interpersonal relations —partisans trust co-partisans more than supporters of the opposing party (Carlin and Love 2013; Iyengar and Westwood 2014; see also Hetherington and Rudolph 2014) and large proportions of both Republicans and Democrats are troubled by the prospect of a family member marrying a supporter of the main opposing party (Iyengar, Sood and Lelkes 2012; see also Huber and Malhotra 2013).[1]

Over the same period that partisan animus has been increasing, the reach of partisan information sources has been expanding. The broadcast news audience of 1975 could "choose" between three largely indistinguishable and devoutly non-partisan network newscasts. Today, aside from a broad array of nonpartisan news sources, including network news, viewers can also tune into "all news" partisan cable channels, partisan "news" shows on numerous other television channels, including two prominent shows on Comedy Central, or one of the countless partisan sources available online.

The reach of partisan media is not limited to discretionary exposure, i.e., those that choose that tune in. Increasingly, even the politically disinterested are exposed to non-trivial doses of partisan news. Online social networks today form the backbone of many Americans' daily information environment. And due to network partisan homophily (Halberstam and Knight 2014; Lewis, Gonzalez and Kaufman 2012), many apolitical individuals find themselves in networks with at least one politically active "friend" who is apt to recirculate news reports and commentary (see Halberstam and Knight 2014; Messing 2013). In fact, by some metrics, Facebook is now a major source of traffic to online news sites (Pew Research Center 2014).

[1] In this paper, we focus on affective polarization or interparty animus rather than ideological polarization. Evidence in favor of increasing ideological polarization in the electorate, however, is mixed (e.g., Abramowitz 2010; Fiorina, Abrams and Pope 2005).

While there are good reasons to believe that the new media environment has contributed to the growth in partisan animus, by facilitating access to partisan news, it is possible that enhanced consumer choice also sends in motion processes that weaken polarization. As choices for news have multiplied, so too have choices for entertainment. The increased availability of entertainment programming enables some to opt out of politics entirely (Prior 2007). But while it is undoubtedly true that forty or so years ago, when during portions of prime time there was nothing to see except news on network television, some people watched news even when they didn't want to, it is likely also true that some people didn't watch news then because it wasn't available at a time they wanted to see it, or available in a format, style, or ideological slant of their liking. Today, a vast buffet of news is available 24/7, both in and outside the house, on mobile phones, and tablets. Thus, the net impact of the increased number of news providers, and the greater empowerment of consumers is as yet mostly unknown.

In this paper we shed some light on this question. We examine whether better access to choice affects political attitudes. In particular, we investigate how access to broadband Internet affects partisan animus. Considerable evidence suggests that media consumption is strongly elastic, increasing sharply with better access. For instance, those with a broadband Internet connection spend considerably more time online—approximately 1300 additional minutes per month according to Hitt and Tambe (2007)—and spend more time reading and sharing news and opinions than those with dial-up connections (Rappoport, Kridel and Taylor 2002). We investigate whether this dramatic increase in consumption of content, including some political content, affects partisan animus.

We identify the causal impact of broadband access on affective polarization by exploiting differences in broadband availability brought about by variation in state right-of-way regulations (ROW), which significantly affect the cost of building Internet infrastructure and thus the price and availability of broadband access. Our results suggest that had all states adopted the least restrictive right-of-way regulations observed in the data, partisan animus would have been roughly 2 percentage points higher. We show that the estimates are robust. We also demonstrate that an alternative set of instruments for broadband availability (surface topography) yields very similar results. Lastly, we present some analyses suggesting that broadband access increases exposure to partisan information, which we take to be the most likely reason for why access to broadband polarizes partisans.

Broadband Internet, Exposure to Political Information, and Partisan Affect

In the Internet era, social scientists have rediscovered the concept of selective exposure, an idea that dates back to the classic studies on attitude change (e.g., Berelson and Steiner 1964; Festinger 1957; Klapper 1960; Lazarsfeld, Berelson and Gaudet 1948; McGuire 1968). Explosive growth in the number of media outlets and the declining cost of access to these choices means that consumers cannot possibly keep up with the increase in available content. Faced with this fire hose of information, people must be selective so as not to be overwhelmed. In selecting what political (and some apolitical) information to consume, partisans have been shown to use, among other things, cues about partisan congeniality (e.g., Iyengar and Hahn 2008; Iyengar et al. 2008; Stroud 2010). However, the evidence suggests that, on average, partisans have only weak preferences for congenial political information (Dvir-Gvirsman, Tsfati and Menchen-Trevino 2014; Garrett 2009; Gentzkow and Shapiro 2011; Prior 2012).

In the age of broadband even a small imbalance in the ratio of congenial to uncongenial political information can prove consequential. As already noted, media consumption is strongly elastic; moving from a dial-up connection to broadband produces a large increase in the amount of time spent online (Hitt and Tambe 2007; see also, Kolko 2010). And while moving to broadband likely causes some substitution in the kind of content people consume, notably a move from text to video, the primary effect of broadband is to increase the amount of media people consume (rather than what they consume). Put more colloquially, access to broadband primarily increases the size of the pie, without having much impact on the ratio of the individual slices. Assuming patterns of consumption remain roughly the same, any increase in consumption necessarily means greater exposure to imbalanced political information. Consider a person who consumes twice as much partisan over balanced media. Keeping the mix of congenial to uncongenial exposure the same, if the person doubles the consumption of total political information each week, the net imbalance also doubles. Hence, relatively small asymmetries in consumption behavior can be magnified by access to broadband.

Quite separate from imbalances in discretionary exposure are imbalances in inadvertent exposure. Studies suggest that the latter is also skewed in the direction of greater exposure to congenial information (Brundidge 2010; Wojcieszak and Mutz 2009). People frequently encounter political discourse during online discussions devoted to music, hobbies, movies, etc., and most of the discussion occurs among individuals with similar political views (Wojcieszak and Mutz 2009). Similarly, expected asymmetries in what news and opinion pieces are shared among "friends" on social networks (An, Quercia and Crowcroft 2014; Flaxman, Goel and Rao 2013) likely produce biases in the information flow within social networks. Thus, high speed Internet access, by facilitating online networking, and participation in general interest online discussion groups also contributes to "de facto" partisan selectivity.

The effect of greater exposure to imbalanced political information is straightforward to hypothesize: People are either persuaded to take more extreme positions on issues, or persuaded to dislike politicians of the opposing party and the people who support them. It also holds that those persuaded to adopt more extreme policy positions will, as a result, develop greater ill-will toward the opposing party.

Exposure to partisan media, however, can polarize audiences in another way —through priming and strengthening their partisan identities. Mere exposure to partisan media primes partisan identity (Knobloch-Westerwick and Kleinman 2012), and strengthens its salience (Horwitz and Nir 2014; Levendusky 2013*a*). Thus, even when a viewer tunes into Fox News to catch up on the latest celebrity scandal, her partisan identity is activated (for evidence that partisans prefer to get their soft news from partisan sources, see Iyengar and Hahn 2008). By this account, partisan media do not have to convince partisans about anything to be consequential, but they need to merely activate their social identity.

Either through persuasion, or by increasing the salience of partisan identity, greater exposure to partisan media is liable to polarize. Unsurprisingly then, some research suggests that voluntary exposure to partisan media causes partisans to trust the opposition less (Levendusky 2013*a*) and increases inter-party animosity (Garrett et al. 2014). Thus, one way broadband Internet may affect political attitudes is by increasing the net imbalance in political information that partisans consume. And as stated above, we expect this greater imbalance to lead to greater partisan animus.

Separately, access to broadband Internet may increase partisans' exposure to "balanced" political information. But even if access to broadband Internet only facilitated greater exposure to "balanced" information, it could still raise partisan animus. A long line of research shows that partisans engage in "motivated reasoning" —they interpret facts and events in a manner that supports their partisan beliefs (for reviews, see Kunda 1990; Lodge and Taber 2000; see also Lord, Ross and Lepper 1979). For instance, partisans think sources conveying uncongenial information are unfair (Vallone, Ross and Lepper 1985). Combined with the fact that in a polarized political environment, even centrist outlets often include abrasive comments from politicians, (see, for instance, Druckman, Peterson and Slothuus 2013) greater exposure to balanced media can increase inter-party animus (see for instance, Arceneaux, Johnson and Cryderman 2013). In all, if access to broadband Internet facilitates greater exposure to partisan rhetoric, either imbalanced or balanced, either discretionary or inadvertent, either via increasing exposure to partisan media channels or to "moderate" outlets, it is all but certain to increase partisan hostilities. Access to broadband Internet, however, may also set in motion processes that reduce affective polarization. Enticed by the plethora of entertainment options, people may become more likely to choose entertainment over news (see Prior 2007). But for a rise in selection into entertainment to offset substantial increase in consumption of all media, and actually reduce exposure to politics, media consumption habits need to undergo a dramatic change. While such dramatic change is likely infrequent, it is liable to be true for some people. Any such reduction in consumption of political news is liable to reduce polarization. Separately, exposure to balanced information may have different effects than suggested by Arceneaux, Johnson and Cryderman (2013). For instance, a study in which subjects were assigned to consonant or dissonant groups shows that exposure to opposing views under certain conditions can increase tolerance (Mutz 2002). In all, while there are strong reasons to think that access to broadband would polarize audiences, countervailing processes may well attenuate net polarization we observe.

Identifying the Effect of Broadband Internet on Affective Polarization

The demand for broadband Internet is likely a function of age, income, and education. But it also likely depends on some other (potentially unmeasured) variables. These variables may, in turn, explain partisan affect. To allay worries about such omitted variables what is needed is an instrumental variable (IV), a variable that causes broadband Internet access, but does not impact affective polarization through any other means.

Fortunately, there exists such an instrument. Section 253 of the Telecommunication Act of 1996 gave municipalities control over public rights-of-way used by telecommunication providers, e.g. use of ground beneath a public park to lay fiber optic cable (Day 2001). Soon after, many states passed laws enabling municipalities to regulate right-of-way, introducing significant variance in the degree to which they could impose fees or otherwise increase the costs faced by Internet service providers entering the local market. These state laws that vary municipal control over right-of-way are a well-known instrument for broadband access (Larcinese and Miner 2012; Suziedelyte 2012; Wallsten 2005). Larcinese and Miner (2012), for instance, find that an index of state regulation of right-of-way laws strongly predicts number of providers in a county, which, as we discuss later, is a good proxy for broadband uptake. Similarly, Wallsten (2005, p. 11) finds that "mandated access to rights-of-way can increase broadband penetration by .006 lines per-capita, or about 10 percent."

In this study, we also seek to exploit ROW regulations as an instrument for broadband access. We begin by establishing the strength and the validity of the instrument. More concretely, we use data from the Federal Communication Commission (FCC) on the number of broadband providers in a county (which we show later is a good proxy for broadband penetration) and regress this on an index of ROW regulations (Beyer and Kende 2003). We also present results in which we add key exogenous county-level features thought to predict broadband access (e.g., racial composition, income, population density, education) to the equation to strengthen the case that the variation in broadband access due to ROW is idiosyncratic.

Formally, our first stage model takes the following form:

$$X_{jk} = \alpha + \beta Z_k + \delta R_j + v_{jk} \tag{1}$$

where j indexes county and k states. X refers to the number of providers in the county j, while Z_k indicates state k's ROW score and $\mathtt{R}_j$ represents a matrix of county-level covariates, and v_{jk} is the error term. Whenever we use ROW scores, we include state-clustered robust standard errors.

Next, we assess the instrument's validity. In particular, we supplement the discussion in Larcinese and Miner (2012), who shed light on the validity of the ROW index. For instance, to assess the concern about less restrictive ROW regulations being adopted by more conservative states, we regress ROW on state-level estimates of ideology (we use measures developed by Tausanovitch and Warshaw 2013). We also regress ROW on median ideology of the state legislature using measures developed by Shor and McCarty (2011). In the same spirit, we check whether ROW laws were related to the party of the governor. Note, that in an effort to address similar kinds of concerns, Larcinese and Miner (2012) show that ROW is unrelated to Democratic vote share between 2004 and 2008, and 1992 and 1996.

To assess concerns that less stringent ROW laws may have been enacted in richer states, we regress an index of ROW laws on median state income. We find the two to be unrelated. We also test whether more educated states were likelier to adopt more liberal ROW laws. We again find the relationship to be weak. Lastly, to test whether partisan affect is endogenous to ROW laws, we test whether affect in 2000 (prior to widespread broadband adoption) predicts ROW laws. Once again, we find no relationship.

Finding the instrument to be strong and valid, we estimate the reduced form model by regressing an indicator of partisan affect on ROW and other exogenous covariates. For data on partisan affect we turn to the 2004 and 2008 National Annenberg Election Studies. (We describe the data and measures in greater detail in the next section.)

The reduced form model is as follows:

$$Y_{ijk} = \alpha + \beta Z_k + \delta R_j + v_{ijk} \tag{2}$$

where i tallies respondents, and Y_{ijk} indicates individual-level partisan affect. The reduced form estimates show that less restrictive right-of-way laws increase affective polarization. After presenting the reduced form estimates, we turn to results from the second stage of the instrumental variable regression using predicted values of broadband access.

In particular, we estimate the following model:

$$Y_{ijk} = \alpha + \hat{X}_{ijk} + \delta R_j + v_{ijk} \tag{3}$$

where $\hat{X}$ are the fitted values following the first stage regression. Results from the second stage regression converged with the reduced form estimates: access to broadband Internet increases partisan animus.

After presenting IV estimates of the effects of broadband availability on affective polarization, we carry out a series of robustness tests. We first test the robustness of our results to different specifications. Next, we present results using a second set of instruments: characteristics of the counties' terrain.

The cost of building broadband infrastructure is known to depend on terrain and weather (Andersen et al. 2012; Government Accountability Office 2006; Jaber 2013; Kolko 2010). For instance, the increased risk of flooding and higher summer temperatures (which disrupts cable heat dissipation) increases the costs of building and maintaining broadband infrastructure in low lying areas. New lines are more difficult to build on steep rather than flat land. Given the relationship between cost for building broadband infrastructure, scholars have used various geographical variables as instruments for broadband penetration, including a region's average elevation (Jaber 2013), average ground steepness (Kolko 2010), and even average number of lightning strikes per year (Andersen et al. 2012), all of which increase the cost of building broadband infrastructure. Hence, as a second set of instruments, we use data from the Economic Research Service's (ERS) Terrain Typography, which places counties into one of 21 landform categories, ranging from those in which building infrastructure is relatively easy to more difficult (expensive) terrains, and data from Kolko (2010), who uses the average slope of the terrain in an area as an instrument for broadband penetration. This second set of instruments helps alleviate concerns due to the small effective sample size (N=48) of the first stage ROW model.

Next, we test for heterogeneous effects. First, we expect the effect of broadband on polarization to have strengthened over time —political campaigns utilized the Internet far more in 2008 than in 2004 (Chadwick and Anstead 2008; Smith et al. 2009). However, one crucial limitation of our estimation strategy —IV estimates are LATE meaning that less restrictive ROW laws "encouraged" a different set of people to subscribe to broadband in 2008 than in 2004 —prevents us from saying much about the question of temporal variation in the effects of broadband.

Second, we test whether the effect of broadband penetration on polarization is concentrated among the politically interested. A number of scholars have argued that media proliferation will affect the more and less engaged differently. For instance, Prior (2007) argues that those uninterested in politics will likely tune-out of potentially polarizing information, and Arceneaux and Johnson (2013) argue that media proliferation will, on average, have minimal effect, as the highly interested tend to be stable in their attitudes, while those that are uninterested are not exposed to news.

Lastly, we test whether the data are consistent with what we think is the most plausible mechanism via which broadband access affects partisan polarization – broadband access changes exposure to partisan media. Many past studies have linked increased consumption of partisan media to increased polarization (Levendusky 2013*b*; Martin and Yurukoglu 2014; Stroud 2010, 2011; although see Arceneaux and Johnson 2013; Arceneaux, Johnson and Cryderman 2013). Using passively observed media data from comScore, and survey data from the 2004 and the 2012 American National Election Studies (ANES), we compare media consumption between those with broadband and those without. We find that those with broadband Internet access consumed far greater amounts of partisan media than those with dial-up connections.

Data and Measures

We use multiple sets of data for the analyses. Briefly, data on Right-of-Way laws come from an index of these laws compiled by Beyer and Kende (2003). The data on broadband access are from the Federal Communication Commission (FCC). For data on partisan affect, we use the 2004 and 2008 National Annenberg Election Studies (NAES). For media data, we turn to comScore.

Right-of-Way Laws

Fifty-two indicators of regulations on deployment of broadband, e.g., a cap on the fees municipalities can charge providers for ROW usages, and the provision of incentives for broadband deployment including low-cost loans to suppliers and tax incentives for broadband subscribers, were combined into a Right-of-Way index (Beyer and Kende 2003). Appendix A provides details on indicators and scoring. These laws were all enacted prior to 2002– two years prior to the collection of the survey data used here. The higher the value of the index, the less restrictive the ROW regulations, and the less costly it is to build Internet infrastructure.[2]

We map the ROW values in Figure 1. The distribution of ROW does not seem to follow any obvious pattern. For instance, rich states and poor states have very similar ROW scores, as do "red" and "blue" states. We more formally test whether ROW is related to ideology in the results section.

Broadband Access

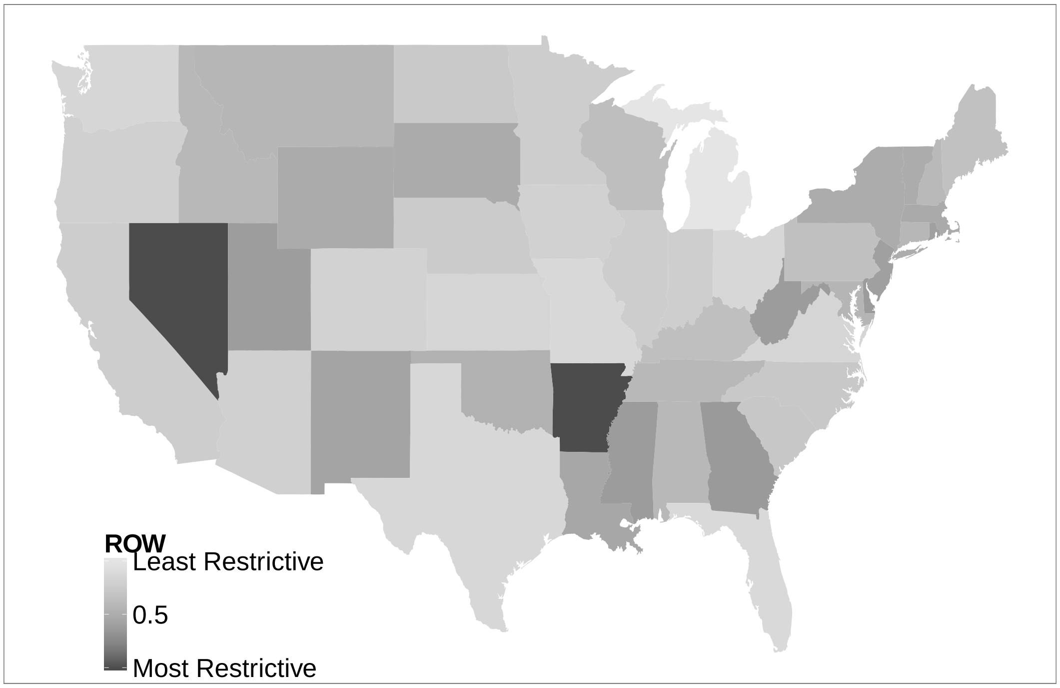
Since we do not have information about whether our survey respondents had a broadband subscription or not, we measure broadband access indirectly, via the total number of broadband providers in a respondent's county. The data on broadband providers are from the FCC, which keeps records of all high speed Internet providers with more than 250 subscribers in a state.[3] The FCC does not distinguish between DSL, cable, broadband, or satellite; broadband service providers are those that enable a transfer speed of at least 200 kb/s in one direction. For each survey-year, we match survey respondents in a county to the number of providers delivering broadband service to that county code.

The number of broadband providers in a given area is a well-known proxy for broadband

[2]We define ROW broadly, to include regulations that affect cost of deployment, as well as regulations that increase supply and demand. These three components are all very highly correlated (average $r = .96$). Results are substantively the same if we limit the ROW index to any one component.

[3]The data come from FCC Form 477, and are available at http://transition.fcc.gov/wcb/iatd/comp.html

Figure 1: Right-of-Way Score by State



penetration. A great deal of research shows that the number of broadband providers is a good measure of the number of broadband subscribers (see, for instance, Jaber 2013; Larcinese and Miner 2012). For instance, Larcinese and Miner (2012) find a strong correlation between the number of providers in a state and the proportion of households in the state with a broadband subscription.

A similar relationship between broadband availability and broadband penetration found at the state level obtains at the zip code level. Kolko (2010) uses survey data from Forrester Research to estimate the relationship between the number of providers in a zip code, and the proportion of households in a zip code with broadband. He finds that the probability of broadband subscription among respondents living in zip codes with only 1-3 providers is .22, while the probability of subscription among those living in zip codes with 20 or more providers is more than doubled (nearly .45). We present a similar analysis, estimating the relationship between the number of providers in a county and the percent of respondents with broadband subscriptions, using data from the comScore 2004 panel. As is clear from

Figure C1, the relationship is monotonic, and roughly linear.

We further replicate the analyses with data at the census tract level from the FCC. The FCC does not provide data on the exact number of subscribers at a lower level of geographic aggregation than the state. However, in December, 2008, the FCC for the first time provided the number of broadband lines per 1000 residents in each census tract split into a few categories: zero to 200 connections, 200 to 400 connections, 400 to 600 connections, 600 to 800 connections, and more than 800 connections. We plot these values against the FCC's tally of broadband providers per census tract (logged) and fit a cubic spline and its 95 percent confidence interval. As shown in Figure 2, the number of subscribers linearly increases with the number of providers in a census tract increase.

Partisan Affect

We rely on data from rolling cross-section surveys conducted as part of the 2004 NAES ($N = 98,711$) and the 2008 NAES ($N = 57,967$) to measure partisan affect.[4] Both surveys are based on a random sample of the US population interviewed by telephone over the course of the presidential campaign. Interviewing occurred between December 2003 and November of 2004 in the 2004 study, and between December 2007 and November 2008 in the case of the 2008 study.

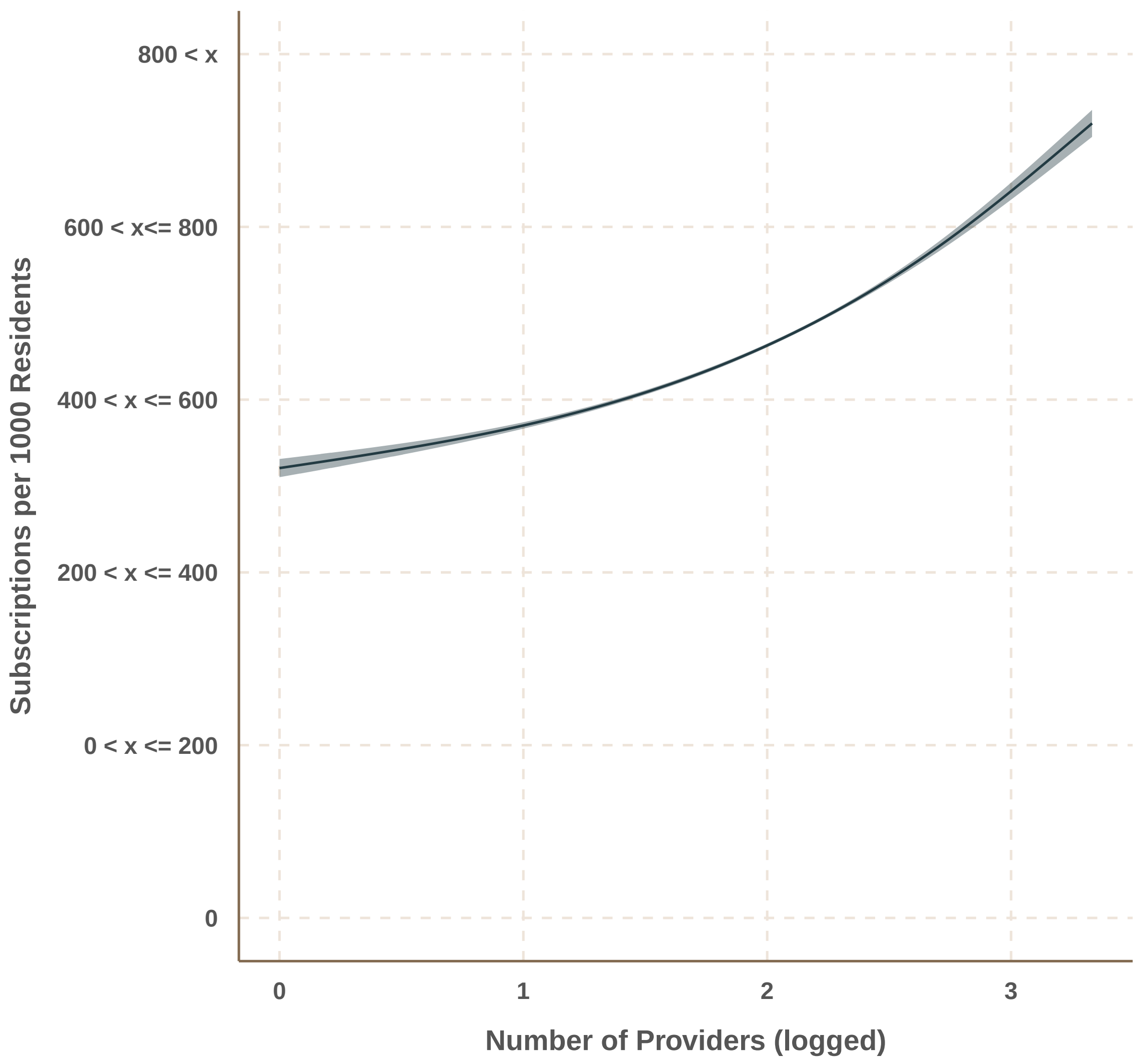
Partisan affect was measured as the difference in feelings toward the in-party and out-party presidential candidate.[5] We define partisans to include leaners, and omit pure independents from all analyses (Keith et al. 1992).[6] In both 2004 and 2008, respondents were

[4]These are the sample sizes before we exclude Independents.

[5]We think affect towards candidates is a very good proxy for partisan affect for two reasons - a) affect towards partisan candidates is strongly endogenous to partisan affiliation (Bartels 2002; Greene 1999), b) even if affect towards partisan candidates was shaped by forces other than partisanship, any growth in candidate affect is still liable to feed into feelings towards parties; people are liable to think badly of parties that nominate candidates they dislike. Lastly, we use data from the 2012 American National Election Studies which measured affect towards the major party candidates and parties and correlate the two. The average correlation was .80.

[6]Excluding Independents from our analysis does not affect our results. Regressing a variable indicating that identifying as an independent versus a partisan on ROW (and county-level controls) reveals that the two variables are unrelated: ($b = -.02$, $p = .59$).

Figure 2: Relationship and 95 percent Confidence Band between Number of Providers and Proportion with Broadband Internet within census tracts in 2008, smoothed using a cubic spline estimator



asked to rate the candidates on favorability and a set of traits on a 11 point scale. In 2004, respondents rated George W. Bush and John Kerry on favorability, and the degree to which they viewed each of the candidates as trustworthy, knowledgeable, reckless (which was reverse coded), cared about "people like [them]," and shared the respondent's values. In 2008, respondents rated John McCain and Barack Obama on favorability, leadership, trustworthiness, experience, and judgment. (See Appendix B for exact question wording and response options.) The average inter-item correlation was .58 (Cronbach's α =.96) in 2004, and .49 (Cronbach's α =.91) in 2008. We took the average of the differences between the in- and out-party candidate ratings, and rescaled it to lie between 0 (out-party candidate rated at 10 on each item and in-party candidate rated at 0 on each item) and 1 (in-party candidate rated at 10 on each item and out-party candidate rated at 0 on each item).[7]

Control Variables

We further included a number of county-level indicators using data from ICPSR's County Characteristics' (2000-2007)[8] file as controls, including the unemployment rate, median age, the male-to-female ratio, percent black, percent white, census region, whether the county is classified as low education, median income, and the population density. In robustness tests we also include several variables as individual-level controls, we used age (divided into quartiles), income (also divided into quartiles), gender, and education (coded as high school or less, some college, Bachelor's degree, post-graduate education). We also created dummy variables that tracked missing values on each of the variables.

[7]Since those that are below the midpoint of this measure do not harbor any out-party animosity, we could also recode those scores as 0. Doing so slightly increases the size of the coefficients in the following models, while maintaining their significance. We stick to the simple difference score as the results are more conservative and the coding is consistent with past studies.

[8]http://www.icpsr.umich.edu/icpsrweb/DSDR/studies/ 20660

Results

Validity and Power of the Instrument

We start by presenting results of the first stage model, predicting the number of providers per county using the ROW index (Column 1, Table 1).[9] The relationship between the ROW Index and number of providers is positive and statistically significant: Loosening ROW restrictions by about 10 percent yields a .5 percent increase in the number of providers (b = .053, s.e.= .018).[10] Other covariates in the model are plausibly related to broadband penetration. For instance, counties with a higher median income or that are more densely populated have more broadband providers. The variation in broadband access caused by ROW is above and beyond these well-known antecedents of broadband diffusion.[11] Although the F-statistic on the ROW coefficient is only about 8.4, because this is a two sample design, concerns about consistency are much less pronounced (Inoue and Solon 2010) (and we find similar results with our second set of instruments).

Next, we conduct three sets of validation checks, testing the bivariate relationship between ROW laws and a number of state characteristics. First, we assess whether worries about *less restrictive* ROW laws being enacted in conservative states is well-founded. In particular, we regress a measure of state-level ideology (Tausanovitch and Warshaw 2013) on ROW. We find no relationship ($b = -.004, p = .88$). Next, we test whether less restrictive ROW laws were more likely to be enacted under Republican governors. Lacking exact date for when ROW laws were enacted, we track the entire time span between 1996 (the year ROW laws began being passed) and 2008. For each year in that range, we regressed party control (Republican=1) on ROW, and find no significant relationship in any of the years (Mean

[9]Because of one outlier (Michigan, see Figure A1) we take the log of the index; the results are substantively unaffected if we do not transform ROW and remove Michigan from the analyses.

[10]These results are the same if we use the merged multilevel dataset.

[11]Another relevant question pertains to the type of people that are affected by ROW laws. Because we do not have individual-level information broadband subscription information, we can only explore whether ROW laws affect certain types of counties, an analysis which we explore in the appendix. However, people that do not have broadband tend to be less educated, older and poorer.

$b = .19$, Mean $p = .58$). Finally, we test whether ROW laws are related to the ideology of the legislature in each state. We use a measure of state legislative ideology (Shor and McCarty 2011), and, for each year between 1996 and 2008, we regress median ideology of each state senate and house on ROW. We find no relationship for upper chambers (Mean $b = .07$, Mean $p = .39$) or lower chambers (Mean $b = .11$, Mean $p = .18$).

Another prominent worry is more expansionary ROW laws were more likely to be enacted in richer states. Once again we find little grounds for such worry. Regressing median income (logged) on ROW yields coefficients that are very close to 0 in all years from 1996 to 2008 (Mean $b = .02$, Mean $p = .38$). Similarly, we test whether ROW laws were related to state-level education. Using census data from 1990 and 2000, we regress the percentage of the state population with a bachelors degree or higher on ROW. The relationship was small and insignificant in both years ($b = .03$, $p = .44$ in 1990 and $b = .03$, $p = .28$ in 2000). Additionally, state-level affective polarization in 2000 (estimated using the 2000 NAES) was not related to ROW ($b = 0.005$,$p = .25$). In all, the results of these analyses are consistent with the assumption that our instrument is valid.

Effect of Broadband Access on Partisan Affect

We next turn to results from the reduced form models, modeling affective polarization as a function of the ROW index and exogenous county and individual-level covariates. As the second column of Table 1 shows, less restrictive ROW laws cause higher levels of affective polarization ($b = .003$) with a 10 percent increase in the restrictiveness of ROW laws causing a .03 percent increase in affective polarization. In other words, if all states were to go from their current value to the least restrictive right-of-way law observed in the data set, partisan animus would increase by roughly 2%, from .65 to about .67.

The IV-based estimates are in line with our reduced form estimates (see Column 3, Table 1). The number of broadband providers in a county increases inter-party animus ($b = .03$, s.e.$= .01$). Translating the logged independent variable into a more intuitive

Table 1: First-Stage, Reduced Form, and IV Estimates Predicting the effects of Broadband Penetration on Affective Polarization

	First Stage	Reduced Form	IV Estimates
ROW Index (logged)	.053*	.003*	
	(.018)	(.002)	
# of Providers (logged)			.032*
			(.014)
Median Income (logged)	.987*	.015*	−.013
	(.058)	(.005)	(.014)
Population Density	2.406*	.048*	.018
	(1.096)	(.007)	(.018)
Unemployment Rate	.014	−.002	−.001
	(.007)	(.001)	(.001)
Low Education County	.028	−.016*	−.018*
	(.023)	(.003)	(.004)
Percent Male	−1.842*	−.083	.084
	(.363)	(.079)	(.107)
Percent White	−.323*	.069*	.100*
	(.115)	(.017)	(.024)
Percent Black	.218	.116*	.104*
	(.134)	(.020)	(.024)
West	−.036	.030*	.027*
	(.049)	(.003)	(.005)
South	.066	.019*	.019*
	(.056)	(.004)	(.004)
Midwest	−.142*	.011*	.015*
	(.055)	(.003)	(.004)
Year: 2008	.821*	−.064*	−.083*
	(.021)	(.002)	(.009)
Intercept	−8.222*	.481*	.628*
	(.715)	(.077)	(.110)
R^2	.770	.034	.032
Adj. R^2	.769	.034	.031
Num. obs.	6034	114803	114803
RMSE	.263	.183	.183

$^*p < 0.05$

metric, increasing the number of providers in a county by 10 percent yields a .003 point increase in affective polarization. Since the average number of providers in a county increased by 32 percent between 2000 and 2004 and 64 percent between 2004 and 2008, our model implies that broadband expansion increased polarization by .01 (between 2000 and 2004) and .02 points (between 2004 and 2008). Moving from a county with the fewest number of providers to a county with the highest number of providers increases affective polarization by roughly .07. Our estimate of the impact of broadband expansion is half as large as the effect of political interest, which is associated with about a .14 point increase in affective polarization.[12]

Robustness Checks

To test the robustness of the estimate to different specifications, we estimate three other reduced form and IV models - a) a model without any controls, b) a model including only the individual-level controls, and c) a model including both county-level and individual level controls. Since controls should not affect the point estimates of a randomly assigned instrument, our confidence in the instrument is strengthened when we compare the columns with and without the controls (including individual-level controls) from the reduced form models (see Table D1)–the coefficients are nearly identical. Similarly, our IV results proved robust to these vastly different specifications (see Table D2): coefficients from the individual and county-level covariate models were almost identical to our preferred specification while estimates from the bivariate and individual covariate-only models were slightly larger.

Next, we use a second set of instruments that capture environmental impacts on broadband penetration. Since "flat terrain constitutes good geography for telecommunications deployment" (Government Accountability Office 2006, p. 19), we use measures that capture terrain. First, we the Economic Research Service's terrain typology, which classifies terrain into 21 categories, ranging from flat plains to high mountains. Similarly, we follow Kolko

[12]As estimated by a model similar to Column 3 in Table D2, but including political interest as a covariate

(2010), and use the average slope of the terrain within a county. The first-stage, reduced form, and IV results appear in the Appendix.[13]

Consistent with our expectations, broadband penetration is highest on flat plains (the omitted category in the regression), and significantly so in over half the categories (Column 1, Table D3). An F-test indicates that the instruments are not weak: F(2, 5932)=10.16. The reduced form estimates show that a number of the terrain-related dummy variables are significantly correlated with affective polarization (Column 2, Table D3). The IV estimates of the effects of broadband access are slightly smaller than in the case of the ROW instrument (b = .02, s.e.= .005, $p < .001$; Column 3, Table D3), but in the same direction and significant. Similarly, there are fewer broadband providers when terrain is steeper (Column 1, Table D4). The IV estimates from the slope model also indicate that the broadband penetration significantly increases affective polarization ((b = .015, s.e.= .006, $p < .05$; Column 3, Table D4).

Combining the three instruments (ROW, terrain, and slope) into one model gives an effect of (b = .02,state-clustered robust s.e.= .01, $p < .05$). This effect represents the weighted-average of the upper-bound estimate from the ROW model and the lower-bound estimate of the terrain model. It indicates that counties with the fewest number of providers are roughly 4 points less polarized than providers with the highest number of providers.

Heterogeneous Effects

To test whether ROW effects are significant in both 2008 and 2004, and whether ROW effects are larger among individuals more interested in politics, we generate the IV estimates of the interaction effect between year and the logged number of providers, on the one hand, and R's political interest and the logged number of providers, on the other, on affective polarization. To estimate the former interaction effect in the first stage regression, we predict the logged number of providers in a county from a model that includes ROW x year

[13] Since these variables are at the county-level, we present county-clustered standard errors.

Table 2: Does the effect of broadband vary by year or by political interest?

	Model 1	Model 2
# of Providers (logged) * Year	.005	
	(.042)	
# of Providers (logged) * Political Interest		.022
		(.050)
Political Interest		.099
		(.099)
Year: 2008	−.093	−.084*
	(.095)	(.010)
# of Providers (logged)	.030	.012
	(.016)	(.037)
Median Income (logged)	−.013	−.015
	(.014)	(.015)
Population Density	.018	.011
	(.018)	(.018)
Unemployment Rate	−.001	−.001
	(.001)	(.001)
Low Education County	−.017*	−.011*
	(.004)	(.005)
Percent Male	.087	.087
	(.125)	(.115)
Percent White	.100*	.088*
	(.024)	(.024)
Percent Black	.103*	.093*
	(.026)	(.023)
West	.026*	.023*
	(.006)	(.005)
South	.019*	.017*
	(.005)	(.004)
Midwest	.014*	.015*
	(.004)	(.003)
Intercept	.629*	.596*
	(.112)	(.150)
R^2	.032	.073
Adj. R^2	.031	.073
Num. obs.	114803	98374
RMSE	.183	.179

$^*p < 0.05$

interaction. In the second stage, we include the predicted values of broadband penetration x year interaction and the county-level covariates to predict affective polarization. To estimate the interaction effect between political interest and ROW, we follow the same steps, but substitute political interest for year.

The IV estimates from these models appear in Table 2.[14] The effect of broadband penetration remains are stable between 2004 and 2008. In addition, the effects were uniform across different levels of political interest.

[14]The N in the second column is lower due to missing values on the political interest variable.

Causal Mechanism

As we note above, we suspect that it is increased exposure to partisan information that accounts for the observed effect of broadband access on affective polarization. Here we present some analyses that test whether moving to broadband actually increases exposure to partisan media. For reasons to do with limitations of data, in this section we move away from instrumental variable estimates and focus on comparing news media consumption of respondents using broadband or dial-up connection. We use coarsened exact matching (CEM) (Iacus, King and Porro 2012) and multiple regression to account for differences between people with broadband and dial-up.

Our first set of analyses relies on 2004 microdata from comScore. The dataset is a random sample of 50,000 panelists in the United States, who allowed the company to track their browsing behavior in exchange for various rewards, including cash and computer software. Although self-selected, the sample is fairly representative of Internet users in terms of age, education, income, and geographic region (De los Santos, Hortaçsu and Wildenbeest 2012; Gentzkow and Shapiro 2011). The major advantage of the comScore data is that web browsing is observed behaviorally rather than through unreliable self-reports (Prior 2009). The disadvantage is that we do not know the partisanship of the respondents. Thus, we are only able to check the extent to which consumption of partisan media varies between respondents with broadband and dial-up connections.

We begin by downloading browsing data for a list of 400 popular news websites used by Gentzkow and Shapiro (Gentzkow and Shapiro 2011). This list includes popular news sites such as nytimes.com, cnn.com, etc. and also important political sites such as democrats.org and votesmart.org. In all, these websites account for a large proportion of the traffic to news sites. We then merge the browsing data with Gentzkow and Shapiro's estimates of website ideology. Next we use a simple cut off to categorize whether a website is partisan or not. (Changing the cut off has little impact on our results). Specifically we classify sites that have scores less than $-.2$ as left-leaning and scores of over .2 as right-leaning. This removes sites such as abcnews.com, economist.com etc. but keeps sites such as msnbc.com, foxnews.com, etc. Without adjusting for covariates, we find that respondents with broadband access consumed on average twice as much content from partisan media than those with dial-up access.[15] Controlling for the entire battery of covariates available in the data: age of the oldest member in the household, household size, number of children in the household, racial background, and country of origin has little impact on the coefficient.

We next use CEM along with multiple regression to compare the frequency with which respondents with a dial-up Internet connection and broadband connection visit partisan websites. (See Appendix E1 for regression results from non-matched data.) CEM is a nonparametric data pre-processing algorithm that reduces imbalance on a set of covariates between, in our case, those with broadband Internet and those with only dial-up. We matched respondents on the entire list, and regressed net exposure to partisan media on the covariates with a weight vector obtained from CEM. The coefficients were again largely unchanged. The covariate-adjusted scores are plotted in Figure 3.

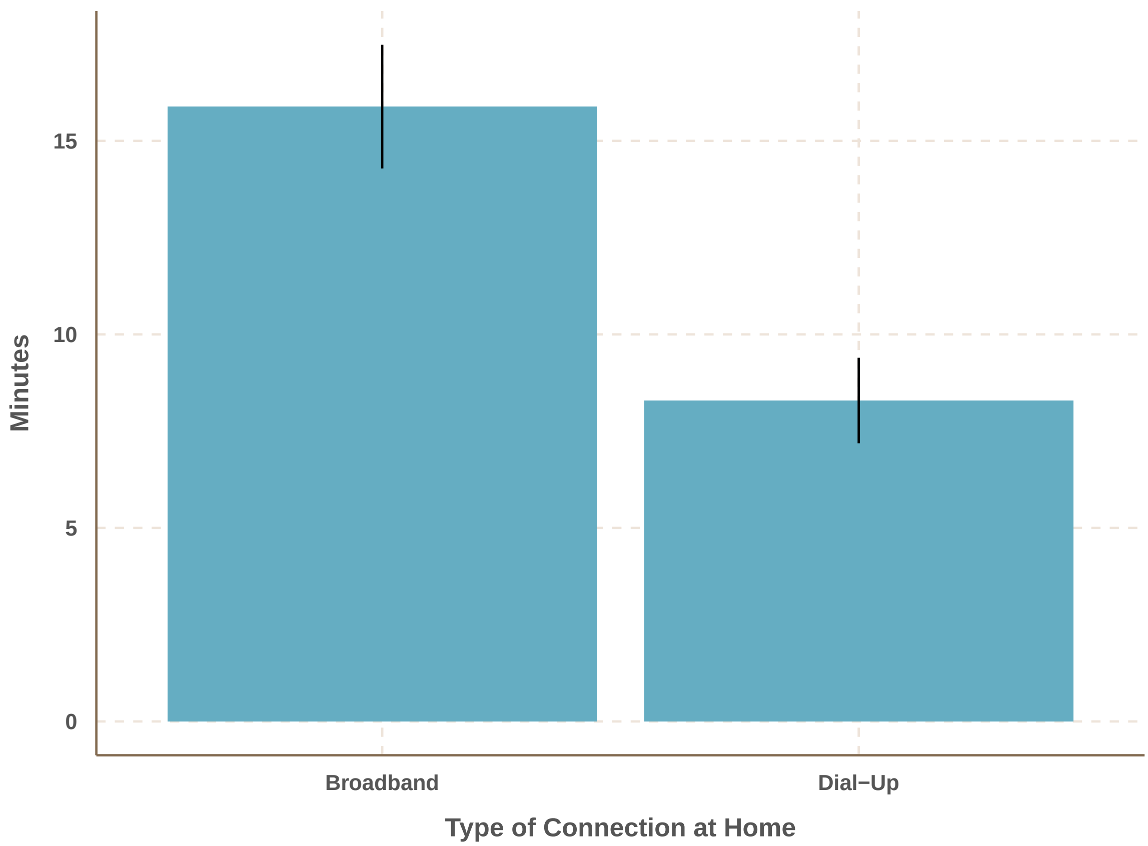
Our second set of analyses is based on data from the 2012 ANES, one of the few publicly available survey datasets to contain indicators of type of Internet access, and self-reported media use measures that have been shown to be reasonably valid and reliable (Dilliplane, Goldman and Mutz 2013; Goldman, Mutz and Dilliplane 2013; LaCour and Vavreck 2014),[16] and a measure of partisan identification. In the 2012 ANES, respondents were asked to indicate if they had visited a website from a long list of websites (Dilliplane, Goldman and Mutz 2013). From the list of websites, we scored Huffington Post and msnbc.com as left-leaning, and Drudge Report and foxnews.com as right-leaning. Next we use CEM, matching respondents on income, age, gender, race, political interest, and education.[17] The matching

[15]Given the skew in visitation patterns, one can also look at medians instead of means. There again we see the same pattern: the median visitation to partisan websites by those with broadband access is roughly double that of those with dial-up connections.

[16]Note though that even if media self-reports were biased (Prior 2013), our inferences are only at risk if error in self-reports is correlated with type of Internet access that the respondent has.

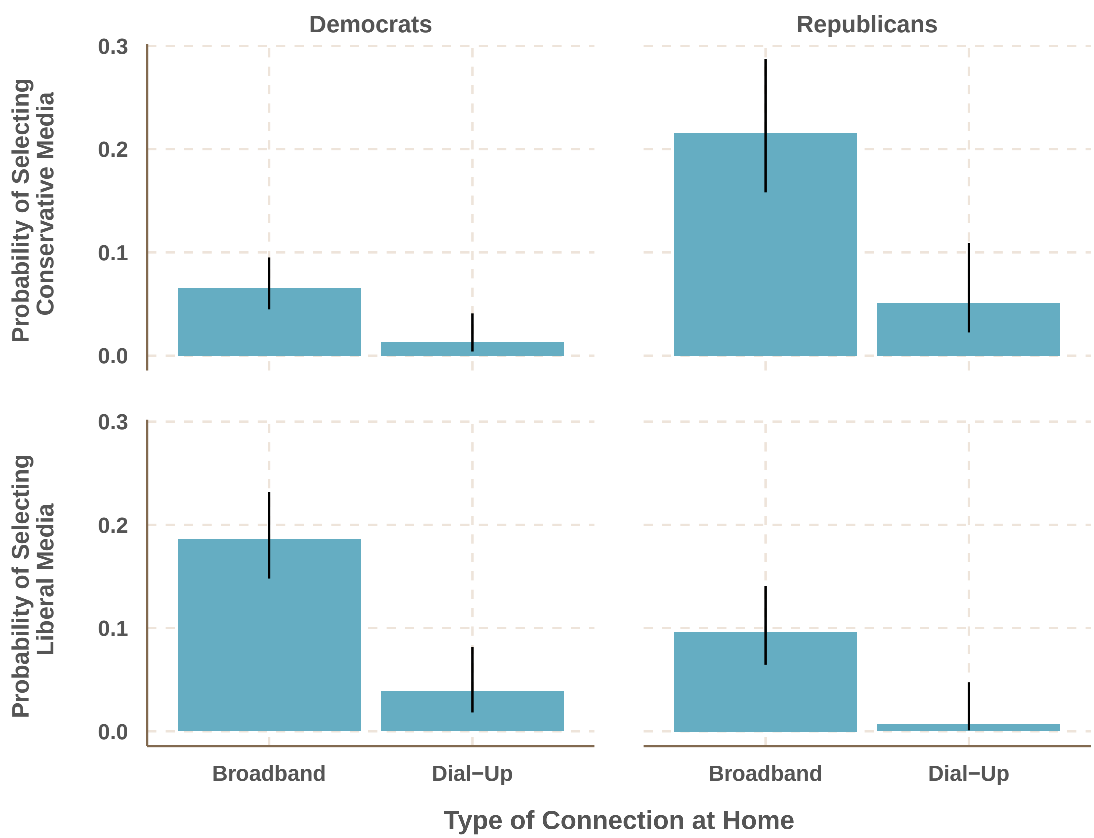
[17]See Table E2 for balance statistics before and after matching. There is some concern that, since, by 2012, those with dial-up were so unique, that matching cannot ensure exchangeability between groups. In Appendix E, we replicate the analysis with Pew data from 2004, when a far larger share of the US population

Figure 3: Visitation to Partisan Websites by Internet Connection



analysis (See Figure 4) shows that 19 percent of Democrats with broadband (compared to 3% with dial-up) say they "regularly" encounter liberal media, and 20 percent of Republicans with broadband (compared to 8% with dial-up) say they "regularly" encounter conservative media. (The results are substantively similar if we test for differences in exposure, and control for covariates in a traditional regression analysis instead; see Appendix E3). In short, broadband access does facilitate exposure to partisan media.[18]

Figure 4: Probability of Visiting Partisan Websites by Internet Connection



The results from both studies suggest that people with broadband access consume a lot more partisan media —especially from sources congenial to their partisanship —than those with dial-up connections. While there are still concerns about selection bias, the sheer size

was still connected to the Internet via dial-up.

[18]Broadband access also is associated with increased exposure to cross-cutting media, but the far larger relationship is exposure to reinforcing media.

of the differences, and their robustness suggests that the differences are real.

Aside from the data assembled here, there is additional evidence consistent with the claims we advance. Jaber (2013), for instance, finds that access to broadband Internet increases political knowledge by about 3.5%. While one can think of numerous mechanisms behind the effect, the most obvious and likely mechanism is increased exposure to public affairs programming. In all, the available data strongly suggest that access to broadband Internet increases exposure to partisan media and news programming. And we think this increase in exposure to partisan programming in particular, and public affairs programming more generally, explains the relationship between access to broadband Internet and affective polarization.

Discussion

Both the supply and the demand for partisan media are considerably greater today than a decade ago. These radical changes have naturally attracted a great deal of scholarly attention. Our study contributes to the burgeoning literature addressing the political implications of changes in the media environment. To overcome concerns over the endogeneity of media consumption to political attitudes, scholars have tried to establish the causal effect of access on attitudes and behaviors, exploiting idiosyncratic variation in the media menus of similar people. For instance, some studies have exploited geographical variation in the introduction of Fox News to estimate the impact of its introduction (not consumption) on voter preferences (DellaVigna and Kaplan 2007; Hopkins and Ladd 2014). Others have exploited the quasi-random location of various partisan channels on the cable menu (Martin and Yurukoglu 2014).

In this paper, we have pursued a similar strategy. We used exogenous variation in access to broadband Internet stemming from differences in right-of-way laws, which significantly boosts access to content, to identify the impact of broadband access on partisan polarization.

We found that access to broadband Internet polarizes rank and file partisans, and the effect amounts to about half the effect partisans political interest.

Although we find that the introduction of broadband has contributed to the rise in affective polarization, we do not think broadband is the only, or perhaps even the primary, cause of the rise in partisan ill-will (Iyengar, Sood and Lelkes 2012). Affective polarization began to increase at least two decades before widespread Internet use. Instead, our claim is that the new media environment exacerbates already rising tensions. The data suggest that access to broadband Internet heightens partisan animus by increasing partisans' exposure to imbalanced partisan rhetoric. Despite the possibility of other mechanisms that mute or neutralize any positive effects of broadband on polarization, we have documented such effects.

While some scholars have concluded – on the basis of data showing only limited exposure to partisan news (for a review, see, Prior 2012) – that partisan media cannot be consequential, we think this verdict is premature and based on insufficient attention to countervailing evidence. For instance, DellaVigna and Kaplan (2007), Hopkins and Ladd (2014), Martin and Yurukoglu (2014), Levendusky (2013*b*), all find that minor changes to a large menu of media choices exert a substantively significant impact on preferences and attitudes.

There are a variety of potential explanations for why relatively small doses of exposure to partisan media can add up to meaningful effects. For one, just because exposure is 'small' doesn't necessarily mean that it is inconsequential. Martin and Yurukoglu (2014), for instance, estimate that watching 4 minutes of Fox News a week is enough to increase the odds of voting for a Republican presidential candidate by .9 percent. Second, as we noted at the outset, even a small partisan imbalance in media exposure, on the order of magnitude found in Gentzkow and Shapiro (2011), Dvir-Gvirsman, Tsfati and Menchen-Trevino (2014) and Flaxman, Goel and Rao (2013), when accumulated over long spans of time, is liable to have substantial effects.

Finally, it may well be that preferences for agreeable partisan content are stronger than what previous research has suggested. Many of the studies investigating selective exposure code media content and consumption at the level of media outlets. This can understate preferences for partisan content. Prominent media outlets categorized as non-partisan, such as the *New York Times* and the *Washington Post* do carry ideological content, although they deliver a wide array of ideological perspectives. Secondly, pooling across various kinds of news stories may understate the strength of preferences for partisan congenial content. For instance, using passively observed data from 1.2 million users (2.3 billion page views), Flaxman, Goel and Rao (2013) find that ideological segregation tends to be much smaller for general news stories than for opinion stories. All told, the extant literature may greatly understate the degree of imbalance in consumers' exposure to congenial over uncongenial information.

Regardless of the strength of media preferences, the fact that greater access to choice causes partisan animus fits well with evidence from some other studies. In many ways the fact that access to broadband Internet causes polarization complements the finding that the introduction of broadcast television reduced polarization (Campante and Hojman 2013). As the political information environment became homogenously non-partisan, as a result of the introduction of national network news, partisans' attitudes moderated. Today, as a result of greater access to partisan news, partisan animus has increased.

In closing, this study shows that the new media environment has contributed to increased partisan animus and that greater exposure to biased news sources is the likely cause. As Americans get better access to providers of partisan information, we can anticipate more "fear and loathing" across party lines.

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Appendix A: ROW Index

The ROW Index was created by first creating scores for each type of regulation–deployment, supply-side, and demand side–by summing across items. Regulation scores were weighted by their relative importance to broadband deployment (4,2,1, respectively) and summed. TechNet weighed deployment regulations most heavily as the impact of the other two categories are contingent on deployment regulations; demand side regulations are weighed least heavily, as the organization believed it to be an indirect route to increased broadband deployment. The results are roughly the same if weights are omitted or if we use only one type of regulation as the instrument.

Deployment Regulations

1. Limit on the number of days that a municipality can take to process a rights-of- way permit request? (Yes=1; No=0)

2. If so, maximum number of days the municipality has to act on the permit? ($\leq$45 days =1;>45 days =0.5)

3. Streamlined resolution of rights-of-way permit disputes? (Yes=1; No=0)

4. Standardized permit application for all municipalities? (Yes=1; No=0)

5. State authority to collect and disburse rights-of-way fees? (Yes=1; No=0)

6. Prohibition on local regulations setting requirements unrelated to rights-of-way usage? (Yes=1; No=0)

7. Does the state encourage coordinated rights-of-way construction by multiple providers? (Yes=1; No=0)

8. Limit on the fees that municipalities can charge for rights-of-way usage? (Yes=1; No=0)

9. Prohibition on municipalities charging in-kind compensation in return for rights-of-way access? (Yes=1; No=0)
10. Rebate on rights-of-way fees for the deployment of a broadband network? (Yes=1; No=0)

Supply-Side Promotion Regulations

1. Formal state plan to increase broadband deployment? (Yes=1, No=0)
2. Dedicated state agency to coordinate broadband deployment? (Yes=1; No=0)
3. Databases or maps of existing broadband deployment? (Yes=1; No=0)
4. Limits to municipal deployment of broadband services? (Yes=0; No=1)
5. If so, are municipalities limited to offering wholesale services? (Yes=1; No=0)
6. State-owned backbone network? (Yes=1; No=0)
7. Does the state aggregate demand for broadband? (Yes=1; No=0)
8. Does the state maintain multiple broadband networks to the same locations (to ensure backup connectivity)? (Yes=1; No=0)
9. Does the state lease its broadband networks to private suppliers? (Yes=1; No=0)
10. Provision of grants to suppliers for broadband deployment? (Yes=1; No=0)
11. Are the grants targeted to deployment in underserved/rural areas? (Yes=1; No=0)
12. Loans to suppliers for broadband deployment? (Yes=1; No=0)
13. Are the loans targeted to deployment in underserved/rural areas? (Yes=1; No=0)
14. Tax incentives to suppliers for broadband deployment? (Yes=1; No=0)

15. Are the tax incentives targeted to deployment in underserved/rural areas? (Yes=1; No=0)

Demand-Side Promotion Regulations

1. Discounts to public sector users for broadband access? (Yes=1, No=0)
2. Are the discounts targeted to public-sector users in underserved/rural areas? (Yes=1; No=0)
3. Grants to public-sector users for broadband access? (Yes=1; No=0)
4. Are the grants targeted to public-sector users in underserved/rural areas? (Yes=1; No=0)
5. Loans to public-sector users for broadband access? (Yes=1; No=0)
6. Are the loans targeted to public-sector users in underserved/rural areas? (Yes=1; No=0)
7. Does the state provide access to state-owned broadband networks for private sector end-users? (Yes=1; No=0)
8. Does the state have a digital divide program? (Yes=1; No=0)
9. Discounts to private sector for broadband access? (Yes=1; No=0)
10. Are the discounts targeted to private-sector users in underserved/rural areas? (Yes=1; No=0)
11. Grants to private sector for broadband access? (Yes=1; No=0)
12. Are the grants targeted to private-sector users in underserved/rural areas? (Yes=1; No=0)

13. Loans to private sector for broadband access? (Yes=1; No=0)

14. Are the loans targeted to private-sector users in underserved/rural areas? (Yes=1; No=0)

15. Tax incentives to the private sector for broadband access? (Yes=1; No=0)

16. Are the tax incentives targeted to private-sector users in underserved/rural areas? (Yes=1; No=0)

17. Does the state encourage public-private consortiums for the purpose of offering or deploying broadband services? (Yes=1; No=0)

18. Does the state provide financial support to developers of broadband applications? (Yes=1; No=0)

19. Does the state have a tele-medicine program for state health institutes? (Yes=1; No=0)

20. Does the state have a distance learning program for state educational institutes? (Yes=1; No=0)

21. Does the state have any public safety/homeland security broadband applications? (Yes=1; No=0)

22. Does the state have tele-work programs for government employees using broadband?(Yes=1; No=0)

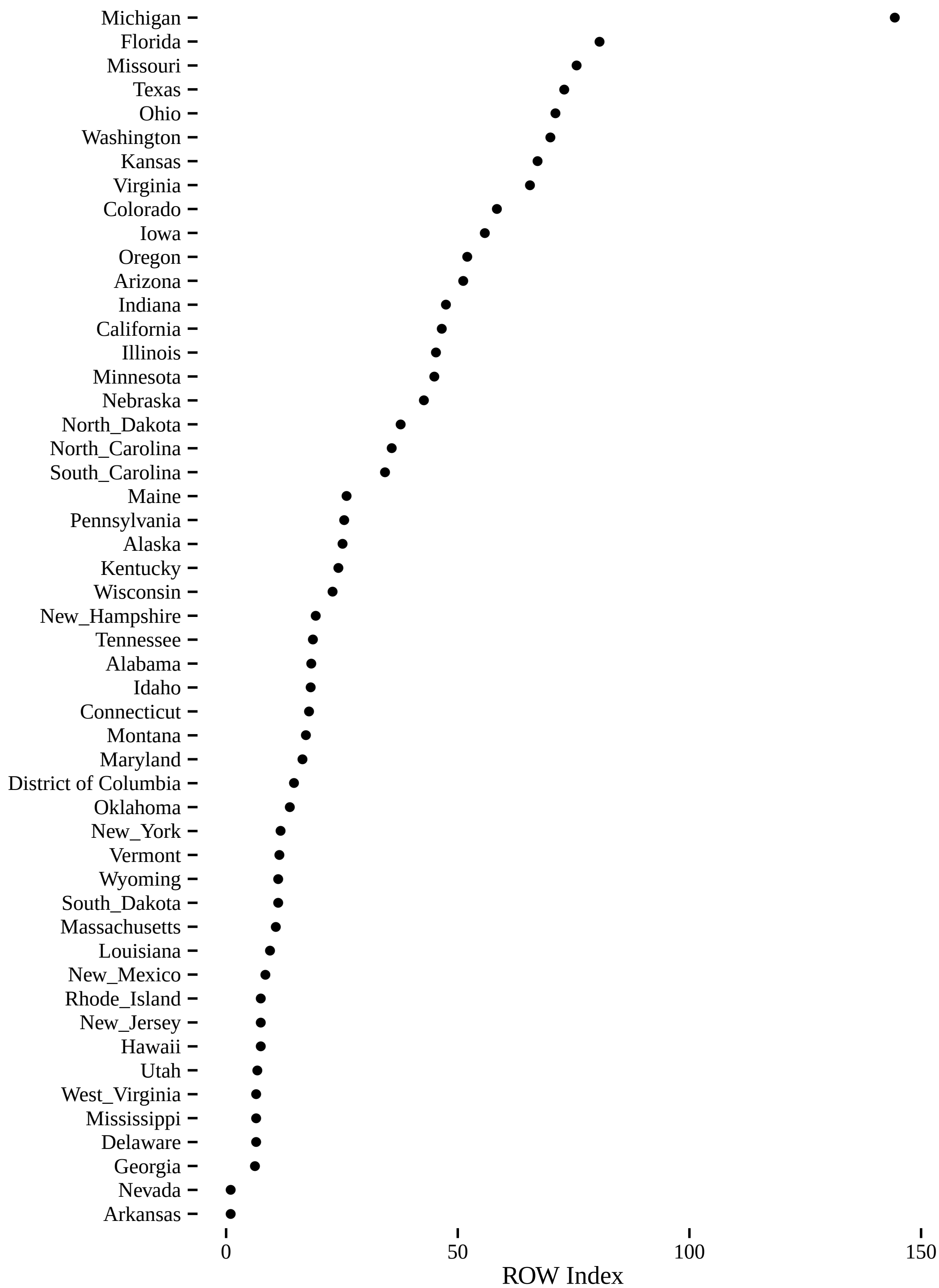
23. Does the state have other broadband applications for public-sector users? (e.g. Department of Corrections using streaming video for remote hearings and sentencing?) (Yes=1; No=0)

24. Does the state have distance learning for the general public using broadband? (Yes=1; No=0)

25. Does the state offer other broadband applications for the general public (e.g. traffic cams using streaming video)? (Yes=1; No=0)

26. Does the state provide streaming video of state legislature proceeding? (Video=1; Audio=0.5)

Figure A1: ROW Value by State



Appendix B: Measures of Affect Towards Candidates

In 2004, respondents were asked the following:

On a scale of zero to 10, how would you rate [George W. Bush/John Kerry]? Zero means very unfavorable, and 10 means very favorable. Five means you do not feel favorable or unfavorable. Of course you can use any number between zero and 10.

Items 2 through 6:

On a scale of zero to 10, how well does [cares about people like me/trustworthy/shares my values/reckless/] apply to [George W. Bush/John Kerry]? Zero means it does not apply at all, and 10 means it applies extremely well.

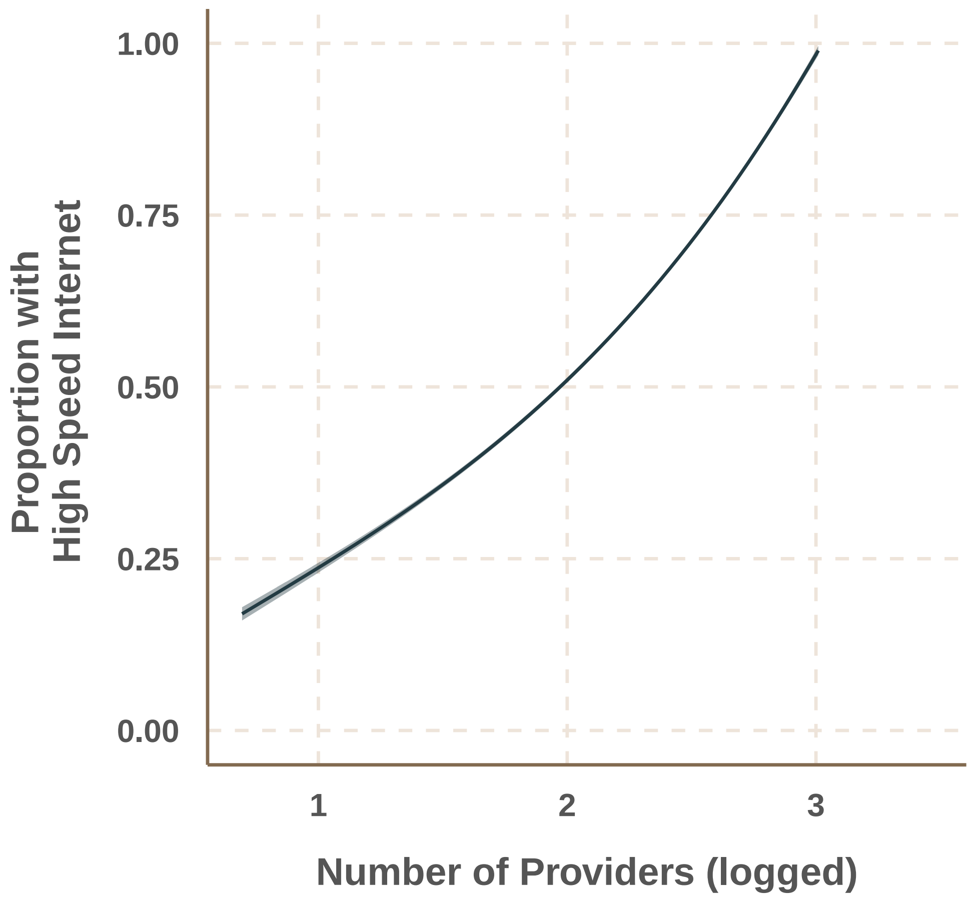
In 2008, respondents were asked the following: Item 1: For the following person, please tell me if your opinion is favorable or unfavorable using a scale from zero to 10. Zero means very unfavorable, and 10 means very favorable. Five means you do not feel favorable or unfavorable toward that person. Of course you can use any number between zero and 10.

On a scale of zero to 10, how would you rate [John McCain/Barack Obama]?

Items 2 through 5: Please tell me how well the phrase [strong leader/trustworthy/has the judgment needed to be president/has the experience needed to be president] applies to [John McCain/Barack Obama]. Please use a scale from zero to 10, where zero means it does not apply at all, and 10 means it applies extremely well. Of course you can use any number in between.

Appendix C: Relationship Between Number of Internet Providers and Broadband Penetration

Figure C1: Percent of ComScore Respondents with Broadband Access in a Zipcode by number of providers.



Smoothed line and 95 percent confidence interval generated using a cubic spline estimator.

What type of people are affected by ROW laws?

Because NAES respondents were not asked detailed questions about their Internet access, we cannot cleanly identify the impact of ROW laws on individual-level broadband access, and therefore we do not know which type of people were impacted by these regulations. We can, however, determine whether certain county-level attributes moderated the impact of ROW on the number of broadband providers in a county.

To estimate whether ROW had heterogeneous effects on the number of providers in a county, we regressed the logged nummber of providers in a county on the county-level controls, the logged ROW index, and the interaction between each control and ROW (see Table C1). These interaction effects indicate that ROW had a larger impact in the Midwest and in more densely populated counties.

Table C1: Does the effect of ROW depend on the type of county?

	Model 1
Intercept	-68.133^{*}
	(20.367)
Year: 2008	4.245^{*}
	(.113)
ROW (logged)	6.426
	(5.524)
Midwest	-4.119^{*}
	(1.463)
South	$-.888$
	(1.704)
West	-2.152
	(1.377)
Percent Black	$-.598$
	(2.979)
Percent White	-5.031
	(2.920)
Percent Male	-6.591
	(6.174)
Low Education County	.344
	(.249)
Unemployment Rate	.050
	(.091)
Population Density	-79.874^{*}
	(28.849)
Median Income (logged)	7.640^{*}
	(1.777)
Midwest x ROW (logged)	1.083^{*}
	(.497)
South x ROW (logged)	.512
	(.568)
West x ROW (logged)	.769
	(.482)
Percent Black x ROW (logged)	.471
	(1.101)
Percent White x ROW (logged)	.822
	(.972)
Percent Male x ROW (logged)	$-.916$
	(1.692)
Low Education County x ROW (logged)	$-.047$
	(.082)
Unemployment Rate x ROW (logged)	$-.006$
	(.030)
Population Density x ROW (logged)	38.647^{*}
	(11.040)
Median Income (logged) x ROW (logged)	$-.681$
	(.478)
R^2	.756
Adj. R^2	.755
Num. obs.	6034

$^{*}p < 0.05$

Appendix D: Robustness Checks

Table D1: Different Reduced-Form Specifications

	Model 1	Model 2	Model 3
ROW Index (logged)	.004*	.004*	.003*
	(.002)	(.002)	(.001)
Year: 2008		−.064*	−.064*
		(.002)	(.002)
Democrat		−.054*	−.054*
		(.007)	(.007)
Race: Black		−.030*	−.026*
		(.006)	(.006)
Race: White		−.049*	−.047*
		(.006)	(.005)
Female		.013*	.013*
		(.001)	(.001)
Age: 2nd Quartile		.008*	.008*
		(.002)	(.002)
Age: 3rd Quartile		.025*	.026*
		(.002)	(.002)
Age: 4th Quartile		.043*	.042*
		(.002)	(.002)
Age: Missing		.014*	.014*
		(.006)	(.006)
Education: Some College		.018*	.016*
		(.001)	(.001)
Education: College		.026*	.023*
		(.002)	(.002)
Education: More than College		.030*	.028*
		(.002)	(.002)
Education: Missing		−.012*	−.011*
		(.004)	(.004)
Income: 2nd Quartile		.011*	.011*
		(.002)	(.002)
Income: 3rd Quartile		.015*	.014*
		(.002)	(.002)
Income: 4th Quartile		.017*	.015*
		(.002)	(.002)
Income: Missing		.019*	.019*
		(.003)	(.002)
Midwest			.010*
			(.004)
South			.013*
			(.004)
West			.027*
			(.004)
Percent Black			.084*
			(.021)
Percent White			.042*
			(.017)
Percent Male			−.089
			(.076)
Low Education County			−.011*
			(.002)
Unemployment Rate			−.001
			(.001)
Population Density			.057*
			(.008)
Median Income (logged)			.006
			(.005)
Intercept	.652*	.683*	.615*
	(.008)	(.012)	(.072)
R^2	.000	.067	.069
Adj. R^2	.000	.067	.069
Num. obs.	114803	114803	114803
RMSE	.186	.180	.180

$^*p < 0.05$

Table D2: Different IV Specifications

	Model 1	Model 2	Model 3
# of Providers	.077	.070	.032*
	(.055)	(.047)	(.014)
Year: 2008		−.105*	−.082*
		(.028)	(.009)
Democrat		−.058*	−.055*
		(.007)	(.007)
Race: Black		−.019	−.026*
		(.011)	(.006)
Race: White		−.053*	−.050*
		(.006)	(.006)
Female		.013*	.013*
		(.001)	(.001)
Age: 2nd Quartile		.008*	.008*
		(.002)	(.002)
Age: 3rd Quartile		.027*	.026*
		(.002)	(.002)
Age: 4th Quartile		.042*	.042*
		(.003)	(.002)
Age: Missing		.015*	.015*
		(.006)	(.006)
Education: Some College		.013*	.016*
		(.003)	(.001)
Education: College		.016*	.022*
		(.006)	(.002)
Education: More than College		.020*	.027*
		(.007)	(.003)
Education: Missing		−.009	−.011*
		(.005)	(.004)
Income: 2nd Quartile		.008*	.010*
		(.003)	(.002)
Income: 3rd Quartile		.008	.014*
		(.005)	(.002)
Income: 4th Quartile		.004	.015*
		(.009)	(.002)
Income: Missing		.012*	.018*
		(.006)	(.002)
Percent Bush			.013*
			(.004)
Midwest			.013*
			(.004)
South			.023*
			(.006)
West			.071*
			(.024)
Percent Black			.071*
			(.024)
Percent White			.076
			(.099)
Percent College			−.012*
			(.004)
Percent Male			−.000
			(.001)
Population Density			.028
			(.016)
Median Income (logged)			−.022
			(.014)
Intercept	.511*	.576*	.760*
	(.111)	(.085)	(.112)
R^2	-.078	.050	.067
Adj. R^2	-.078	.050	.067
Num. obs.	114803	114803	114803
RMSE	.193	.181	.180

$^*p < 0.05$

Table D3: First Stage, Reduced Form, and IV Estimates Using Terrain Instrument

	First Stage	Reduced Form	IV Estimates
# of Providers (logged)			.020*
			(.005)
High mountains	−.852*	.005	
	(.281)	(.005)	
Low mountains	−1.416*	−.012*	
	(.205)	(.005)	
High hills	−1.814*	−.019*	
	(.196)	(.006)	
Hills	−.875*	−.023*	
	(.248)	(.009)	
Open high mountains	−1.711*	.006	
	(.323)	(.008)	
Open low mountains	−.714*	−.010*	
	(.203)	(.004)	
Open high hills	−.769*	−.010*	
	(.297)	(.005)	
Open hills	−1.099*	−.010*	
	(.176)	(.005)	
Open low hills	−1.253*	−.007	
	(.213)	(.006)	
Plains with high mountains	−1.125*	−.001	
	(.271)	(.008)	
Plains with low mountains	−1.705*	.005	
	(.333)	(.017)	
Plains with high hills	−.354	−.011*	
	(.311)	(.005)	
Plains with hills	−.622*	−.007	
	(.264)	(.004)	
Tablelands, very high relief	−1.839*	−.160*	
	(.148)	(.002)	
Tablelands, high relief	−2.208*	−.016	
	(.291)	(.016)	
Tablelands, considerable relief	−1.442*	−.052*	
	(.378)	(.007)	
Tablelands, moderate relief	.262	−.002	
	(.317)	(.003)	
Irregular plains	−.570*	−.004	
	(.162)	(.002)	
Irregular plains, slight relief	−.977	.050	
	(.751)	(.027)	
Smooth plains	−.521*	−.009*	
	(.217)	(.005)	
Year: 2008	4.239*	−.065*	−.077*
	(.027)	(.001)	(.003)
Intercept	4.320*	.699*	.660*
	(.146)	(.002)	(.008)
R^2	.506	.031	.031
Adj. R^2	.504	.030	.031
Num. obs.	5452	95279	95279
RMSE	2.153	.184	.184

$^*p < 0.05$

Table D4: First Stage, Reduced Form, and IV Estimates Using Slope Instrument

	First Stage	Reduced Form	IV Estimates
# of Providers (logged)			.015*
			(.006)
Average Slope	−1.019*	−.015*	
	(.107)	(.007)	
Year: 2008	.622*	−.065*	−.074*
	(.013)	(.001)	(.004)
Intercept	1.793*	.694*	.668*
	(.027)	(.001)	(.011)
R^2	.387	.030	.031
Adj. R^2	.387	.030	.031
Num. obs.	95279	95279	95279
RMSE	.415	.184	.184

$^*p < 0.05$

Appendix E: Causal Mechanisms

Table E1: Regression estimates of effect of Internet speed on visitation of partisan websites.

	Year	
	1	2
Broadband	8.65***	8.18***
	(1.00)	(1.01)
Household Size 2		−5.43***
		(1.61)
Household Size 3		−7.70***
		(1.78)
Household Size 4		−7.47***
		(1.94)
Household Size 5		−7.24***
		(2.28)
Household Size 6		−5.64**
		(2.58)
Oldest Age 2		0.38
		(4.96)
Oldest Age 3		2.58
		(4.80)
Oldest Age 4		3.31
		(4.67)
Oldest Age 5		3.25
		(4.66)
Oldest Age 6		7.09
		(4.61)
Oldest Age 7		4.53
		(4.61)
Oldest Age 8		4.45
		(4.63)
Oldest Age 9		8.30*
		(4.73)
Oldest Age 10		8.62*
		(4.84)
Oldest Age 11		9.83**
		(4.71)
Income 2		−0.91
		(2.30)
Income 3		0.99
		(2.11)
Income 4		1.73
		(2.04)
Income 5		3.52*
		(2.01)
Income 6		5.96***
		(2.25)
Income 7		6.95***
		(2.29)
Children		−0.27
		(1.28)
Children		−3.36**
		(1.65)
Racial Background		−0.97
		(2.54)
Racial Background		−1.81
		(1.93)
Country of origin		−2.21
		(1.47)
Constant	7.89***	6.82
	(0.56)	(4.79)
Observations	39,393	39,393
R^2	0.002	0.004
Note:	*p<0.1; **p<0.05; ***p<0.01	

Table E2: Balance statistics for ANES data

Covariate	All		Matched	
	Means Treated	Means Control	Means Treated	Means Control
Distance	0.16	0.07	0.16	0.16
Political Interest	1.68	1.68	1.64	1.64
Income 1st Quartile	0.62	0.25	0.63	0.63
Income 2nd Quartile	0.25	0.25	0.25	0.25
Income 3rd Quartile	0.09	0.24	0.09	0.09
Income 4th Quartile	0.04	0.26	0.03	0.03
Black	0.16	0.15	0.16	0.16
Hispanic	0.19	0.15	0.18	0.18
Race, Other	0.05	0.06	0.04	0.04
Education, HS or Less	0.47	0.30	0.48	0.48
Education: More than College	0.06	0.14	0.04	0.04
Education: Some college	0.34	0.34	0.35	0.35
Age, 2nd Quartile	0.22	0.26	0.22	0.22
Age 3rd Quartile	0.32	0.24	0.33	0.33
Age 4th Quartile	0.39	0.23	0.38	0.38
Female	0.53	0.50	0.53	0.53

Pew Data

We used the November 2004 Daily Tracking Survey from the Pew Internet & American Life Project.

Respondents that had Internet access at home were asked whether they "connect to the Internet through a dial-up telephone line, or do you have some other type of connection, such as a DSL-enabled phone line, a cable TV modem, a wireless connection, or a T-1 or fiber optic connection?" We dichotomized the measure into dial-up versus other.

The survey also asked respondents how often they read "opinion sites such as [the liberal] Slate.com or [the conservative] National Review" in one question, and "blogs such as [the liberal] Talking Points Memo, [the liberal] Daily Kos, or [the conservative] InstaPundit."

As with the NAES data, we matched respondents Internet access on income, age, race, political interest, and race. The balance statistics appear in Table E4.

We then regressed a measure that indicated whether R visited either opinion sites or blogs regularly (1 if yes; 0 otherwise) on the connection type variable (See Table E5). The

Table E3: Log-odds of selecting partisan media condition on internet speed by Party ID

	Dependent variable:			
	Selected Conservative Media		Selected Liberal Media	
	Republicans	Democrats	Republicans	Democrats
Dial-Up	-1.028^{***}	-1.316^{**}	-2.654^{***}	-1.663^{***}
	(0.373)	(0.598)	(1.014)	(0.398)
Political Interest	-1.098^{***}	-0.564^{***}	-0.302^{**}	-0.960^{***}
	(0.125)	(0.156)	(0.119)	(0.099)
Income: 2nd Quartile	-0.022	0.167	0.274	0.278^{*}
	(0.210)	(0.225)	(0.248)	(0.156)
Income: 3rd Quartile	0.007	-0.498^{*}	0.132	0.100
	(0.203)	(0.292)	(0.243)	(0.168)
Income: 4th Quartile	0.044	-0.394	0.440^{*}	0.246
	(0.206)	(0.299)	(0.239)	(0.169)
Race: Black	-0.295	0.067	-0.866	-0.320^{**}
	(0.519)	(0.227)	(0.742)	(0.138)
Race: Hispanic	-0.023	-0.041	-0.643^{**}	-0.323^{**}
	(0.228)	(0.262)	(0.304)	(0.164)
Race: Other	-0.059	0.497	0.216	0.054
	(0.274)	(0.352)	(0.283)	(0.239)
Education: Some College	0.643^{***}	0.392^{*}	0.608^{***}	0.528^{***}
	(0.190)	(0.224)	(0.218)	(0.154)
Education: College	0.916^{***}	-0.110	0.666^{***}	0.980^{***}
	(0.199)	(0.310)	(0.227)	(0.173)
Education: More than College	0.973^{***}	0.075	0.845^{***}	0.917^{***}
	(0.228)	(0.332)	(0.259)	(0.185)
Age: 2nd Quartile	-0.177	-0.450^{*}	-0.367^{*}	-0.131
	(0.197)	(0.242)	(0.212)	(0.152)
Age: 3rd Quartile	-0.134	-0.727^{***}	-0.155	-0.130
	(0.196)	(0.270)	(0.208)	(0.157)
Age: 4th Quartile	0.037	-0.527^{**}	-0.298	-0.317^{*}
	(0.193)	(0.264)	(0.216)	(0.166)
Female	-0.418^{***}	0.018	0.093	0.159
	(0.132)	(0.185)	(0.146)	(0.111)
Intercept	-0.120	-1.595^{***}	-1.863^{***}	-0.376
	(0.311)	(0.402)	(0.353)	(0.257)
Observations	1,670	2,353	1,670	2,353
Log Likelihood	-756.433	-479.741	-638.064	$-1,048.538$
Akaike Inf. Crit.	1,544.866	991.482	1,308.129	2,129.076
Note:				*p<0.1; **p<0.05; ***p<0.01

probability that a respondent with high speed internet visited a partisan site was p=.10. The probability that a respondent with a dial-up connection visited a partisan site was $p = .03$ ($b = 1.43$, s.e.= .53, $p < .05$).

Table E4: Balance statistics for Pew data

Covariate	All		Matched	
	Means Treated	Means Control	Means Treated	Means Control
Distance	0.58	0.48	0.60	0.60
Age, 1st T	0.49	0.40	0.44	0.44
Age, 2nd T	0.37	0.35	0.45	0.45
Age, 3rd T	0.13	0.24	0.11	0.11
Income 2nd T	0.32	0.32	0.22	0.22
Income 3rd T	0.45	0.26	0.56	0.56
Education: College	0.56	0.35	0.64	0.64
Education: Some college	0.26	0.34	0.22	0.22
Interest	1.45	1.56	1.27	1.27
Race: White	0.85	0.84	0.94	0.94
PID: Independent	0.36	0.31	0.27	0.27
PID: Republican	0.34	0.36	0.42	0.42
Female	0.45	0.49	0.44	0.44

Table E5: Probability of selecting partisan media by internet type.

	Dependent variable:	
	Visits Partisan Sites	
	(1)	(2)
Highspeed Internet	1.360^{**}	1.483^{**}
	(0.654)	(0.677)
Age: 2nd T		-1.212^{*}
		(0.657)
Age: 3rd T		-1.695
		(1.229)
Income, 2nd T		-2.596
		(1.628)
Income, 3rd T		-0.882
		(1.244)
Education: College		1.275
		(1.424)
Education: Some College		-0.437
		(1.235)
Political Interest		-1.563
		(1.084)
Race: White		1.456
		(1.261)
PID: Independent		-1.120
		(0.789)
PID: Republican		-1.737^{**}
		(0.807)
Female		-1.765^{**}
		(0.737)
Constant	-3.528^{***}	-1.025
	(0.595)	(2.002)
Observations	248	248
Log Likelihood	-61.813	-50.893
Akaike Inf. Crit.	127.626	127.786
Note:	*p<0.1; **p<0.05; ***p<0.01	