Importing Political Polarization? The Electoral Consequences of Rising Trade Exposure[*]

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Abstract

Has rising import competition contributed to the polarization of U.S. politics? Analyzing multiple measures of political expression and results of congressional and presidential elections spanning the period 2000 through 2016, we detect an ideological realignment in trade-exposed local labor markets that commences prior to the divisive 2016 U.S. presidential election. Exploiting the exogenous component of rising trade with China, we find that greater exposure to import competition led to an increasing market share for the FOX News channel, stronger ideological polarization in campaign contributions, and a disproportionate rise in the likelihood of electing a Republican to Congress. Trade-exposed counties with an initial majority white population became more likely to elect a GOP conservative, while trade-exposed counties with an initial majority-minority population become more likely to elect a liberal Democrat, and in both sets of counties, these gains came at the expense of moderate Democrats. In presidential elections, counties with greater trade exposure shifted towards the Republican candidate. We interpret these results as supporting a political economy literature that connects adverse economic shocks to sharp ideological realignments that cleave along racial and ethnic lines and induce discrete shifts in political preferences and economic policy.

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1 Introduction

The ideological divide in American politics is at a historic high. Ranking congressional legislators on a liberal-conservative scale, based either on their roll-call votes (McCarty et al., 2016) or the political orientation of their campaign contributors (Bonica, 2013), reveals that the gap between Democrats and Republicans has been widening since the late 1970s.[1] A sizable rightward shift among the GOP and a modest leftward shift among Democrats has left few centrists in either party. The 21^{st} century has also seen greater polarization in the policy preferences and media viewing habits of the American public. In the 1990s and early 2000s, roughly half of respondents took moderate positions on prominent political issues; by the late 2000s the centrist share had shrunk to under 40%, as individuals adopted more strident views on the left or right (Pew Research Center, 2014; Gentzkow, 2016). These trends are also reflected in, and amplified by, popular media (DellaVigna and Kaplan, 2007; Levendusky, 2013; Martin and Yurukoglu, 2017). After decades of ostensibly nonpartisan news programming on major TV networks, the FOX news channel launched in 1996 with a candidly pro-conservative slant. It has been the top cable news channel since 2001.

The causal linkages between economic shocks and sustained increases in partisanship remain poorly understood. Mian et al. (2014) find that while congressional voting patterns become more polarized following financial crises, these movements are temporary.[2] Although the widening ideological divide in Congress tracks rising U.S. income inequality (McCarty et al., 2016; Voorheis et al., 2016), the coincidence of these two phenomena does not reveal which underlying shocks intensify partisanship. The rise in political polarization appears causally unrelated to the structure of primary elections, rule changes in Congress, gerrymandering, or immigration.[3] Like the rise in income inequality (Katz and Autor, 1999), polarization may be the result of multiple causal factors. Yet in contrast to research on inequality, we do not know which factors should populate this list.

In this paper, we examine whether the exposure of local labor markets to increased foreign competition has contributed to rising political polarization in the U.S. since 2000. We begin by documenting the effect of exposure to rising manufacturing imports from China on political expression, as captured by the stated political values of voting-age adults, the TV-news viewing habits of households, and the ideological orientation of campaign contributors to congressional elections. Next, we estimate the impact of greater import competition on changes in the party affiliation and ideology

[1]In the 1990s and especially the 2000s, greater polarization is also evident in the content of political speech used by Democratic and Republican legislators (Gentzkow et al., 2019).

[2]Their analysis spans 1879 to 2010. Over the 1948-2008 subperiod, financial crises are followed by increases in voter identification with more extreme ideological positions. For related analysis of Europe, see Funke et al. (2016).

[3]See Gelman (2009), McCarty et al. (2009), Barber et al. (2015), and McCarty et al. (2016) Other factors potentially related to polarization include intensified media partisanship (DellaVigna and Kaplan, 2007; Levendusky, 2013; Prior, 2013) and stronger ideological sorting of voters by party (Levendusky, 2009). See Canen et al. (2018) on the contribution of the practices and structures of the major political parties to polarization.

of congressional legislators. We then consider how greater trade exposure has affected presidential voting. These exercises allow us to evaluate whether adverse shocks related to trade cause voters and campaign donors to take more extreme positions and to elect more extreme politicians.

The appeal of studying the China trade shock is the abundant evidence linking foreign competition to the decline in U.S. manufacturing jobs. Whereas in earlier decades manufacturing helped workers without a college degree reach the middle class, the sector's steep decline has left U.S. employment more partitioned between highly paid professional occupations and low-wage service jobs (Autor and Dorn, 2013a; Barany and Siegel, 2018). Industries more exposed to trade with China have seen higher exit of plants (Bernard et al., 2006), larger contractions in employment (Pierce and Schott, 2016; Acemoglu et al., 2016a), and lower incomes for affected workers (Autor et al., 2014; Galle et al., 2017; Caliendo et al., 2019; Autor et al., 2019). The local labor markets that are home to more-exposed industries have endured sustained job loss and increases in unemployment, non-participation in the labor force, uptake of government transfers (Autor et al., 2013, 2019), and declines in tax revenues and housing prices (Feler and Senses, 2015). The steepest increase in U.S. imports occurred just after China's accession to the World Trade Organization in 2001; China's share of world manufacturing exports surged from 4.8% in 2000 to 15.1% in 2010. This export boom was driven by reform-induced productivity growth in Chinese manufacturing (Naughton, 2007; Brandt et al., 2017a), where China's reform push and the productivity gains associated with it appeared to have largely abated by the late 2000s (Brandt et al., 2017b).[4] The concentrated impact of the China shock on specific industries and regions makes the economic consequences of trade acutely recognizable and therefore politically salient (Margalit, 2011; Di Tella and Rodrik, 2019).

While U.S. political polarization did not originate with the China shock, U.S. politics have further polarized amid this expansion of trade. Moderate Democrats have become increasingly rare in the House of Representatives, while Tea Party and like-minded conservatives have risen to prominence in the GOP (Madestam et al., 2013). The surprise election of Donald J. Trump to the Presidency in 2016 has further heightened partisan divisions and injected ethnic nationalism into Republican policy positions.[5] Among voters, ideological divides by race and education have also widened, as seen most notably in a shift of less-educated whites toward the GOP (Pew Research Center, 2017). We test whether the China shock contributed to these ideological and electoral shifts.[6]

[4]Over 2000 to 2010, China's annual growth rate in manufacturing exports was 11.5 percentage points faster than the world as a whole; over 2010 to 2016, its growth advantage was just 2.8 percentage points. Although China's manufacturing exports continued to grow after 2010, the transitional growth associated with China's aggressive post-Mao market-oriented reforms—on which we base our empirical identification strategy—appears to have been largely exhausted by the end of that decade (Naughton, 2018; Lardy, 2019).

[5]Such nationalism is not new to the Republican Party. Trump's policy positions strongly echo the ethnocentrism of Pat Buchanan, a GOP presidential candidate in 1992 (Buchanan, 2010), and the isolationism and protectionism of Senator Robert Taft, a GOP presidential candidate in 1940, 1948, and 1952 (Kabaservice, 2012).

[6]For analysis of the rise of right-wing populism in high-income countries, see Inglehart and Norris (2016), Algan

To perform this analysis, we must address two empirical challenges. One is that local labor markets, which we take to be Commuting Zones (CZs), do not map one-to-one into congressional districts. Whereas CZs aggregate contiguous counties, gerrymandering creates congressional districts that often span parts of several commuting zones. We resolve this issue by dividing the continental U.S. into county-by-congressional-district cells, attaching each cell to its corresponding CZ, and weighting each cell by its share of the district voting-age population. To measure regional trade exposure, we use the change in industry import penetration from China, weighting each industry by its initial share of CZ employment. We isolate the component of U.S. import growth that is driven by export-supply growth in China, rather than U.S. product-demand shocks, following the identification strategy in Autor et al. (2014) and Acemoglu et al. (2016b).

A second empirical challenge, potentially more formidable than the first, is the discontinuous changes in voter composition and frequent turnover of incumbent representatives caused by the redrawing of U.S. congressional districts after each decennial census. We surmount this issue by studying changes in the ideology of legislators first over the 2002 to 2010 period, during which most district boundaries are fixed and the mapping of districts to CZs is stable, and then over the 2002 to 2016 period, which requires us to account for redistricting-induced changes in the mapping of districts to CZs. We exclude electoral outcomes across the immediate 2010-2012 redistricting seam, except for the small number of districts that did not change boundaries.

To begin our study, we evaluate how trade shocks affect political expression. We use survey data from the Pew Research Center to examine the political beliefs of voting-age adults, ratings data from Nielsen Media to examine the news viewing habits of households, and the Database on Ideology, Money in Politics, and Elections (DIME: Bonica, 2013, 2014) to examine the ideological leanings of campaign donors. Over 2004 to 2016, regions more exposed to import competition from China became more likely to see support for positions on the right of the political spectrum as reported in Pew survey data, to shift consumption of TV news to the right-leaning FOX News Channel, and to attract campaign contributions from left-wing and right-wing donors, but not from moderates. Each of these outcomes manifests differently among non-Hispanic whites than among other racial and ethnic groups. This analysis indicates that exposure to import shocks moves political sentiment away from the ideological center along three dimensions: the valence of expressed political beliefs, the ideological skew of media consumption, and the ideological stripe of campaign donations.

Following this analysis of political sentiment, we examine the impact of exposure to trade shocks on the party and ideological composition of elected congressional legislators.[7] Although shocks

et al. (2017), Gidron and Hall (2017), Guiso et al. (2017), Dehdari (2018), and Dal Bo et al. (2019). For related work on populism, see Karakas and Mitra (2017), Rodrik (2017), and Pastor and Veronesi (2018).

[7] A large literature, beginning with Fair (1978), finds that economic downturns hurt sitting politicians. Margalit

increase campaign contributions from both liberal and conservative donors, across all districts the net beneficiaries in terms of electoral results are Republican candidates, and conservative candidates in particular. Districts exposed to larger increases in import competition became more likely to elect a GOP legislator in each election from 2010 to 2016, where conservative rather than moderate Republicans absorbed these electoral gains.[8]

To understand why, despite inducing ideological polarization, trade shocks primarily improve electoral outcomes for Republicans, we break voting districts down by their racial and ethnic composition. This step reveals that rising trade exposure simultaneously raises the odds that majority white non-Hispanic areas elect GOP conservatives and that majority non-white areas elect liberal Democrats. Because in both sets of districts candidates advantaged by adverse trade shocks pull support from moderate Democrats, it is the GOP that gains on net from these shocks.

Why does greater trade exposure induce ideological polarization? One potential explanation is resource competition. Adverse trade shocks increase local uptake of government transfers (Autor et al., 2013) and reduce local tax revenues (Feler and Senses, 2015), meaning that they are likely to intensify competition for government funds. In the U.S., support for the provision of public services tends to divide along racial and ethnic lines (Alesina et al., 1999; Alesina and La Ferrara, 2005). Since the post-civil rights movement realignment of the major political parties, stronger group identification among whites has typically meant a stronger association with the GOP (Kuziemko and Washington, 2018). These divisions are embodied in the Tea Party movement, which objects to multiculturalism and to government benefits being captured by constituencies defined by race, ethnicity, or nativity (Parker and Barreto, 2014). To the extent that white voters disadvantaged by economic changes see GOP conservatives as favoring their interests over those of other groups—while disadvantaged minority voters see liberal Democrats as their champions—we would expect the political response to a common shock to vary by race, an implication that our analysis supports.

Perhaps a deeper explanation for why politicians would endorse both group identity and protectionism in the face of rising trade exposure is found in the behavioral general-equilibrium framework of Grossman and Helpman (2018).[9] Introducing social-identity theory into an otherwise standard trade model, they show that adverse economic shocks—due, e.g., to globalization—may engen-

(2011) and Jensen et al. (2017) show that voters also punish incumbents for the adverse effects of import competition.

[8] We find that the GOP was especially successful in increasing its vote share in competitive districts, while the Democratic party gained vote shares but not congressional seats in non-competitive districts. The net impact of the China shock on the Republican vote share in congressional elections is small, and in some years even negative, as is the case in a county-level vote share analysis by Che et al. (2016). Whereas it can be challenging to interpret the electoral consequences of county-level congressional vote shares that are aggregated across multiple gerrymandered districts, one can readily study vote shares of presidential elections, given that all counties choose among the same candidates. We find that over the 2000 to 2016 period, counties with a larger increase in trade exposure saw a larger increase in the vote share of the GOP presidential candidate.

[9] Gennaioli and Tabellini (2019) also model how economic shocks affect group identity and political preferences.

der both a psychological response that strengthens one's identification with a particular social group—e.g., the white working class—and a material interest in stronger trade protection. Intensified foreign competition (or other shocks) may thus increase the salience of racial and ethnic identities among voters, along with support for nationalist economic policies.

Although we cannot definitively separate resource-based versus identity-based explanations for why trade shocks catalyze the political outcomes that we observe, the Grossman-Helpman framework does make a prediction that is not directly implied by the resource-based view, which is that trade shocks increase support for protectionism. Consistent with this implication, Feigenbaum and Hall (2015) demonstrate that support for protectionist trade bills was stronger among legislators whose congressional districts were more exposed to the China shock during the 1990s and 2000s.[10]

Given that the GOP has endorsed the principle of free trade since the 1950s (Irwin, 2017), it may appear paradoxical that trade shocks simultaneously advantage Republican candidates, as we show below, and increase support for protectionism, as demonstrated by Feigenbaum and Hall (2015).[11] One resolution to this puzzle, found in Feigenbaum and Hall (2015), is that support for protectionism is greater in trade-exposed districts irrespective of whether these districts are classified as safe Democratic, safe Republican, or competitive.[12] The election of Donald Trump may further suggest that what it means to be a conservative is changing (Mann and Ornstein, 2016). Continuing a shift that began with the Tea Party movement—which tends to see trade agreements as compromising American sovereignty—Donald Trump has pushed the GOP in a more nationalist and white-identitarian direction, mirroring the evolution of conservatism and the rise of right-wing populism in Europe (Inglehart and Norris, 2016; Colantone and Stanig, 2018b).[13]

Because right-wing populist movements tend to arise during times of economic hardship (Hutchings and Valentino, 2004; Inglehart and Norris, 2016; Algan et al., 2017), a third strand of explanation for the pattern of results we report is that animus towards minorities and foreigners stems from political opportunism. Glaeser et al. (2005) formalize this intuition in a model in which politicians

[10] In related work, Kleinberg and Fordham (2013) and Kuk et al. (2015) find that legislators from districts harder hit by the China trade shock are more likely to support foreign-policy legislation that rebukes China. For work on how congressional representatives vote on trade legislation, see Bailey and Brady (1998), Baldwin and Magee (2000), Hiscox (2002), and Milner and Tingley (2011). On labor-market shocks and support for protectionism, see Colantone and Stanig (2018b) on Europe and Di Tella and Rodrik (2019) on the U.S.

[11] By contrast, in the late 19^{th} and early 20^{th} centuries, the GOP tended to be strongly protectionist, given its geographic bases of support in manufacturing-oriented states in the Midwest and Northeast (Irwin, 2017).

[12] Feigenbaum and Hall (2015) show that import competition raises support for protectionist trade bills by an equal extent in safe Democratic and safe Republican districts, and about twice as much in competitive districts.

[13] Members of the Republican Liberty Caucus and the House Freedom Caucus, two groups of right-wing GOP legislators in the House, opposed the Trans-Pacific Partnership, a recent major trade deal, and are frequent critics of the WTO. Suspicion of trade agreements on the right is not historically novel. The conservative stalwart Senator Barry Goldwater opposed the Trade Expansion Act of 1962, which enabled the president to negotiate tariff reductions in the Kennedy Round of the General Agreement on Trade and Tariffs. In terms of public opinion, GOP voters are wary of trade accords. In a 2016 Pew Research Center survey, 53% of voters who identify or lean Republican, as compared to 34% of voters who identify or lean Democrat, saw free-trade agreements as a "bad thing for the U.S."

engage in strategic extremism (e.g., inflaming wedge issues) to raise voter turnout and campaign contributions among their core supporters. Our results do not adjudicate between opportunism-based and identity-based explanations for rising political polarization; indeed, one can argue that these mechanisms are strategic complements, with adverse shocks triggering group-identity shifts, and politicians exploiting these shifts for electoral gain. One alternative explanation that we *can* reject is that our findings are a byproduct of a secular trend favoring conservatives. That trade shocks led to ideological polarization means that more-extreme actors in *both* parties became more engaged at the expense of moderates—though conservatives have benefited disproportionately.

The U.S. is not alone in seeing economic adversity expand support for right-wing politicians. During the Great Depression, far-right movements had greater success in European countries that had more prolonged downturns (De Bromhead et al., 2013). Today, French and German regions more exposed to trade with low-wage countries have seen larger increases in vote shares for the far right (Dippel et al., 2017; Malgouyres, 2017), British regions more exposed to trade with China voted more strongly in favor of Brexit (Colantone and Stanig, 2018a), and EU regions more exposed to the Great Recession have seen a greater rise in voting for anti-establishment, Euro-skeptic parties (Algan et al., 2017; Dehdari, 2018; Dal Bo et al., 2019). Our work differs from existing literature by demonstrating that trade shocks generate, first, shifts in elected representatives in the highest offices, and, second, a polarized response to these shocks—both in campaign contributions and in the party and ideological repositioning of majority-white versus majority-non-white regions.

In section 2, we describe our data on political beliefs, media viewership, and campaign contributions, and next summarize our data on local labor markets, how we match these markets to congressional districts, and how we account for congressional redistricting in section 3. We present our empirical results on the impacts of trade shocks on political expression in section 4, on legislator ideology in section 5, and on presidential voting in section 6. Section 7 concludes.

2 National Trends in Political Expression and Partisanship

We begin our analysis by considering how political expression in the U.S. has changed over time. To account for myriad forms of political engagement, we study three disparate types of expression. Surveys of public opinion from the Pew Research Center provide direct information on the political beliefs of potential voters; Nielsen data on the ratings of cable news networks capture the relative standing of right-leaning FOX News and more left-leaning MSNBC and CNN; and DIME measures of campaign contributions indicate how donor support for candidates has shifted along the ideological spectrum. In concert, these data reveal the demand side for ideology, which we will later use to examine which viewpoints and sentiments have been most emboldened by adverse trade shocks.

2.1 Changes in Voter Beliefs on Political Issues

To measure changes in voter ideology over time, we use data from the Pew Research Center.[14] Pew periodically asks U.S. adult survey participants a consistent set of questions about their political beliefs (see Appendix Table A1). In each of ten questions, participants choose which of two opposing statements on a topic—one left-leaning, one right-leaning—best reflects their opinion. For example, the first pair of statements is "government is almost always wasteful and inefficient" versus "government often does a better job than people give it credit for." Subsequent questions cover government regulation of business, corporate profits, government assistance to the poor, environmental regulations, the role of the military in national security, and attitudes toward African Americans, immigration, and homosexuality. By coding agreement with left-leaning and right-leaning statements as -1 and +1, respectively, Pew constructs a measure of the left–right distribution of political beliefs ranging from -10 to 10, which we refer to as the Pew ideology score.[15] We use data on political beliefs, rather than party identification, because beliefs directly reflect ideology whereas party attachment may not (Abramowitz and Webster, 2016).[16]

Table 1: Share of Population by Pew Ideology Score, 2004 to 2015

Year	Mean Score (1)	% Liberal (2)	% Moderate (3)	% Conservative (4)
2004	-0.91	32.6	48.7	18.6
2011	-0.30	31.2	42.2	26.6
2014	-0.59	34.5	39.3	26.2
2015	-0.62	35.4	37.6	26.9
Δ2004-15	0.29	2.8	-11.1	8.3

Notes: The Pew Ideology score ranges from -10 (most liberal) to +10 (most conservative). Columns 2-4 define liberals as those with scores of -10 to -3, moderates as those with scores from -2 to 2, and conservatives as those with scores from 3 to 10. Sample sizes for survey participants who reside in the 48 mainland states and who have complete demographic information are 1,994 in 2004, 3,016 in 2011, 9,868 in 2014, and 5,907 in 2015. Observations are weighted by survey weights.

We obtained from Pew unpublished geocoded microdata for its surveys in 2004, 2011, 2014, and 2015, in which participants were asked about the 10 pairs of belief statements, yielding a pooled

[14] Other surveys of political attitudes include the American National Election Studies, the General Social Survey, and the Cooperative Congressional Election Survey. None are suitable for our purposes. The first two have commuting zone-level sample sizes that are very small, while the third does not commence until late in our study period.

[15] For further analyses of the Pew ideological consistency scale, see Pew Research Center (2014). If respondents do not have a preference between the two statements (i.e., they say they disagree with both or do not know), their answer to this statement is coded as a 0 for the construction of the Pew ideology score.

[16] Other work that uses Pew data to study polarization includes Gentzkow et al. (2019). Gentzkow (2016) discusses alternative measures of political polarization used in the literature.

sample of 20,785 observations.[17] The data show both a rightward shift and a strong polarization in participant political beliefs over the 2000s.[18] In Table 1, the mean ideology score increased from -0.91 to -0.62 from 2004 to 2015, corresponding to one more survey item with a right-leaning answer for every seven respondents.[19] The fraction of participants whose ideology was centrist (Pew score of -2 to 2) fell from 48.7% in 2004 to 42.2% in 2011 and declined further to 37.6% in 2015. By contrast, the fraction of participants whose ideology was mostly or strongly conservative (Pew score of 3 to 10) rose from 18.6% in 2004 to 26.9% in 2015, with most of this change occurring by 2011. The fraction whose ideology was mostly or strongly liberal (Pew score of -3 to -10) rose more modestly from 32.6% to 35.4% over the 2004-2015 time frame.

Changes in political beliefs vary markedly by race and ethnicity. The rightward shift in ideology evident in Table 1 is due almost entirely to the preferences of non-Hispanic whites (see Appendix Table A2). Whereas their mean ideology score increased from -0.63 to 0.09 from 2004 to 2015 (corresponding to one more right-leaning answer for every three whites), the mean ideology score of Hispanics and non-whites shifted leftwards from -1.65 to -1.97 (one more left-leaning answer for every six respondents), thus doubling the ideological distance between whites and other groups. The share of whites with conservative beliefs rose sharply from 22.2% to 35.0%, while the prevalence of liberal beliefs increased from 37.3% to 44.0% among Hispanics and non-whites. These patterns reveal increasing polarization of left-right beliefs between non-Hispanic whites and other groups.

2.2 Changes in Cable News Viewing Habits of Households

As a second measure of the ideological orientation of the American public, we exploit the distinct role of the FOX News Channel in national political life. In a break with long-standing convention in network TV programming, FOX News has openly supported Republican politicians and viewpoints and opposed Democratic ones. Utilizing geographic variation in the post-1996 rollout of access to FOX News, DellaVigna and Kaplan (2007) find that, relative to towns without FOX News access, towns exposed to the channel had larger 1996-to-2000 gains in votes shares for GOP presidential and senatorial candidates. Martin and Yurukoglu (2017) document that over the 2000 to 2012 period, the conservative slant of FOX News intensified further. Tapping regional variation in the local channel number of FOX News, they find that in areas in which FOX News has a lower channel number, its viewership was higher, and higher viewership translated into stronger support for GOP presidential

[17] We retain all survey respondents who reside in the 48 mainland states, and drop the 0.6% of observations that have incomplete demographic information. Microdata prior to 2004 were unavailable.

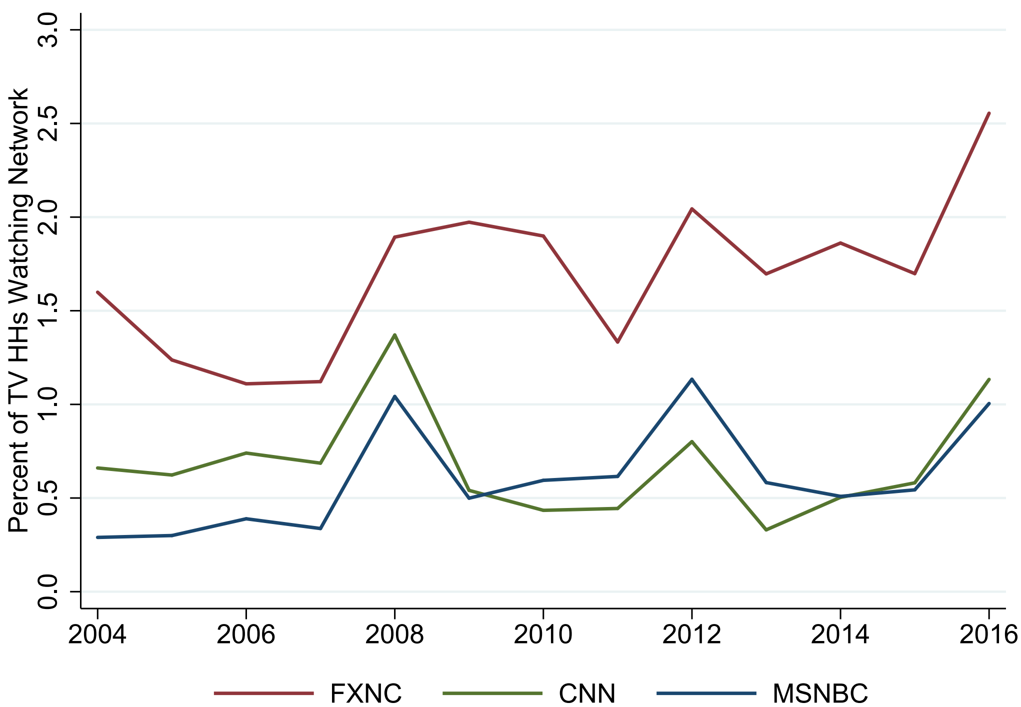
[18] Survey data that measure respondents' views of the other party rather than their views on specific issues find a sharp rise in polarization in the mid-1990s—suggesting an increase in party salience. Data that track polarization of specific issue positions, however, do not find a rise of polarization until the mid-2000s (Gentzkow, 2016).

[19] Changing a survey response from left-leaning to right-leaning raises the ideology score by 2 points (+1 instead of -1); the increase in average score by 0.29 corresponds to $0.29/2 = 0.15$ additional right-leaning answers per person.

candidates. This positive impact of FOX News on the GOP strengthened over 2000 to 2008.

Based on the documented connection between FOX News and conservative politics, we use viewership of the channel as an indication of household demand for partisan media content. We compare ratings for FOX News with ratings for the other two large cable news networks, CNN and MSNBC.[20] CNN, which launched in 1980, and MSNBC, which launched in 1996, were long seen as less partisan than FOX News, though both have greater viewership among Democratic voters than among Republicans. Using text analysis of TV news transcripts over 1998 to 2012 and the speech of congressional representatives with known ideological positions, Martin and Yurukoglu (2017) classify the content of FOX News as strongly right of center, the content of CNN as modestly left of center, and the content of MSNBC as similar to CNN until 2009, and further left thereafter. More recently, CNN has had antagonistic interactions with Donald Trump since his election in 2016.

Figure 1: Nielsen Ratings for Cable TV News Networks, November 2004 to November 2016



Notes: Nielsen ratings indicate the fraction of all TV-owning households that are tuned to a particular program at a particular time. Figure plots average ratings for the 5pm to 11pm time-slot, Monday through Friday, during the month of November for years 2004 through 2016. Sample size for each November estimate ranges from 99,000 to 119,000 households.

Our ratings data are from Nielsen Local TV View, which tracks TV viewing in U.S. households. Nielsen measures the size of the audience for a given programming hour of a given TV show on

[20]We focus on cable news networks rather than network TV news (ABC, CBS, NBC, and PBS) because cable news provides news programming over all or nearly prime-time hours, whereas the other networks typically broadcast nightly news for just 30 minutes, which amounts to a small share of their total content and messaging.

a given network. Based on electronic monitors attached to household TV sets and viewer diaries, the Nielsen ratings indicate the fraction of all TV-owning households that are tuned to a particular program at a particular time. We obtained average ratings for the 5pm to 11pm time-slot, Monday through Friday, which is prime time for cable news programming. Our data cover 2004 to 2016, with sample sizes ranging from $99,000$ to $119,000$ households in each month during which Nielsen conducts ratings "sweeps" (February, May, July, November). To align news viewership with the demand for political content as closely as possible, we focus on ratings for cable news in the month of November during presidential election years. Ratings for cable news spike during presidential election months, averaging 3.9% versus 2.7% in non-election months during our sample period. The data record average ratings for the households of each county, which we use in later analysis to examine how exposure to trade shocks affects the viewership of cable news networks.

Figure 1 shows average national November ratings for the three major cable TV news channels. Aggregating the networks' viewing audience, overall ratings for cable TV news rose from 2.37% in 2004 to 3.86% in 2016, implying that over the sample period, the average fraction of households that were tuned to a cable news channel during prime-time hours in November rose by 1.5 percentage points. In all years, FOX News is the dominant network. Its ratings rose from 1.36% in 2004 to 2.04% in 2016, while its share of cable news viewers declined modestly from 57.5% in 2004 to 52.8% in 2016. Ratings for CNN rose from 0.70% in 2004 to 1.00% in 2016, which resulted in a drop in its cable-news market share from 29.7% to 26.0%. For MSNBC, ratings rose from 0.30% in 2004 to 0.82% in 2016, as its market share expanded from 12.8% to 21.2%.[21] Since 2004, the news viewing of the American public has modestly polarized. Ratings during presidential-election months have increased more substantially for right-leaning FOX News (+0.7 percentage points) and left-leaning MSNBC (+0.5 percentage points) than for less-partisan CNN (+0.3 percentage points).[22]

In Appendix Figure A1, we separate Nielsen households according to the race and ethnicity of the household head. The 2004 to 2016 FOX News gain in ratings is large among households headed by non-Hispanic whites ($1.03\% = 2.66\% - 1.63\%$) and negligible among households headed by non-whites ($0.06\% = 0.52\% - 0.46\%$), relatively few of which are FOX viewers. For MSNBC, the ratings gains among white-headed households ($0.46\% = 0.79\% - 0.33\%$) are slightly smaller than among non-white-headed households ($0.68\% = 0.90\% - 0.21\%$). These patterns reveal further evidence of divergence in political expression between non-Hispanic whites and other groups.

[21]Over all four ratings sweeps months, cable-news market shares rose for FOX NEWS, from 61.3% (1.47/2.40) in 2004 to 66.2% (2.59/3.91) in 2016, and for MSNBC, from 8.8% (0.21/2.40) in 2004 to 14.8% (0.58/3.91) in 2016, while they fell for CNN from 30.0% (0.72/2.40) in 2004 to 19.2% (0.75/3.91) in 2016.

[22]Because over time FOX News has become more conservative and MSNBC has become more liberal, their ratings gains may understate the increase in consumption of politically charged news from partisan outlets.

2.3 Ideology of Campaign Contributors and Congressional Legislators

To measure the political ideology of campaign contributors and legislators, we use the Database on Ideology, Money in Politics, and Elections (DIME; Bonica, 2013), which tabulates campaign contributions by donor and recipient for all amounts in excess of $200, based on reports mandated by the Federal Electoral Commission (FEC). DIME encapsulates the ideology of campaign donors and electoral candidates in a campaign finance (CF) score. The CF score is based on the solution to a spatial model of contributions. Bonica (2013) proposes that donors choose contributions to each candidate to maximize the difference between the net benefit they derive from giving to candidates in general and the loss they experience when giving to particular candidates whose ideological positions differ from their own. Applying the model to the universe of FEC-registered campaign donors and candidates for state and national electoral offices, he estimates the ideal points for each entity in the data (i.e., campaign donors and candidates for elected office), which are the CF scores.

In our data, the largest conservative donors include Associated Builders and Contractors (anti-regulation), the National Rifle Association (pro-gun rights), and the National Right to Life Political Action Committee (anti-abortion); the largest liberal donors include the Association of Trial Lawyers of America (pro-plaintiff rights), the Service Employees International Union (pro-labor), and the American Federation of State, County, and Municipal Employees (pro-public sector). Because donors in the first group give to similar candidates, few if any of whom are supported by donors in the second group, and vice versa, the model solution will give extreme CF scores in one direction to donor-candidate combinations in the first group and extreme CF scores in the other direction to donor-candidate combinations in the second group. Donors who give widely to candidates (e.g., the National Auto Dealer's Association or the National Beer Wholesaler's Association) and candidates who receive contributions from a wide variety of donors will have intermediate CF scores.[23]

The use of spatial models to measure the ideology of political actors was pioneered by Poole and Rosenthal (1985, 2007), who developed an approach for detecting the ideology of members of Congress from roll-call votes. Their Nominate scores of legislator ideology have been widely emulated and extensively applied (see, e.g., Nokken and Poole, 2004; McCarty et al., 2016). A notable advantage of CF scores is that they provide ideology measures not only for election winners, who subsequently cast votes in Congress, but also for election losers (as long as they receive campaign contributions), even though their subsequent legislative records are not observed.

Basic versions of CF and Nominate assume that a politician's ideology is time-invariant even

[23]Even controlling for legislator party affiliation, CF scores for members of Congress are strongly positively correlated with the likelihood that a representative voted in favor of legislation deemed as conservative (e.g., stronger immigration enforcement) and strongly negatively correlated with the likelihood that a representative supported legislation deemed liberal (e.g., the Affordable Care Act). See Bonica (2019).

over a decades-long tenure in Congress, which is unappealing for our analysis that studies changes in ideology over time. To address this limitation, we derive time-variant ideology scores for candidates by computing the contribution-weighted-average of the time-invariant CF donor scores of each candidate's donors.[24] In so doing, we follow the political science literature (e.g., Bonica, 2013; McCarty et al., 2016) in interpreting a donor's choice of candidates to support to be a genuine expression of the donor's ideology. Aggregating over contributions to candidates in a given election reveals the relative demand for ideology by donors in that election, and aggregating over the CF scores of donors to a particular candidate reveals the relative demand for that candidate's ideological position. For the Congress elected in 2002, the correlation between our time-variant legislator CF-score and the time-invariant CF-score of Bonica (2013) is 0.97, while the correlation with the time-variant DW-Nominate score of Poole and Rosenthal (2007) is 0.92.[25] In addition to studying outcomes that capture the ideology of a legislator by a single number, the DIME data allow us to study the extent to which a legislator draws financial support from donors at different quantiles of the donor ideology distribution.

Figure 2 summarizes campaign contributions to all candidates in primary and general congressional elections from 2002 to 2016, where we group donors based on terciles of CF scores in 2002.[26] By keeping tercile cutoffs fixed over time, we evaluate how contributions shift across the distribution of CF scores, which reveals changes in relative demand by campaign donors for ideological positioning among candidates for Congress. The first tercile comprises the most liberal donors, while the third tercile comprises the most conservative donors. By construction, each group accounts for one-third of contributions in the initial year, 2002. Over time, the contribution shares of each group will deviate from one-third, if contributing donors skew to the right and (or) to the left. Such skewing is abundantly evident. The share of contributions by conservative (3^{rd} tercile) donors rises to 0.42 in 2010, a level maintained through 2016. The share of contributions by liberal (1^{st} tercile) donors

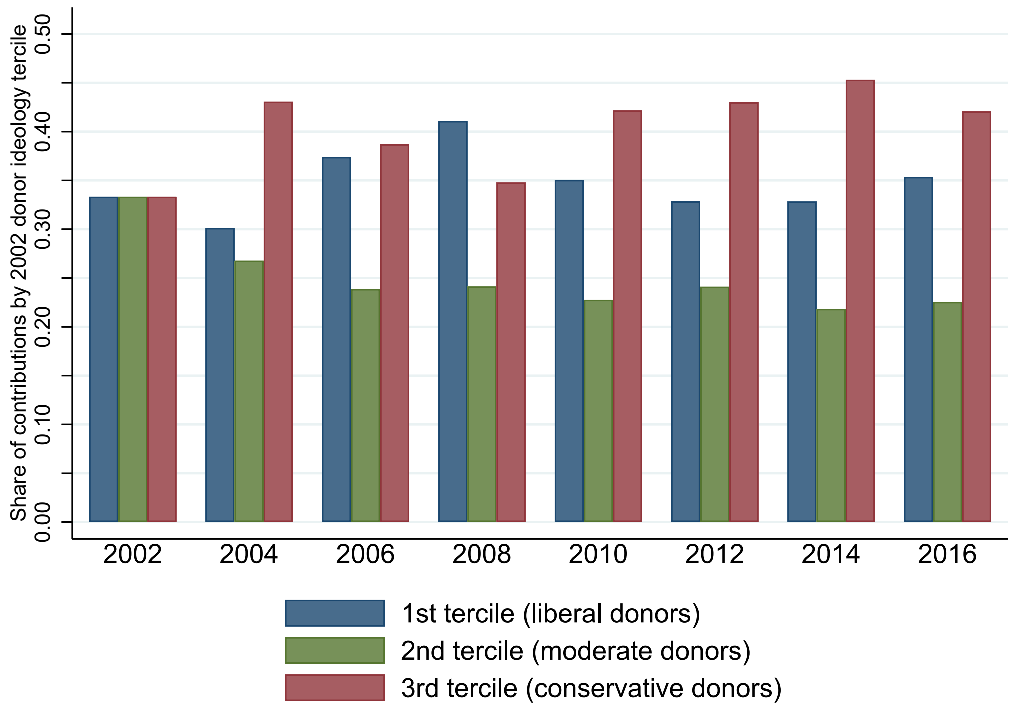
[24] The DIME database similarly provides a series of time-variant CF scores for politicians that are derived from constant donor CF scores. These dyanmic CF scores, which re-estimate the ideology of each politician for each congressional election cycle, do not cover our full period of analysis, however.

[25] The computation of legislator ideal points from roll-call votes faces the challenge that each Congress votes on a different set of bills that represent different topical issues to a varying degree (Bonica, 2017). In order to compare the ideology of legislators who have served in different Congresses and never cast votes on the same bills, one usually has to impose strong parametric restrictions on the change over time in ideology of legislators who served in multiple Congresses, such as a linear time trend in the time-variant DW-Nominate scores of Poole and Rosenthal (2007), which we studied in an earlier version of this paper (Autor et al., 2016b). During the period of 2002 to 2010, there is a high correlation of 0.66 between the change in the time-variant CF score and the change in the linear-trend DW-Nominate, and both legislator ideology scores yield similar results in our empirical analysis. The CF scores, however, allow us to analyze a broader set of outcomes, including changes in campaign contributions by different groups of contributors, and changes ideology among both the winners and losers of Congressional elections.

[26] In monetary terms, FEC-registered contributions to candidates for election in the House of Representatives rose from $729 million in 2002 to $1.2 billion in 2016 (2016 USD), with a spike in 2012 following the *Citizens United v. FEC* case (Dawood, 2015), which relaxed restrictions on campaign financing. See Bonica (2016) on why *Citizens United* did not lead to a more sustained increase in campaign contributions.

first rises to 0.42 in 2008 and then declines to 0.35 in 2010, a level maintained through 2016. These changes imply that the share of contributions by centrist donors has declined over time, dropping to 0.23 in 2010 and remaining at that level through 2016. The composition of campaign contributions has thus become more polarized. Donations have shifted strongly right and modestly left, leaving a relative decline in contributions by political moderates.

Figure 2: Polarization in Campaign Finance Scores for Campaign Donors, 2002 to 2016



Notes: Calculations based on Data Database on Ideology, Money in Politics, and Elections (DIME; Bonica, 2013). Donor ideology is divided into ideology terciles based on campaign contributions in 2002 ranked by dollar-weighted CF scores. The first tercile comprises the most liberal donors, while the third tercile comprises the most conservative donors. The height of each bar in each reported year reflects the share of all contributions (in dollars) falling within each 2002 ideology tercile.

The DIME database identifies whether donors are individuals, corporations, or non-corporate organizations (e.g., labor unions; single-issue, single-candidate or single-party political action committees). Over 2002 to 2016, donations by individuals remained roughly stable at around one-half of all contributions, while the corporation share in contributions fell (from 27.9% to 11.0%) and the non-corporate-organization share rose (from 19.9% to 37.7%). In Appendix Figure A2, we decompose these contributions by donor type according to the same CF-score terciles used in Figure 2.[27]

[27] To facilitate comparisons across donor categories, we continue to define the boundaries of CF terciles across all donors. As a consequence, donations by individuals, corporations, and organizations may be unequally distributed across these terciles, even in the initial year, 2002.

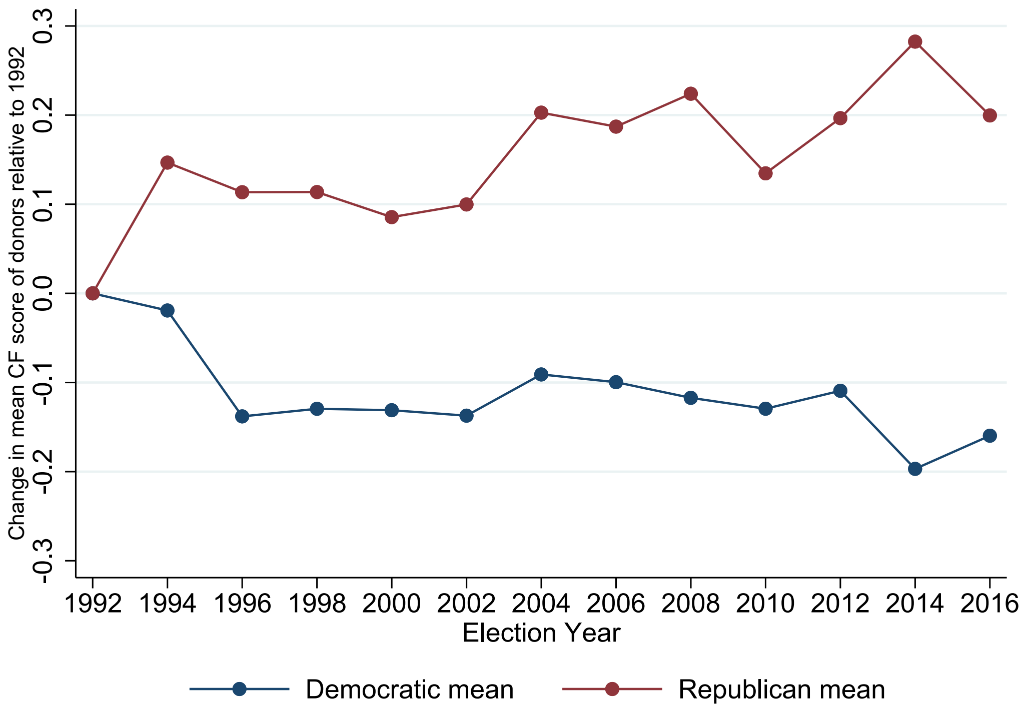
Appendix Figure A2 reveals cleavages in ideological positioning by donor type. While centrist donors dominate among corporations—perhaps reflecting the desire of many corporations to remain in the good graces of whichever party is in control of Congress—liberals and conservatives dominate among individual and non-corporate-organization donors.[28] Over time, the share of moderates in contributions by type fell for both corporations (from 75.6% in 2002 to 67.7% in 2016) and individuals (21.6% in 2002 to 15.0% in 2016), while it rose from low levels among non-corporate organizations (from 6.7% in 2002 to 19.4% in 2016). The share of conservatives in contributions rose most strongly for non-corporate organizations (from 31.0% in 2002 to 41.4% in 2016). In concert, rightward and leftward shifts in aggregate contributions by individual donors have combined with a rightward shift in non-corporate donors and a decline in (relatively moderate) corporate donations to generate the polarization of campaign finance seen in Figure 2.

We turn next to how the ideological positioning of general-election winners has changed over time, as revealed by the identities of their donors. Figure 3 shows the central tendency of contribution-weighted-average CF scores for Democratic and Republican congressional election winners from 1992 to 2016, where we normalize CF scores by the party-specific mean CF score in 1992 in order to highlight between-party polarization.[29] Republican legislators have become more conservative in terms of the donors that support their elections, with their average CF score rising by 0.11 from 1992 to 2002 and by 0.11 again from 2002 to 2016, with the total 1992-2016 change equal to one ($[0.60 - 0.40]/0.20$) standard deviation of the Republican CF score in the initial year. Democrats, for their part, have become more liberal, with average CF scores falling by 0.14 from 1992 to 2002 and by 0.05 from 2002 to in 2016, where the total drop equals a 0.62 ($[0.39 - 0.23]/0.26$) standard-deviation change. Ideological polarization is thus evident whether we examine the ideological composition of campaign donors or average ideology by party among elected representatives.

[28] Bonica (2016) shows that whereas corporate political action committees tend to have moderate CF scores, the scores of corporate executives and directors are decidedly more ideological, with a strong majority of these elites giving to GOP candidates and having relatively high (i.e., conservative) CF scores. In the 2000s and 2010s, nearly all senior executives and directors of Fortune 500 companies gave to political campaigns.

[29] Over 1992 to 2016, the mean CF score for congressional representatives is roughly centered on zero ($\mu = 0.11$, $\sigma = 0.52$ in the average year). Naturally, parties already demonstrated strong ideological differences in 1992. The initial mean CF score for GOP congressional legislators was 0.40 ($\sigma = 0.26$), whereas that for Democrats was -0.23 ($\sigma = 0.20$).

Figure 3: Polarization in Campaign Finance Scores for Congressional Legislators, 1992 to 2016



Notes: Figure reports the central tendency of contribution-weighted-average CF scores for Democratic and Republican congressional election winners from 1992 to 2016, where CF scores are normalized by the party-specific mean CF score in 1992. The initial mean CF score for GOP congressional legislators was 0.40 (σ= 0.26) and -0.23 (σ= 0.20) for Republicans and Democrats, respectively, in 1992.

3 Measuring Local Economic and Political Change

In our analysis of congressional elections, we examine changes over 2002 to 2016 in the ideological positioning of contributors to election campaigns and the candidates who win these elections. Within our period, the 2002 and 2010 elections are respectively the first and last whose congressional district boundaries are based on the 2000 Census. In 2012, states defined new districts, based on population counts in the 2010 Census and the constitutional mandate that each district contain approximately $1/435^{th}$ of the U.S. population. When analyzing 2002 to 2010, we study a period that spans the primary force of the China trade shock (Autor et al., 2016a; Brandt et al., 2017a), and within which district boundaries are stable. When extending our analysis beyond 2010, we address the longer run impacts of economic shocks on electoral outcomes, but must confront the measurement inconsistencies created by redistricting. To balance concerns over measurement error due to redistricting with interest in the persistence of shocks on our outcomes, we study both the 2002-2010 and 2002-2016 periods. In most cases, we omit observations spanning the 2010-2012 seam except for the small set of districts that retain consistent boundaries during this window.

3.1 Local Labor Market Exposure to Trade

Our empirical analysis employs the specification of local trade exposure in commuting zones (CZs) derived by Autor et al. (2014) and Acemoglu et al. (2016a). CZs are clusters of adjoining counties that have the commuting structure of a local labor market (Tolbert and Sizer, 1996; Dorn, 2009). For each CZ j, we measure the shock experienced by a local labor market as the average change in Chinese import penetration in that CZ's industries, weighted by the share of each industry k in the CZ's initial employment:

$$\Delta IP_{j\tau}^{cu} = \sum_k \frac{L_{jkt}}{L_{jt}} \Delta IP_{k\tau}^{cu}. \tag{1}$$

In this expression, $\Delta IP_{k\tau}^{cu} = \Delta M_{k\tau}^{cu}/(Y_{k0} + M_{k0} - X_{k0})$ is the growth of Chinese import penetration in the U.S. for an industry k over period τ, computed as the growth in U.S. imports from China during the outcome period, $\Delta M_{k\tau}^{cu}$, divided by initial absorption (U.S. industry shipments plus net imports, $Y_{k0} + M_{k0} - X_{k0}$) in the base period 1991, near the start of China's export boom. The fraction L_{jkt}/L_{jt} is the share of industry k in CZ j's total employment, as measured in County Business Patterns data prior to the outcome period in the year 2000.

In (1), the difference in $\Delta IP_{j\tau}^{cu}$ across commuting zones stems from variation in local industry employment structure at the start of period t. This variation arises from two sources: differential concentration of employment in manufacturing versus non-manufacturing activities and specialization in import-intensive industries within local manufacturing.[30] In our main specifications, we control for the start-of-period manufacturing share within CZs so as to focus on variation in exposure to trade arising from differences in industry mix within local manufacturing.

An issue for the estimation is that realized U.S. imports from China in (1) may be correlated with industry import-demand shocks. In this case, OLS estimates of the relationship between changes in imports from China and changes in U.S. manufacturing employment may understate the impact of the pure supply shock component of rising Chinese import competition, as both U.S. employment and imports may rise simultaneously in the face of unobserved positive shocks to U.S. product demand. To identify the causal effect of rising Chinese import exposure on local-level political outcomes, we employ an instrumental-variables strategy that accounts for the potential endogeneity of U.S. trade exposure. We exploit the fact that during our sample period, much of the growth in Chinese imports

[30] Differences in manufacturing employment are not the primary source of variation. In a bivariate regression, the start-of-period manufacturing employment share explains less than 40 percent of the variation in $\Delta IP_{j\tau}^{cu}$.

stems from the rising competitiveness of Chinese manufacturers, which is a supply shock from the perspective of U.S. producers. China's lowering of trade barriers (Bai et al., 2017), dismantling of the constraints associated with central planning (Naughton, 2007; Hsieh and Song, 2015), and accession to the WTO (Pierce and Schott, 2016; Handley and Limao, 2017) have contributed to an immense increase in the country's manufacturing productivity and a concomitant rise in the country's manufacturing exports (Hsieh and Ossa, 2016; Brandt et al., 2017a).[31] China's aggressive market opening appears to have ended in the late 2000s (Naughton, 2018; Lardy, 2019), after which point the government took a heavier hand in guiding the country's industrial development.

We identify the supply-driven component of Chinese imports by instrumenting for growth in Chinese imports to the U.S. using the contemporaneous composition and growth of Chinese imports in eight other developed countries.[32] Specifically, we instrument the measured import-exposure variable $\Delta IP^{cu}_{j\tau}$ with a non-U.S. exposure variable $\Delta IP^{co}_{j\tau}$ that is constructed using data on industry-level growth of Chinese exports to other high-income markets:

$$\Delta IP^{co}_{j\tau} = \sum_{j} \frac{L_{jkt-10}}{L_{jt-10}} \Delta IP^{co}_{k\tau}. \tag{2}$$

This expression for non-U.S. exposure to Chinese imports differs from the expression in equation (1) in two respects. In place of U.S. imports by industry ($\Delta M^{cu}_{k\tau}$) in the computation of industry-level import penetration $\Delta IP^{cu}_{k\tau}$, it uses realized imports from China by other high-income markets ($\Delta M^{co}_{k\tau}$) in $\Delta IP^{co}_{k\tau}$, and it replaces all other variables with lagged values to mitigate any simultaneity bias.[33] As documented by Autor et al. (2016a), all eight comparison countries used for the instrumental variables analysis witnessed import growth from China in at least 343 of the 397 total set of four-digit SIC manufacturing industries. Moreover, cro(1)ss-country, cross-industry patterns of imports are strongly correlated with the U.S., with correlation coefficients ranging from 0.55 (Switzerland) to 0.96 (Australia). That China made comparable gains in penetration by detailed sector across numerous countries in the same time interval suggests that China's falling prices, rising quality, and declining trade costs in these surging sectors are root causes of its manufacturing export growth.[34] Because China's market-oriented reforms accelerated with its WTO accession in 2001 and

[31] China may have intentionally undervalued its exchange rate in the early 2000s, which may have contributed to its export growth in the first half of the decade (Bergsten and Gagnon, 2017).

[32] The eight other high-income countries are those that have comparable trade data covering the full sample period: Australia, Denmark, Finland, Germany, Japan, New Zealand, Spain, and Switzerland.

[33] The start-of-period employment shares L_{jkt}/L_{jt} are replaced by their 10-year lag, while initial absorption in the expression for industry-level import penetration is replaced by its 3-year lag.

[34] A potential concern for our analysis is that it ignores U.S. exports to China, focusing instead on trade flows in the opposite direction. This focus is dictated by the fact that our instrument, by construction, has less predictive power for U.S. exports to China. To a first approximation, China's economic growth during the 1990s and 2000s generated a substantial shock to the supply of U.S. imports but only a modest change in the demand for U.S. exports. During our sample period, imports from China were nearly five times as large as manufacturing exports from the U.S. to China.

had largely run their course by the end of the 2000s, we define our measure of the China trade shock in (1) and our instrument for this shock in (2) to span the period 2002 to 2010.

Data on international trade are from the UN Comtrade Database, which gives bilateral imports for six-digit HS products.[35] To concord these data to four-digit SIC industries, we first apply the crosswalk in Pierce and Schott (2012), which assigns ten-digit HS products to four-digit SIC industries (at which level each HS product maps into a single SIC industry), and then aggregate to six-digit HS products and four-digit SIC industries (at which level some HS products map into multiple SIC entries). For this aggregation, we use data on U.S. import values at the ten-digit HS level, averaged over 1995 to 2005. Dollar amounts are inflated to dollar values in 2015 using the PCE deflator. Data on CZ employment by industry from the County Business Patterns for the years 1990 and 2000 is used to compute employment shares by industry in (1) and (2).

Appendix Table A3 summarizes trade exposure defined at the CZ level and then matched to, variously, Pew survey respondents by CZ of residence, Nielsen households by CZ of residence, electoral outcomes in county-congressional district cells, and voting results by CZ in presidential elections. Through most of our analysis, we use the 2002 to 2010 period to characterize the rise in import competition from China.[36] This period corresponds with China's post-WTO-accession productivity boom and its most intense increase in import penetration in the U.S. On average, Chinese import penetration grew by 0.71 percentage points between 2002 and 2010 (column 1 of Appendix Table A3). In section 4, we use the interquartile range of the increase in trade exposure as a metric to scale estimated treatments of trade exposure on political outcomes in more versus less-exposed districts. This range is 0.49 percentage points across the full set of congressional districts in our analysis.

3.2 Political Outcomes in County-by-Congressional-District Cells

To map economic outcomes in commuting zones to political outcomes in congressional districts, we define the geographic unit of our main analysis to be the county-by-congressional-district cell. The building blocks of congressional districts are census tracts, whose amalgamation allows officials to construct districts that meet the requirements of contiguity and equal population size within a state. The resulting map of congressional districts frequently splits counties and CZs between multiple districts. We overlay this map with the map of county boundaries to obtain county-by-district cells, which allow us to nest the geographies of economic and political outcomes. We ascribe to each county-district cell the CZ-level trade shock that corresponds to the county and weight each

[35] See http://comtrade.un.org/db/default.aspx.

[36] The 2002 to 2010 window is the initial outcome period for our analysis of congressional elections. We also study presidential elections starting from 2000 to 2008 and beyond, and correspondingly measure the growth of import competition over the period of 2000 to 2008 there.

cell by its share of the voting-age population in the district, such that each congressional district has equal weight in the analysis. If a district spans multiple CZs, the economic factors that are mapped to the district will be a population-share-weighted average of the values in these CZs.[37]

To illustrate how we construct country-by-congressional-district cells, consider North Carolina's 12th congressional district, which connects parts of the cities of Charlotte, Greensboro and Winston-Salem along a narrow corridor (see Appendix Figure A3). Rowan County overlaps with the 12th district in its center, but also with the 5th district in its Northwest, and with the 8th district in its Southeast. Our data contain a separate observation for each of these county-district cells. To each cell, we attach information on the elected representatives for the corresponding district (for cells in Rowan Country from the 5th, 8th, or 12th), and the economic conditions of the commuting zone (Charlotte) to which the county (Rowan) belongs. In our analysis, the weight attached to each cell equals the cell's share of the voting-age population in its corresponding congressional district.

Data on election outcomes in county-district cells are from Dave Leip's Atlas of U.S. Elections, which tracks votes received by Democratic, Republican, and other candidates for Congress in each county within each congressional district, and in each election year. We use these data to tabulate the shares of votes won by Democratic and GOP congressional candidates in each county-by-district cell in 2002, as well as the change in these values between 2002 and later years. The Leip data also provides the number of registered voters at the county level, which allows us to compute voter turnout by county.[38] The information from the DIME database (Bonica 2013, 2014) on campaign contributions and the inferred ideology of congressional legislators matches into our county-by-district geography at the congressional district level.

In addition to our analysis of congressional elections, we also study Leip data on vote shares in presidential elections, Nielsen data on TV consumption, and survey responses from PEW. All of these data are reported at the level of counties, and do not depend on the (changing) boundaries of congressional districts. In our analysis of presidential voting, we use county-level vote shares for the nominees of the two major parties in 2000, 2008, and 2016.

3.3 Adjusting for Redistricting

The main period for our analysis of congressional elections is 2002 to 2010. This period encompasses the most rapid rise of import competition from China, and the measurement of changes in district-level outcomes is facilitated by stable district boundaries in almost all states. Appendix Table A4

[37]From the full sample of 435 congressional districts, we omit Alaska's one congressional district and Hawaii's two congressional districts because CZs are difficult to define for these states. The resulting set of 3,772 county-district cells covers all 432 congressional districts on the U.S. mainland.

[38]Data on registered voters are missing in some years for Georgia, Mississippi, North Dakota and Wisconsin. These four states are omitted from our empirical analysis of voter turnout.

shows the extent of redistricting in congressional elections from 2002 to 2016. In the period of 2002 to 2010, only four states implemented adjustments to their district boundaries. Between 2010 and 2012, however, nearly all districts on the US mainland (425 out of 432) changed boundaries. To extend our analysis beyond 2010 by including outcomes from congressional elections in 2012, 2014, and 2016, we thus need to account for the sweeping congressional restricting of 2012.

To match county-by-congressional-district cells across time, we construct a new crosswalk that apportions county-district cells for the 113^{th} Congress (elected in 2012), and the next two congresses, to county-district cells as defined for the 108^{th} Congress (elected in 2002). We begin by splitting each county-district cell of the 113^{th} Congress into Census Blocks of the 2010 Census. We next create a weighted crosswalk between 2010 Census Blocks and 2000 Census Blocks, which indicates the fraction of population of a 2000 Block that maps into the boundaries of a given 2010 Block. The 2000 Blocks in turn can be mapped to county-district cells for the 108^{th} Congress. We finally aggregate the Block-to-Block crosswalk to the level of county-district cells, such that the final crosswalk indicates the fraction of population (measured in 2000) of a county-district cell for the 108^{th} Congress whose location of residence falls into a given county-district cell for the 113^{th} Congress. We also construct similar crosswalks to account for several intracensal period episodes of redistricting of individual states in the elections of 2004, 2006 and 2016.[39]

The crosswalks allow us to map outcomes from years following redistricting into the boundaries of the initial county-district cells for the 108^{th} Congress. However, we need to additionally address the fact that redistricting elevates churning in political outcomes. Panel B of Appendix Table A4 indicates the fraction of congressional districts that replaced a Democratic Representative with a Republican or vice versa, separately for districts that changed boundaries and those that did not. In each election following redistricting (in 2004, 2006, 2012, and 2016), districts with boundary changes experienced much greater levels of party churning than those whose geography remained unchanged. Averaging over these elections, one out of every six districts with boundary changes (15.7%) switched parties, while only one out of every 26 districts without boundary change (3.8%) did so. The much higher churning in the former group of districts is likely a consequence of redistricting, rather than an expression of rapidly changing political preferences among voters in these districts.

Consider the example of Montgomery County, Alabama. From the 108^{th} Congress (elected in 2002) to the 112^{th} Congress (elected in 2010), the northwestern part of the county, which includes most of the state capital city of Montgomery, belonged to Alabama's 2nd district while the more rural southwestern part belonged to the 3rd district. Both districts elected GOP candidates in 2010, and after redistricting in 2012, the entire congressional delegation of Alabama was reelected. Despite this

[39] To construct our crosswalks of county-district cells, we draw on data of the Census Bureau, the Missouri Census Data Center, and the IPUMS National Historical Geographic Information System.

maximum stability in election outcomes from 2010 to 2012, some residents of Montgomery county were no longer represented by the same politician or even the by same party after the 2012 election. The 2012 redistricting moved large swaths of inner-city Montgomery from the 2nd and 3rd to the 7th district, which is Alabama's only district with a majority black population. According to our county-district crosswalk, 15% of the Montgomery county residents who belonged to the 2nd district until 2010 found themselves in the Democrat-controlled 7th district as of 2012. This change in party is likely a mechanical outcome of redistricting, and not informative about changes in the political views of Montgomery county residents. To purge the considerable noise caused by redistricting, we compute changes in outcome variables that omit any two-year period during which a district changed its boundaries. Our outcome variables thus take the form,

$$\Delta Y^{r}_{cd\tau} = \sum_{t \in \tau} (1 - R_{dt+2}) \left(\sum_{d'} \frac{p_{cdd'}}{p_{cd}} Y_{cd't+2} - \sum_{d'} \frac{p_{cdd'}}{p_{cd}} Y_{cd't} \right) \quad (3)$$

where $\Delta Y^{r}_{cd\tau}$ is the redistricting-adjusted change of an outcome Y over a period τ for the cell of county c and district d of the 108^{th} Congress. The variable $Y_{cd't}$ indicates the level of the same outcome variable in a year t that is the start year of a two-year period contained in period τ. It is measured for county c and the districts d' that are used during the election in year t. The fraction $\frac{p_{cdd'}}{p_{cd}}$ indicates the population share of the initial county-district cell cd that maps to the new county-district cell cd', and R_{dt+2} is an indicator variable that takes a value of one if district d experienced boundary changes in election year $t+2$.[40] Consider as an example the change in the Republican vote share over the period 2002 to 2016 for the overlap between Montgomery county and Alabama's 2nd district of the 108^{th} Congress. Adding up over the four two-year periods from 2002 to 2010, the Republican vote share declined from 64% in 2002 to 41% in 2010 in this cell. We omit the two-year change during redistricting in 2010-2012, and then compute the subsequent change in Republican vote share from 2012 to 2016 as a weighted average of the change in the 85% overlap of the cell with the new 2nd district (where the Republican share increased from 47% in 2012 to 58% in 2016) and the 15% overlap the new 7th district (where the Republican vote share declined from 4% in 2012 to 0% in 2016). The redistricting-adjusted change in Republican vote share during the 2002 to 2016 period is thus $(41\% - 64\%) + 0.85(58\% - 47\%) + 0.15(0\% - 4\%) = -14\%$.

If a district experienced no boundary changes during outcome period τ, then $d' = d$, $\frac{p_{cdd'}}{p_{cd}} = 1$, and $R_{dt+2} = 0$, so that equation (3) simplifies to $\Delta Y^{r}_{cd\tau} = \sum_{t \in \tau}(Y_{cdt+2} - Y_{cdt})$. This sum of two-year changes contained in period τ is equivalent to the first difference of outcome variable Y between the start and end of period τ. Since there was little redistricting from 2002 to 2010, outcome

[40] We restrict the value of $\Delta Y^{r}_{cd\tau}$ to lie within the range of values that could be observed for a non-adjusted change of that outcome. For instance, adjusted changes in party vote shares are restricted to -100% or +100% in the rare cases where equation (3) yields an adjusted change of more than 100%.

variables during our main period of analysis correspond to simple first differences in most states.

4 Impact of Trade Shocks on Political Expression

We address the political consequences of exposure to greater import competition from China in three stages: in this section, by examining changes in media viewership, political beliefs, and campaign contributions; in section 5, by considering congressional election outcomes; and in section 6, by assessing presidential voting. This sequence of analyses allows us to evaluate how greater trade exposure affects political expression, resource mobilization for electoral campaigns, and the political orientation of candidates who win congressional elections and presidential contests.

4.1 Cable News Market Shares for FOX, CNN, and MSNBC

We begin our analysis of how trade exposure affects political expression by using our Nielsen data on the cable-news-viewing habits of U.S. households. We prioritize the Nielsen data over the Pew survey data used in the next section, given the much larger sample sizes in the former (99k to 119k households per time period) relative to the latter (2k to 10k respondents per time period). The rankings of cable news channels indicate household relative demand for ideological content. According to fivethirtyeight.com, the percentage of FOX News viewers who voted for the GOP presidential candidate exceeded the percentage of viewers who voted for the Democratic candidate by the stunning margins of 62% in 2004 and 66% in 2016.[41] Viewers of CNN and MSNBC tend to lean Democratic. CNN and MSNBC viewers favored the Democratic over the GOP presidential candidate by 32% and 28%, respectively, in 2004, and by 47% and 70%, respectively, in 2016.

Because the Nielsen data contain repeated cross sections of households, we cannot study changes in viewership for a longitudinal panel of viewers. Instead, we use aggregate data on households by CZ and age-race groups to study whether news viewership changed differentially in CZs that faced greater trade exposure. For this analysis, we aggregate the Nielsen data to the level of CZ i by age-race group g (based on the household head ages 18-34, 35-54, 55+ for non-Hispanic whites and those with other race/ethnicity) by time period t (weekday prime-time hours during 28-day windows in November of two presidential election years $t = t_1$ and $t = t_2$).[42] Our regression equation is:

$$Y_{jgt} = \gamma_j + \gamma_g + \gamma_t + \beta_1 \Delta IP^{cu}_{j\tau} \times 1[t = t_2] + \gamma_g \times 1[t = t_2] + \beta_2 X_{jt_1} \times 1[t = t_2] + \epsilon_{jgt}. \qquad (4)$$

The dependent variable Y_{jgt} is either the combined rating of the three major news channels or the

[41] See https://fivethirtyeight.com/features/how-roger-ailes-polarized-tv-news/.

[42] As discussed in Section 2, cable news networks have high ratings during presidential elections. These periods may be particularly indicative of viewers' political leanings, as news programming is dominated by election coverage.

cable-news market share of a given channel (both in percentage points), measured for each CZ and age-race group in two time periods (November 2004 and November 2008, 2012, or 2016). We control for CZ, age-race group and time-period main effects $(\gamma_j, \gamma_g, \gamma_t)$ and interact the age-race-group indicators with the time dummy $(\gamma_g \times 1[t = t_2])$ to allow for time trends in TV preferences within these groups. The 2002-2010 import shock $\Delta IP^{cu}_{j\tau}$ is interacted with a dummy variable $1[t = t_2]$ indicating the end period of the analysis and we allow for region-specific time trends in a vector of control variables X_{jt_1} via an interaction with the time dummy $1[t = t_2]$. This control vector includes dummy variables for the Census geographic region to which CZ j belongs, start-of-period economic conditions in CZ j (the share of manufacturing in CZ employment, the offshorability index and the Autor and Dorn (2013a) routine-task-intensity index for CZ occupations, each measured in 2000) and start-of-period political conditions in CZ j (the two-party vote share of the Republican nominee in the 1996 and 2000 presidential elections).

We first consider the years 2004 and 2012, a time period that overlaps with the 2002 to 2010 period for which we begin our analysis of congressional elections.[43] Panel A of Table 2 shows that greater exposure to import competition triggers little change in the combined rating of the three cable news channels. The coefficient on the trade-shock, end-year interaction is positive and marginally significant in the column 3 regression only, and falls to near zero and becomes highly imprecisely estimated when full controls are added in column 6. These results indicate that there is no apparent effect of trade shocks on households' *overall* consumption of TV news, even though cable news viewership does rise in the aggregate across CZs during the sample period.

[43]Nielsen data were not available prior to 2004.

Table 2: Exposure to Chinese Import Competition and Cable TV News Viewership, November 2004-November 2012. Dependent Variable: Change in TV Rating or News Channel Market Share (in % pts)

	(1)	(2)	(3)	(4)	(5)	(6)
	A. Combined Nielsen Rating of TV News Networks					
Δ CZ Import Penetration x	-0.18	-0.10	0.67	0.54	0.16	0.03
[t=2012]	(0.09)	(0.17)	(0.37)	(0.37)	(0.40)	(0.41)
	B. Market Share FOX News					
Δ CZ Import Penetration x	2.86	5.87	5.09	7.31	8.32	10.48
[t=2012]	(1.49)	(2.77)	(5.03)	(4.81)	(5.21)	(5.30)
	C. Market Share CNN					
Δ CZ Import Penetration x	-0.51	-1.28	-6.51	-5.37	-4.34	-4.18
[t=2012]	(1.25)	(2.03)	(3.73)	(3.44)	(3.47)	(3.44)
	D. Market Share MSNBC					
Δ CZ Import Penetration x	-2.34	-4.58	1.42	-1.95	-3.98	-6.28
[t=2012]	(1.47)	(1.96)	(3.94)	(3.14)	(3.46)	(3.72)
Estimation Method	OLS	2SLS	2SLS	2SLS	2SLS	2SLS
F-statistic First Stage		45.4	40.6	38.3	27.7	27.7
CZ FE	yes	yes	yes	yes	yes	yes
Year FE	yes	yes	yes	yes	yes	yes
Age-Race Group FE	yes	yes	yes	yes	yes	yes
CZ Industry/Occ x [t=2012]			yes	yes	yes	yes
Pres. Election Ctrls x [t=2012]				yes	yes	yes
Census Divisions x [t=2012]					yes	yes
Age-Race Group FE x [t=2012]						yes

Notes: $N = 6,923$ CZ-year-age-race cells in Panel A and N=5,079 cells in panels B-D. The Combined Nielsen Rating in Panel A indicates the percentage of households that own TVs that were watching one of the three major TV news networks. Panels B-D indicate the market share of each major TV news network in their combined market. In November 2004, the average combined rating was 2.5%, and the TV news market shares were 59.2% for FOX News, 27.7% for CNN, and 13.1% for MSNBC. All regressions control for Commuting Zone, year, and age-race (three age times two race groups) fixed effects. Industry and occupation controls in column 3 include the fraction of CZ employment in the manufacturing sector and the Autor and Dorn (2013b) routine share and offshorability index of a CZ's occupations, all of which are measured in 2000 and interacted with the dummy for the 2012 period. Election controls in column 4 comprise the Republican two-party vote shares of the CZ in the presidential elections of 1996 and 2000, each interacted with the period dummy. Census division dummies interacted with the period dummy in column 5 allow for different time trends across the nine geographical Census divisions. Age-race-time interactions in column 6 allow for different time trends by age-race group. Observations are weighted by Nielsen's estimate of the number of TV households in each cell, and standard errors are clustered on Commuting Zones.

In panels B to D of Table 2, we examine how greater trade exposure affects the market share of individual cable-news channels. Consider the results for FOX News in panel B. In column 1, we estimate a parsimonious OLS regression that controls for CZ, age-race group, and year fixed effects only. The coefficient estimate is positive and significant at the 10% level ($t = 1.92$). Turning to the 2SLS regression in column 2, the trade-shock coefficient estimate doubles in magnitude and becomes more precisely estimated ($t = 2.12$). In the Autor et al. (2013) analysis of the labor market impact of increased import competition from China, instrumental variables regressions consistently indicate more adverse impacts of trade than OLS regressions. To the extent that import shocks affect political beliefs via deteriorating labor market conditions, one would expect the greater impact of imports on ideology when moving from OLS in column 1 to 2SLS in column 2. The column 2 estimates indicate that CZs with a one-percentage-point larger increase in trade exposure had a 5.9 percentage-point larger FOX News market share in 2012 relative to 2004, a period during which FOX News' presidential-election-month ratings rose but its market share fell modestly. Once we include the full set of controls for economic and political conditions in column 5, the trade-shock impact rises to 10.5 percentage points ($t = 2.00$), which implies that when comparing CZs at the 75^{th} versus 25^{th} percentiles of trade exposure, the former would have a 5.2 percentage-point (10.5×0.49) larger increase in the market share of FOX News.

That trade shocks have a positive impact on the FOX News market share implies that they must diminish market shares for CNN and (or) MSNBC, evidence for which we see in panels C and D of Table 2. The results in column 5 with full controls indicate that approximately three fifths ($-6.3/10.5$) of the FOX News gain in market share in trade-impacted CZs was at the expense of MSNBC while two-fifths ($-4.2/10.5$) of the FOX gain was at the expense of CNN, although only the first impact reaches the 10% significance level ($t = 1.70$).[44] We interpret these results to mean that greater regional exposure to import competition caused an increase in the relative demand for television news with a conservative political slant.

In Appendix Table A5, we expand the analysis to the 2004-2008 and 2004-2016 periods, using the specification in column 6 of Table 2 with full controls. The results in Table 2 are fully replicated for these alternative horizons. Greater exposure to import competition yields no change in cable-news viewership overall, while it does reallocate market share to FOX News from MSNBC and CNN. The impact of the trade shock on FOX News market shares for the 2004-2008 ($\gamma_1 = 8.8$, $t = 2.1$) and 2004-2016 ($\gamma_1 = 10.5$, $t = 2.7$) time periods are similar to that for 2004-2012 ($\gamma_1 = 10.5$, $t = 2.0$), indicating that four-fifths of the long-run trade-shock-induced impact on FOX News was realized by 2008, by which point the China trade shock itself had almost entirely unfolded.

Motivated by the differential trends in cable news ratings for households separated by race and ethnicity in Figure A1, in Appendix Table A6 we report regressions in which we interact the trade

[44] In supplementary estimates, we expand the sample to include all ratings months (February, May, July, November). The impact of import competition on the FOX News market share has a slightly smaller magnitude and remains precisely estimated when we include non-presidential-election months in the analysis.

shock with dummy variables for the six age-race groups, where the specifications are otherwise the same as in column 6 of Table 2. While the trade shock leads to an increase in the FOX News market share for most age-race groups, these gains tend to be larger and more precisely estimated for non-Hispanic whites. For the 2004 to 2016 period, these impacts—which range from 12.4 percentage points ($t = 2.1$) in FOX ratings gain per percentage point of extra import shock for the 18-34 year-old white age group to a 13.5 percentage-point gain ($t = 2.7$) for the 35-54 year-old white age group and to a 12.6 percentage-point gain ($t = 2.8$) for the 55+ year-old white age group—are $2.7 = (13.5/7.3)$ to $3.8 = (12.6/3.3)$ times larger than the imprecisely estimated impacts for the corresponding non-white and non-Hispanic age groups. The trade-induced gains in market share for right-leaning FOX News thus appear to be concentrated among non-Hispanic-white households.

4.2 Political Beliefs in Pew Survey Responses

Next, we examine how rising trade exposure affects expressed political beliefs using the Pew data presented in Section 2.1. We include surveys in 2004, 2011, 2014, and 2015, where we treat the latter three years as a single time period to maximize sample size. Our local labor market approach puts high demands on the data as we observe only 25 observations on average per CZ.[45] Having this caveat in mind, we proceed with an analysis that follows the structure of (4) by estimating an equation of the form:

$$Y_{ijt} = \gamma_j + \gamma_1 \Delta IP^{cu}_{j\tau} \times 1[t = t_2] + Z^{'}_{ijt} (\gamma_3 + \gamma_4 \times 1[t = t_2]) + X^{'}_{jt_1} \gamma_6 \times 1[t = t_2] + \epsilon_{ijt}, \qquad (5)$$

where the dependent variable Y_{ijt} is the Pew ideology score (on a scale of -10 to $+10$, from more liberal to more conservative) for survey participant i who resided in CZ j and who was interviewed in survey year t, with $t_1 = 2004$ and $t_2 = \{2011, 2014, 2015\}$; γ_j is a fixed effect for CZ j; and $1[t = t_2]$ is a dummy variable for the second time period. The main variable of interest is the change in import exposure $\Delta IP_{j\tau}$ in CZ j over 2002 to 2010, for which we instrument using (2). The control variables include Z_{ijt}, a vector of characteristics corresponding to participant ijt (a quadratic in age and dummy variables for gender, race, and three categories of education); and X_{jt_1}, the set of regional dummies and initial conditions used in equation (4). As in (4), we include CZ main effects and time-varying coefficients in equation (5) to examine whether average political beliefs change systematically over time within CZs as a function of CZ trade exposure.

[45] The pooled sample of $20,914$ participants includes 650 commuting zones that appear in at least one of the years and 412 commuting zones that appear both in both time periods (i.e., 2004 and at least one of 2011, 2014 or 2015).

Table 3: Exposure to Chinese Import Competition and Pew Ideology Scores, 2004 to 2011/14/15. Dependent Variable: Change in Pew Ideology Score

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Δ CZ Import Penetration x [t>2010]	0.38	1.24	1.61	1.30	1.30		
	(0.27)	(0.67)	(0.70)	(0.77)	(0.77)		
Δ CZ Import Penetration x [t>2010]						0.58	1.35
x Non-Hispanic White						(0.24)	(0.70)
Δ CZ Import Penetration x [t>2010]						-0.36	0.91
x Hispanic or Black						(0.33)	(0.75)
Wald Test Equal Coefficients						p<0.01	p<0.14
Estimation Method	2SLS	2SLS	2SLS	2SLS	2SLS	2SLS	2SLS
F-statistic First Stage	40.8	35.1	34.7	26.4	26.9	39.0	28.9
CZ FE	yes	yes	yes	yes	yes	yes	yes
Year FE	yes	yes	yes	yes	yes	yes	yes
Demographic Controls	yes	yes	yes	yes	yes	yes	yes
CZ Industry/Occ x [t>2010]		yes	yes	yes	yes		yes
Pres. Election Ctrls x [t>2010]			yes	yes	yes		yes
Census Divisions x [t>2010]				yes	yes		yes
Demographic Controls x [t>2010]					yes		yes

Notes: $N = 20,914$ in columns 1-5, $N = 19,556$ in columns 6-7. The Pew Ideology Score has a minimum of -10 (most liberal) and maximum of +10 (most conservative). Controls for individual demographics include a quadratic in age and indicators for sex, three race/ethnicity groups (non-Hispanic whites, Hispanics, all others), and three education groups (college, some college, high school and less). Industry and occupation controls in column 2 include the fraction of CZ employment in the manufacturing sector and the Autor and Dorn (2013b) routine share and offshorability index of a CZ's occupations, all of which are measured in 2000 and interacted with the dummy for the 2011/14/15 period. Election controls in column 3 comprise the Republican two-party vote share in the presidential elections of 1996 and 2000, measured at the county level and interacted with the period dummy. Census division dummies interacted with the period dummy in column 4 allow for different time trends across the nine geographical Census divisions. Demography interactions in column 5 interact the demographic control variables with the dummy for the 2011/14/15 period. Models in columns 6 and 7 retain only individuals who are white, Hispanic or black. Observations are weighted by each individual's share in the sum of Pew survey weights of a given year, and standard errors are clustered on CZs.

Results for the Pew sample data appear in Table 3. In Column 1 of Table 3, which presents a parsimonious 2SLS regression, the coefficient on the interaction between the CZ trade shock and the second-period dummy is positive but small and not precisely estimated ($t = 1.41$). The coefficient magnitude increases substantially in value and becomes more precisely estimated when adding controls for initial economic conditions in column 2 ($t = 1.85$) and political conditions in column 3 ($t = 2.30$). The addition of full controls in column 5—for the Census region and interactions between individual demographic characteristics and the second-period dummy—reduces the trade-shock coefficient somewhat and leaves it marginally significant ($t = 1.69$). These results suggest that demographically comparable survey respondents residing in commuting zones that were subject to larger increases in Chinese import competition in the 2000s became more likely to express conservative political beliefs over the course of a decade. The magnitude of the coefficient estimate in column 5 indicates that if we compare CZs at the 75^{th} and 25^{th} percentiles of trade exposure, the Pew ideology score would be predicted to increase by 0.65 points (1.30×0.49) between 2004 and 2011/14/2015, or one more right-leaning answer for every three survey respondents, in a CZ at the 75^{th} versus the 25^{th} percentile of trade exposure.[46]

Table A2 suggests that the rightward shift in political beliefs over the 2000s was stronger among non-Hispanic whites than among other racial and ethnic groups. In the final columns of Table 3, we show results in which we estimate separate trade-shock coefficients for non-Hispanic-white participants versus Hispanic or non-white participants. Whereas the interaction between trade exposure and the second-period dummy is positive and precisely estimated for whites ($t = 2.41$ for partial controls in column 6; $t = 1.93$ for full controls in column 7), for racial and ethnic minorities it ranges from negative to positive and is imprecisely estimated in each case. These findings, though estimated on a relatively small set of Pew survey datasets, provide suggestive evidence that trade shocks have engendered a stronger rightward shift in political beliefs among whites than among non-whites, consistent with our findings on FOX News market shares using Nielsen viewership data.

4.3 Ideology of Congressional Campaign Donors

We turn now to analysis of the effect of rising trade exposure on political expression as represented by the contributions of campaign donors. Contributions reveal support for candidates that arises from their appeal to donors, where larger contributions indicate, in part, a stronger ideological match between the candidate and the donor (Bonica, 2014; McCarty et al., 2016). Because we know the ideology of donors via their CF scores, we can use the distribution of contributions across scores to assess the total demand for candidates at different points along the ideological spectrum.

In this analysis, and in our later analysis of congressional and presidential elections, we estimate equations of the form:

$$\Delta Y_{cdj\tau} = \gamma + \beta_1 \Delta IP^{cu}_{j\tau} + X^{'}_{cdjt}\beta_2 + e_{cdj\tau}. \qquad (6)$$

The dependent variable $\Delta Y_{cdj\tau}$ is the change in an outcome for time period τ (2002 to 2010 in our baseline specifications) that corresponds to county-congressional-district cell cd in CZ j. To our trade-exposure measure $\Delta IP_{j\tau}$, we pair an expanded vector of regional controls X_{cdjt}, which

[46] With the interquartile range of import exposure equal to 0.49, an increase in the ideology score of 0.65 corresponds to one in every three respondents changing an answer from the left-leaning to the right-leaning position.

includes Census-region dummies and initial CZ economic and political conditions, as in regression equations (4) and (5), and now start-of-period demographic characteristics (population shares for nine age and four racial groups, shares of the population that are female, college educated, foreign born, and Hispanic, each measured at the county level).

We estimate (6) using as the dependent variable the change in campaign contributions for primary and general elections combined. By incorporating donations to both primary and general election campaigns, we capture the total demand for candidate ideology expressed during an electoral cycle. To aggregate contributions, we define bins based on quantiles for CF scores in 2002, match each donor to the bin to which the donor's CF score corresponds, and sum contributions across donors in each bin in each year for each district. To allow for zero values in some cells, we measure the change in contributions ΔC_{bjt} for bin b in district d between time periods t_1 and t_2 as,

$$\Delta C_{bd\tau} = \frac{C_{bdt_2} - C_{bdt_1}}{0.5 \times [C_{bdt_2} + C_{bdt_1}]},$$

which approximates the log change in the value.

4.3.1 Baseline Results for 2002-2010

We begin by studying the 2002-2010 period, during which congressional district boundaries are stable. Panel A of Table 4 shows the impact of greater trade exposure on the 2002-to-2010 change in campaign contributions for all donor types across primary and general congressional elections. In column 1 of panel A, which includes no controls, districts with larger increases in trade exposure have larger increases in contributions, where this impact is significant at the 10% level ($t = 1.71$). As we add controls for initial economic conditions in column 2, geographic region in column 3, demographic characteristics in column 4, and political conditions in column 5, the coefficient estimate doubles in magnitude and remains marginally significant. The estimate with full controls in column 5 of panel A ($t = 1.77$) implies that if we compare congressional districts at the 75^{th} versus 25^{th} percentiles of trade exposure, the more-exposed district would have a 18.2% (37.2×0.49) larger increasein campaign contributions, which is equivalent to a 0.23 standard-deviation change in contributions across all districts over 2002 to 2010. If higher donations indicate more fiercely contested campaigns, then these results suggest that greater trade exposure increases campaign intensity.

Table 4: Exposure to Chinese Import Competition and Campaign Contributions, 2002-2010. Dependent Variable: Relative Change in Contributions (in log points)

	(1)	(2)	(3)	(4)	(5)
	A. All Contributions				
Δ CZ Import Penetration	15.84	28.69	33.30	38.97	37.23
	(9.24)	(21.30)	(21.12)	(21.23)	(21.05)
	B. Left-Wing Contributions (1st Tercile of Donor CF Score)				
Δ CZ Import Penetration	25.22	58.26	64.68	71.05	71.00
	(11.69)	(27.75)	(31.31)	(31.10)	(31.06)
	C. Moderate Contributions (2nd Tercile of Donor CF Score)				
Δ CZ Import Penetration	8.77	14.34	22.19	25.29	23.60
	(8.15)	(18.73)	(18.62)	(19.70)	(19.63)
	D. Right-Wing Contributions (3rd Tercile of Donor CF Score)				
Δ CZ Import Penetration	9.30	34.02	44.97	49.49	46.05
	(13.00)	(26.94)	(26.98)	(27.37)	(27.15)
Estimation Method	2SLS	2SLS	2SLS	2SLS	2SLS
F-statistic First Stage	35.8	37.0	27.3	29.1	29.2
2000 Industry/Occ Controls		yes	yes	yes	yes
Census Division Dummies			yes	yes	yes
2000 Demographic Controls				yes	yes
1996/2000 Pres. Election Ctrls					yes

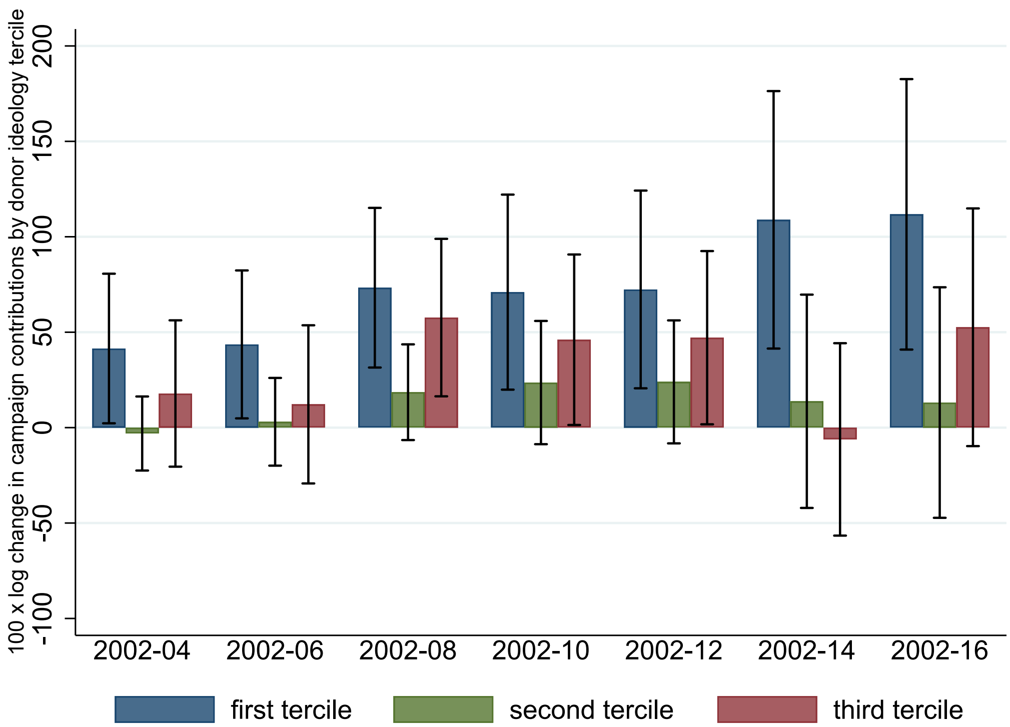
Notes: N=3,772 county-district cells. Panels B to D indicate the change in contributions from donors whose CF score falls into the first, second, and third tercile of the dollar-weighted distribution of donor ideology in 2002. Industry and occupation controls in column 2 are measured at the CZ level and comprise the fraction of CZ employment in the manufacturing sector and the Autor and Dorn (2013b) routine share and offshorability index of a CZ's occupations. Census division dummies in column 3 allow for different time trends across the nine geographical Census divisions. Demographic controls in column 4 comprise the percentage of a county's population in nine age and four racial groups, as well as the population shares that are female, college-educated, foreign-born, and Hispanic. Election controls in column 5 comprise the Republican two-party vote share in the presidential elections of 1992 and 1996, measured at the county level. Observations are weighted by a county-district cell's share in the total year-2000 voting age population of a district, so that each district has a total weight of one. Standard errors are two-way clustered on counties and congressional districts.

The remainder of Table 4 shifts the analysis to how trade shocks affect the ideological composition of campaign donations. Consider first Panel C, which shows the impact of import competition on campaign contributions by relatively moderate donors, those whose CF scores fall in the middle tercile of CF scores as of 2002. In all specifications, the coefficient estimate is small relative to the panel A estimates, and imprecisely estimated (t = 1.20 with full controls in column 5). By contrast, panel B shows that greater trade exposure increases contributions by left-leaning donors, defined as the sum of contributions by donors whose CF scores fall in the first tercile (most liberal) of 2002 CF scores. This impact is positive and precisely estimated in all 2SLS specifications. The coefficient estimate in column 5 ($t = 2.29$) indicates that when comparing more-versus-less trade-exposed congressional districts, the more-exposed district would have an approximately 35% (71.0×0.49) larger increase in campaign contributions by left-leaning donors, which is equivalent to a 0.33 standard-deviation change in first-tercile contributions across all districts over 2002 to 2010. In panel D, we see a qualitatively similar pattern of impacts for contributions by right-leaning donors, defined as total contributions by donors whose CF scores fall the third tercile (most conservative) of 2002 CF scores. In column 5, the marginally significant coefficient estimate ($t = 1.70$) indicates that when comparing more-versus-less trade-exposed districts, the more-exposed district is predicted to experience a 22.6% (46.1×0.49) larger increase in contributions by right-leaning donors, which is equivalent to a 0.22 standard-deviation change in third-tercile contributions across districts over 2002 to 2010. These results provide our first evidence that greater trade exposure heightens polarization in campaign finance by increasing contributions among more-partisan donors on the left and right relative to contributions by moderate donors in the ideological center.

4.3.2 Extended Results for 2002-2016

We next consider results for a longer time horizon. Using the specification with full controls in column 5 of Table 4, we estimate regressions for time periods beginning in 2002 and ending in each congressional election year from 2004 to 2016. The dependent variables are those in panels B to D of Table 4, which represent changes in campaign contributions by 2002 terciles of donor CF scores. The trade shock variable in these regressions is always the growth of Chinese import competition from 2002 to 2010, so that the results up to 2010 are informative about the timing of the changes in campaign contributions that Table 4 reported for the 2002-2010 period, while the results for subsequent years indicate the persistence of these effects. During the early periods, 2002-2004 and 2002-2006, the full impact of the 2002-2010 China trade shock is yet to be felt; for the 2002-2008 and 2002-2010 periods forward, China's reform-driven export boom is largely complete. Up to 2010, congressional districts are stable and defined based on the 2000 Census. The periods ending in 2012, 2014, and 2016 include elections based on congressional districts whose boundaries reflect the redrawing of districts after the 2010 Census. Because of the need to match county-congressional-district cells to commuting zones across periods of redistricting, as discussed in Section 3.3, outcomes for these later time periods may be measured with more noise relative to outcomes for earlier time periods. In constructing changes in outcomes across redistricting periods, we therefore discard the two-year change in districts in which redistricting occurs, as shown in equation (3), such that results for 2002-2012 are very similar to those for 2002-2010, because a large majority of districts (but not all districts) changed their boundaries between 2010 and 2012.

Figure 4: Exposure to Chinese Import Competition and Campaign Contributions, 2002-2004/2016. Dependent Variable: Relative Change in Contributions (in log points)



Notes: Figure reports estimates of equation (6) for the relationship between changes in China trade exposure between 2002 and 2010 and 100× log changes in campaign contributions within ideology terciles (based on 1992 contributions, as per Figure (2)) within county-district cells across designated year pairs. Each bar represents a coefficient from a separate regression while whiskers indicate 90% confidence intervals. All regressions include the full vector of control variables from column 5 of Table 4. Observations are weighted by a county-district cell's share in the total year-2000 voting age population of a district, so that each district has a total weight of one. Standard errors are two-way clustered on counties and congressional districts. Full regression results are reported in Appendix Table A7.

Figure 4 presents the point estimates and 90% confidence intervals for these regression estimates. Each bar represents a coefficient from a separate regression while whiskers indicate 90% confidence intervals based on standard errors that are clustered both at the level of CZs and congressional districts.[47] Consider first the impact of trade exposure for the middle tercile of centrist donors. In all periods except the first, 2002-2004, the impact is positive, but it is always small and imprecisely estimated. Congressional districts more exposed to import competition see no differential increase in campaign contributions from moderate donors at any time horizon. Consider next the impact of trade exposure on contributions by first-tercile liberal donors. These impacts are positive and

[47]Full regression details appear in Appendix Table A7.

precisely estimated in each time period.[48] They are small in the first two time periods, roughly double in magnitude value in the middle three time periods, 2002-2008, 2002-2010, and 2002-2012, and increase further in the final two time periods. When examining the long-period change, 2002-2016, the coefficient estimate of 111.8 ($t = 2.59$) indicates that when comparing more-versus-less trade-exposed congressional districts, the more-exposed district would have approximately 54.8% (111.8×0.49) higher campaign contributions by liberal first-tercile donors, which is equivalent to a 0.37 standard-deviation change in first-tercile contributions across districts over 2002 to 2016.

As with the 2002-2010 results, the impacts of trade exposure on campaign contributions by right-leaning donors are qualitatively similar to those for left-leaning donors when we expand the time horizon under analysis. Impacts of trade exposure on increased contributions by conservative donors are small and imprecise in the first two time periods. Coefficient estimates increase substantially in magnitude and become significant in the 2002-2008 period, and remain comparable in subsequent periods apart from a dip in 2014.[49] When examining the coefficient estimate for the full-period change, 2002-2016, the now less-precise coefficient estimate of 52.6 ($t = 1.39$) indicates that when comparing more-versus-less trade-exposed districts, the more-exposed district would have approximately 25.8% (52.6×0.49) higher campaign contributions by left-leaning donors, which is equivalent to a 0.17 standard-deviation change in third-tercile contributions across districts over 2002 to 2016. Overall, these results suggest that greater trade exposure induces a polarization in campaign contributions in the 2000s that is largely maintained through 2016. Contributions from liberal and conservative donors, but not from moderate donors, differentially expand in more-trade-exposed districts.

In Appendix Figure A4, we revisit the results in Figure 4 by estimating regressions in which we split counties according to whether or not a majority of their voting-age residents were non-Hispanic whites according to Census 2000 enumeration data. The lion's share of U.S. county-district cells had a majority non-Hispanic white population in that year: $3,491$ of $3,772$ cells, corresponding to 370 of the 432 electoral districts (85.6%) that are used in our analysis.[50] This demographic split is, not surprisingly, correlated with the political affiliation of elected representatives: 58.5% of the population in majority white counties was represented by a Republican in 2002; conversely, 76.8% of the population in minority-dominated areas was represented by a Democrat in 2002.

For districts with a majority non-Hispanic white population in panel (a), the polarization results in Figure 4 are preserved. Greater exposure to import competition leads to larger increase in

[48] The corresponding t-values for the change in import competition are 1.74 for 2002-2004, 1.85 for 2002-2006, 2.88 for 2002-2008, 2.29 for 2002-2010, 2.30 for 2002-2012, 2.66 for 2002-2014, and 2.59 for 2002-2016.

[49] The corresponding t-values for the change in import competition are 0.77 for 2002-2004, 0.48 for 2002-2006, 2.30 for 2002-2008, 1.70 for 2002-2010, 1.71 for 2002-2012, -0.20 for 2002-2014, and 1.39 for 2002-2016.

[50] Our sample comprises $3,108$ counties, of which $2,924$ are majority-white. Minority-dominated counties are more populous on average, so that the reported fraction of minority-dominated districts is larger than the fraction of minority-dominated counties. Majority- and minority-dominated areas have an average trade shock value of 0.71.

contributions from liberal and conservative donors, but not from moderate ones. For liberal donors, impacts are precisely estimated for each end year from 2008 onward; for conservative donors, impacts are statistically significant for end years 2008, 2010, and 2012. When considering districts with majority-minority populations in panel (b), a materially distinct pattern emerges. There is a positive and significant impact of trade exposure on contributions by liberal donors, which is precisely estimated for all end years from 2006 forward. By contrast, impacts on contributions by moderate and conservative donors are small and imprecisely estimated in all years. As with news viewership in the Nielsen data and political beliefs in the Pew data, trade shocks appear to elicit different political responses among non-Hispanic whites than among other racial and ethnic groups.

Together, our results on political expression suggest that localized economic shocks stemming from rising trade pressure in the 2000s increased the relative demand for conservative media content, support for conservative viewpoints, and campaign contributions by more ideologically extreme donors. In our subsequent analysis, we examine changes in the composition of elected representatives in order to explore how these changes in political sentiment affected electoral outcomes.

5 Impact of Trade Shocks on Congressional Election Outcomes

In shifting our focus to the outcomes of congressional elections, we first consider the impact of import exposure on the standard election measures of voter turnout and party vote shares, which in turn allows us to square our results with current literature. We then examine how trade shocks have affected the composition of election winners, measured by party affiliation and ideological orientation. To analyze these outcomes, we continue to use regression specification (6) above.

5.1 Campaign Competitiveness, Party Vote Shares, and Party Win Percentages

We initiate our study of congressional elections by considering how rising exposure to import competition affects the number of registered voters who cast ballots and the share of votes cast captured by the GOP. In column 1 of Table 5, the dependent variable is the change in fraction of registered voters who turn out to vote in the general congressional election, where outcomes are for the 2002 to 2010 period. In all regressions we include the full set of controls for initial economic conditions, political conditions, and demographic characteristics, matching the specification in column 5 of Table 4. Voter turnout is higher in congressional districts subject to larger increases in trade exposure in their corresponding CZs. The coefficient estimate of 5.27 ($t = 2.72$) implies that when comparing districts at the 75^{th} versus 25^{th} percentiles of trade exposure, the more exposed district would have a 2.6 percentage-point (5.27×0.49) larger increase in voter turnout, relative to mean turnout in

2002 of 46.7% and a mean 2002-2010 change in turnout of 3.3 percentage points.[51] These results accord with the findings of Table 4 suggesting that rising trade exposure increases the intensity of campaigns, as indicated by increases in campaign contributions and higher voter participation.

Table 5: Exposure to Chinese Import Competition and Electoral Results, 2002-2010. Dependent Variables: Change in Turnout among Registered Voters, Change in Republican Two-Party Vote Share, or Change in Republican Win Probability (in % pts).

	Turnout in % of Reg. Voters	*Two-Party Republican Vote Share by District Sample*				Prob Republican Elected
		All Districts	Solid Democrat	Compe-titive	Solid Republican	
	(1)	(2)	(3)	(4)	(5)	(6)
Δ CZ Import Penetration	5.27	-1.08	-0.95	6.10	-6.24	24.08
	(1.94)	(5.98)	(1.80)	(4.93)	(3.93)	(12.07)

Notes: $N = 3,772$ county-district cells, except $N = 2,772$ in column 1. Turnout among registered voters is measured at the county level and excludes counties in districts with uncontested elections in 2002 or 2010, as well as district that were redistricted in 2004 or 2006. The Republican two-party vote share is the ratio of Republican votes to the sum of Democratic and Republican votes. Column 3 indicates the change in vote share in the 129 districts where the Democratic party maintained a two-party vote share of >55% in every election from 2002 to 2010. Its sets the outcome variable to zero for all districts where this condition was not met. Column 5 correspondingly indicates the change in vote share in the 124 districts where the Republican party maintained a two-party vote share of >55% in every election from 2002 to 2010, while column 4 comprises the 179 remaining districts. All regressions include the full vector of control variables from column 5 of Table 4. Observations are weighted by a county-district cell's share in the total year-2000 voting age population of a district, so that each district has a total weight of one. Standard errors are two-way clustered on counties and congressional districts.

Does this trade-induced increase in electoral competitiveness tend to favor one political party or the other? In column 2 of Table 5, the outcome is the change in the GOP share of two-party vote in the general election.[52] We see that greater trade exposure has a modest negative impact on the change in the Republican vote share, where the coefficient estimate is small (-1.1 percentage points per percentage-point change in trade exposure) and highly imprecise. These results are broadly in line with Che et al. (2016), who document vote share gains for Democrats in counties with greater exposure to Chinese import competition from 2002 to 2010. Our analysis at the county-district level allows us to identify the districts that accounted for these vote share gains of the Democratic party. In columns 3 to 5, we split districts into "safe districts" that were consistently held by the same party with vote shares of 55% or higher in each election from 2002 to 2010, and "competitive" districts where neither party consistently attained at least 55% of the vote. This classification yields 129 "safe Democratic" districts, 124 "safe Republican" districts, and 179 "competitive" districts. Columns 3

[51] This finding is consistent with the classic quiescence hypothesis in political science (Edelman, 1971), which views low voter turnout as indicative of voter satisfaction, and conversely, implies that rising voter dissatisfaction will spur turnout. The adverse impacts of trade shocks could plausibly be one source of such dissatisfaction.

[52] The Republican two-party vote share is the share of Republican votes among the total of Republican and Democratic votes. In our analysis, we count as a Democrat the lone independent member of Congress, Bernie Sanders of Vermont. On two occasions, Sanders later sought the presidential nomination of the Democratic Party.

to 5 respectively interact the vote share outcome of column 2 with dummies for safe Democratic, competitive, and safe Republican districts, such that the regression coefficients across these columns add up to the total effect in column 2. The column 3 and 5 results indicate that the Democratic party increased its vote share in trade-exposed districts that remained under safe control of the incumbent party. Column 4 however shows that the Republican party gained in trade-exposed districts where both parties were competitive. While none of the columns 2 to 5 results is precisely estimated, the coefficient pattern suggests that modest overall vote share gain for the Democratic party masks gains for the Republican party in the subset of districts that were not firmly controlled by one party. Indeed, column 6 of Table 5 shows that districts more exposed to import competition became more likely to elect a GOP legislator, where this impact is significant at the 5% level ($t = 2.00$). Comparing more-versus-less trade-exposed congressional districts, the more exposed district would have a substantial 11.8 percentage-point (24.08×0.49) larger increase in the probability of electing a Republican.

But this effect did not arise immediately. Consistent with findings in Feigenbaum and Hall (2015), Figure 5 shows that import competition had little impact on the party balance in congress through 2008.[53] Electoral gains for the GOP emerge in the 2010 election that brought many Tea Party Republicans into congress, and these gains persist thereafter. In Figure 5a, over the 2002-2012, 2002-2014, and 2002-2016 time periods, greater trade exposure had positive and precisely estimated impacts on the change in probability that a Republican candidate wins the election, where the coefficient magnitudes for these later-ending periods are either equal to or larger than those for 2002-2010: 24.1 ($t = 1.99$) for 2002-2012, 26.1 ($t = 2.14$) for 2002-2014, and 27.1 ($t = 2.19$) for 2002-2016. Over these same time horizons, Figure 5b shows that the trade-exposure impact on the GOP vote share is small, negative, and highly imprecisely estimated, as is the case for the 2002 to 2010 period.

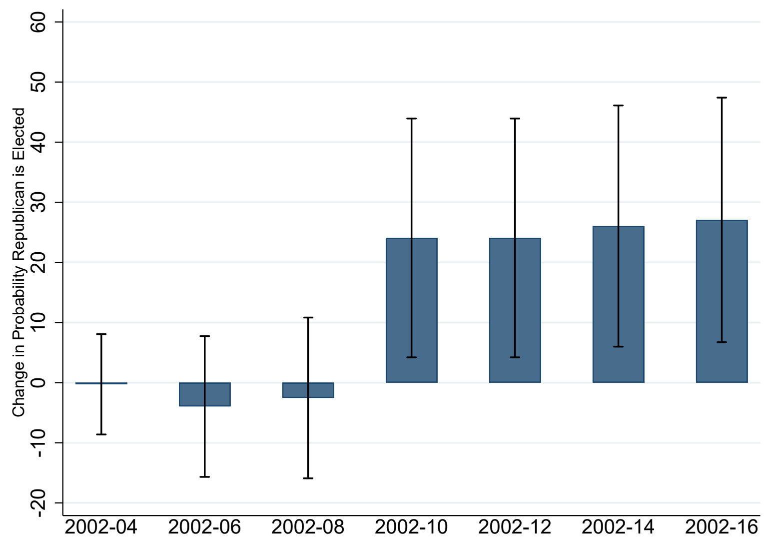
How do we reconcile trade shocks weakly lowering the GOP vote shares in Figure 5b while raising GOP win probabilities in Figure 5a? Column 4 of Table 5 offers suggestive evidence that the Republican party may have improved its electoral results in the competitive districts where a few additional percentage points of the vote share matter could prove decisive for victory. In supplemental analysis, we indeed find that greater trade exposure has a positive impact on the likelihood that Republicans win an election with a narrow GOP vote margin of 1 - 20% , while reducing the likelihood of a dominant GOP victory with a margin exceeding 20%. These outcomes

[53]See Appendix Table A8 for regression details. Feigenbaum and Hall (2015) report that through the Congress elected in 2008, trade-exposed districts had no greater likelihood of a contested primary, no greater likelihood of a loss by the incumbent candidate, and no lower vote share for the incumbent candidate. While they did not study the electoral success of the Democratic and Republican parties in those districts, they observe that elected legislators did not systematically adjust their voting behavior in Congress apart from greater support for protectionists bills.

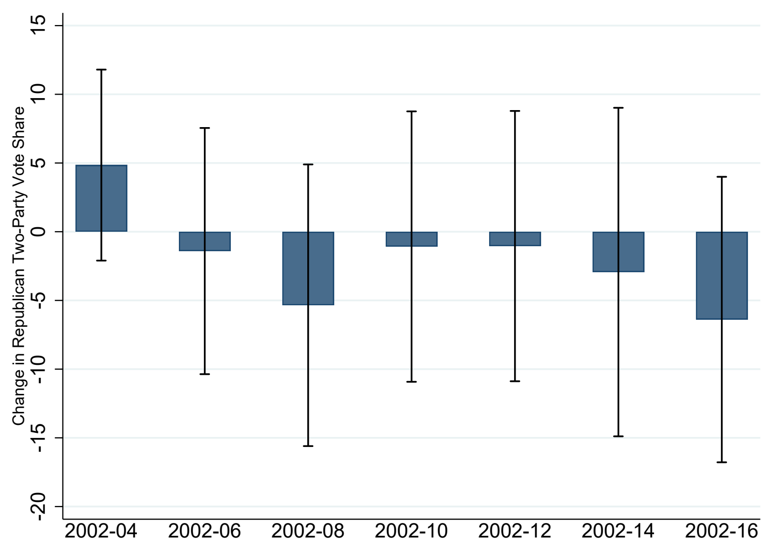
accord with the results in Tables 4 and 5 on how trade shocks increase campaign intensity.

Figure 5: Exposure to Chinese Import Competition and Electoral Results, 2002-2010. Dependent Variables: Change in Republican Two-Party Vote Share, or Change in Republican Win Probability (in % pts)

a. Change in Probability Republican is Elected



b. Change in Republican Two-Party Vote Share

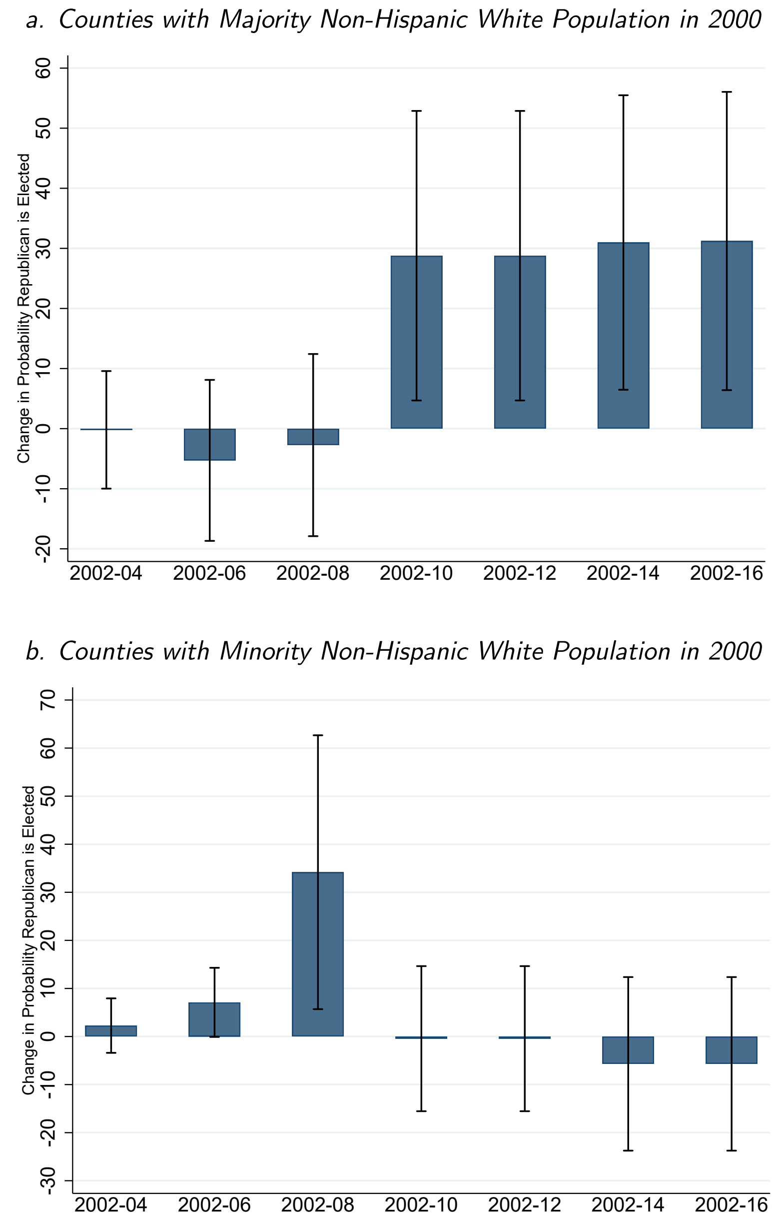


Notes: Estimates of equation (6) for the relationship between the change in China trade exposure 2002-2010 and (panel A) the change in the probability that a Republican is elected, and (panel B) the change in the Republican two-party vote share. Each bar represents a coefficient from a separate regression while whiskers indicate 90% confidence intervals. All regressions include the full vector of control variables from column 5 of Table 4. Observations are weighted by a county-district cell's share in the total year-2000 voting age population of a district, so that each district has a total weight of one. Standard errors are two-way clustered on counties and congressional districts. See Appendix Table A8 for full regression results.

More competitive elections could in turn be the consequence of parties running more centrist candidates against each other, who, because they compete for similar groups of voters, realize narrower electoral margins. However, this explanation runs counter to the perception that the GOP has shifted demonstrably to the right, as supported by the increase in CF scores in Figure 3, the emergence of the Tea Party movement in 2010 (Madestam et al., 2013; Parker and Barreto, 2014), and the much-discussed ongoing decline of the party's moderate wing (Kabaservice, 2012). Alternatively, the model of Glaeser et al. (2005) suggests that greater competitiveness of elections could be the result of more extreme candidates exploiting wedge issues during campaigns to catalyze turnout and financial contributions among their core supporters. In the framework of Grossman and Helpman (2018), inflaming wedge issues is the rough equivalent of strengthening group identity, such as Tea Party acolytes declaring their opposition to immigration, affirmative action, and social protections for disadvantaged groups. In either case, extreme candidates—by virtue of their extremism—may be more likely to win elections narrowly when they happen to come out on top.

If GOP success in more trade-exposed congressional districts is in part a consequence of candidates taking more ideologically extreme or polarizing positions, we would expect the impacts on GOP vote share to differ according to districts' racial and ethnic composition. In Figure 6 we repeat the analysis in Figure 5a, now splitting counties according to whether or not a majority of their voting-age residents were non-Hispanic whites according to the 2000 Census. Regression details appear in Appendix Table A9. In majority non-Hispanic-white congressional districts, shown in Figure 6a, the results in Figure 5a are preserved. For all end years from 2010 onward, districts exposed to greater import competition had a larger increase in the likelihood of a GOP congressional election victory. The coefficient rises from 28.8 ($t = 1.97$) for 2002-2010 to 31.0 ($t = 2.08$) for 2002-2014 and to 31.3 ($t = 2.07$) for 2002-2016. Taking the coefficient estimate for the full 2002-2016 time period, when comparing more-versus-less trade-exposed majority-white congressional districts, the more-exposed district would have a sizable 15.3 percentage-point (31.3×0.49) larger increase in the GOP win probability. By contrast, in majority-minority congressional districts, no such impacts obtain over these later time periods (although they do obtain for 2002-2008). From the 2002-2010 period onward, greater trade exposure leads to sizable increases in the likelihood of GOP victory in majority white non-Hispanic congressional districts but not in majority-minority districts.

Figure 6: Exposure to Chinese Import Competition and Electoral Results, 2002-2010. Dependent Variables: Change in Republican Win Probability (in % pts)



Notes: Estimates of equation (6) for the relationship between the change in China trade exposure 2002-2010 and the change in the probability that a Republican is elected in (panel A) counties with a majority non-Hispanic white population in 2000, and (panel B) counties with a minority non-Hispanic white population in 2000. Each bar represents a coefficient from a separate regression while whiskers indicate 90% confidence intervals. All regressions include the full vector of control variables from column 5 of Table 4. Observations are weighted by a county-district cell's share in the total year-2000 voting age population of a district, so that each district has a total weight of one. Standard errors are two-way clustered on counties and congressional districts. See Appendix Table A9 for full regression results.

It may seem puzzling that greater trade exposure helps GOP candidates in majority white congressional districts but matters little for GOP win prospects in majority-minority districts. One might instead expect these results to be symmetric, with trade exposure causing the GOP to lose ground in majority non-white areas. To explore this issue, we assess which types of candidates—in terms of their ideological orientation—have become more likely to win elections.

5.2 Ideology of Congressional Election Winners

We now characterize the impact of rising trade exposure on congressional elections according to the political party *and* ideological orientation of those elected, where the ideology of winners is measured by the contributed-weighted-average CF score of the donors to their election campaign.[54] We define "moderate Democrats" and "moderate Republicans" as legislators whose contribution-weighted-average CF score would place them in the more centrist half of their party's legislators in 2002. By contrast, "liberal Democrats" have an ideology score below the median of their party in 2002, while "conservative Republicans" have an ideology score above the 2002 party median.[55]

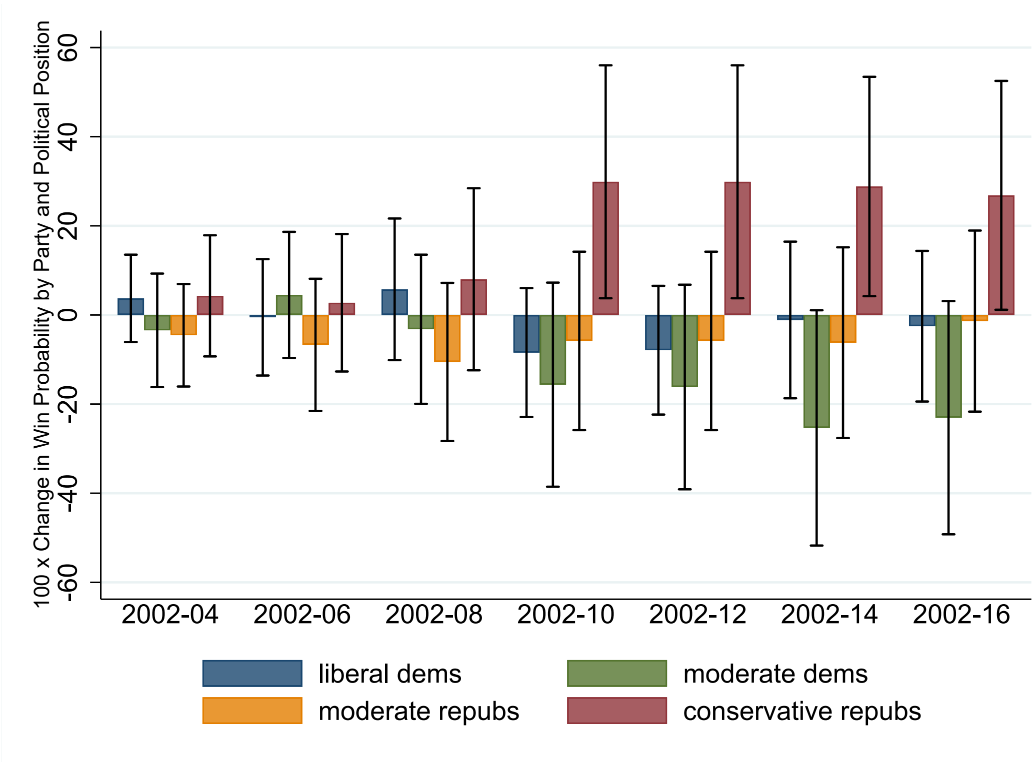
Figure 7 displays estimates of the impact of trade exposure on the probability that candidates from equal ideological partitions are elected to the House of Representatives for time periods ranging from 2002-2004 to 2002-2016. The specification is that in equation (6), with full controls for initial economic conditions, political conditions, and demographic characteristics. Each bar represents a separate regression in which the dependent variable is the change in the likelihood that a given type of candidate wins the election. Because the four categories—liberal Democrat, moderate Democrat, moderate Republican, conservative Republican—are exhaustive and mutually exclusive, the heights of the four bars sum to zero within each time period, except for small deviations caused by the redistricting adjustments of section 3.3. Regression details appear in Appendix Table A10.

Consider first electoral outcomes for conservative Republicans. In all time horizons from 2002-2010 onward, districts subject to greater import competition became substantially more likely to elect a GOP conservative. This effect is significant at the 10% level in 2002-2010 ($t = 1.88$), 2002-2012 ($t = 1.88$), 2002-2014 (t-value $t = 1.95$), and 2002-2016 ($t = 1.72$) for which the coefficient magnitude falls in the narrow range of 29.9 to 26.8. Using results for the 2002-2016 period, when comparing a more-versus-less trade-exposed district the former would have a 13.2 percentage-point (26.8×0.49) higher likelihood of electing a conservative Republican.

[54] Implicitly, we assume that candidates are ideologically and (or) legislatively aligned with their donors.

[55] DIME records campaign contributions for 96.2% of all election winners from 2002 to 2016. In the rare cases where a winner received no contributions during an electoral cycle, we impute the winner's ideology value using the next previous or subsequent election in which the same candidate obtained contributions, or absent that, from other same-party election winners of the same district or state. These imputations do not materially affect our results

Figure 7: Exposure to Chinese Import Competition and Change in Ideological Position of Election Winner, 2002-2016. Dependent Variable: 100 x Change in Indicators for Election of Politician by Party and Political Position



Notes: Estimates of equation (6), with full controls for initial economic conditions, political conditions, and demographic characteristics. Each bar represents a separate regression in which the dependent variable is the change in the likelihood that a given type of candidate wins the election. The four categories—liberal Democrat, moderate Democrat, moderate Republican, conservative Republican—are exhaustive and mutually exclusive. Whiskers indicate 90% confidence intervals. Observations are weighted by a county-district cell's share in the total year-2000 voting age population of a district, so that each district has a total weight of one. Standard errors are two-way clustered on counties and congressional districts. Regression details appear in Appendix Table A10.

These trade-shock-induced improvements in electoral prospects for GOP conservatives must come at the expense of other candidate types. For 2002-2010 onward, the impacts of greater import competition on the election probabilities of each of the other three candidate types is negative, though none is precisely estimated. Figure 7 indicates that trade shocks do not cause a monotone shift towards the political right. Instead, moderate politicians experience the largest decline in election probability in each period, with moderate Democrats accounting for the bulk of the losses from conservative GOP gains from 2010 onwards. For 2002-2010, the trade-shock-induced decline in election probability for moderate Democrats is 52.3% ($-15.6/29.9$) of the gain for GOP conservatives, a fraction that reaches 85.9% ($-23.1/26.8$) for 2002-2016. These results align with the noted demise of moderate congressional legislators in recent decades (e.g., Layman et al., 2006).[56] Figure 5a above showed

[56]We alternatively classified the ideology of elected legislators using DW-Nominate scores that are based on roll call

that over 2002 to 2016, greater exposure to import competition enhanced the electoral prospects of Republicans seeking congressional office. Figure 7 makes clear that the primary beneficiaries of these gains were conservative candidates. These results present something of a puzzle. Given that trade shocks induced a polarization of campaign contributions (section 4)—with donations rising differentially both among left-leaning and right-leaning contributors—it is not obvious why the electoral winners are lopsidedly drawn from the right of the ideological spectrum. To gain additional leverage, we focus on the districts that drive the polarization of campaign contributions. Figure A4b established that the left-leaning pole of the polarized response of campaign contributions to trade exposure was driven by majority-minority congressional districts. If more liberal candidates did in fact benefit electorally from these polarized contributions, we would expect to see these effects in majority-minority districts.

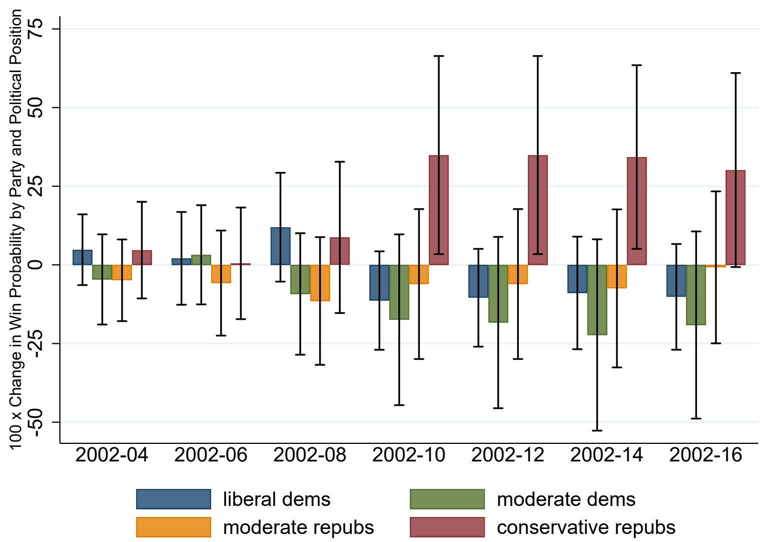
Figure 8 examines the effects of trade shocks on electoral outcomes, where county-district cells are split according to whether or not a majority of a county's voting-age residents in 2000 were non-Hispanic whites. The upper panel of Figure 8 shows that in counties with majority non-Hispanic White populations, trade exposure catalyzed movements towards conservative Republicans between 2002 and 2010 and in all later periods. The impacts are marginally significant for 2002-2010, 2002-2012, and 2002-2014 (t-values $t = 1.8$, 1.82, and 1.93, respectively) and slightly less precise for 2002-2016 ($t = 1.61$). Scaling by the interquartile range of trade exposure, the 2002-2016 point estimate implies that a conservative Republican became 14.8 percentage points (30.2×0.49) more likely to take office in a more-versus-less-trade-exposed congressional district, with these gains coming at the expense of moderate Democrats (-9.4 points) and liberal Democrats (-5.0 points).

Focusing attention on the subset of counties where less than half of the voting-age population is non-Hispanic white (Figure 8b), we find a largely complementary pattern: liberal Democrats made strong gains in these locations in the probability of taking office for 2002-2010 and later periods. For the full sample period of 2002-2016, the standardized effect size is a 21.5 percentage-point ($t = 2.74$) increase in the probability that a liberal Democrat wins office. These gains came almost entirely at the expense of moderate Democrats, whose standardized loss in win probability is 21.1 percentage points ($t = 3.25$).

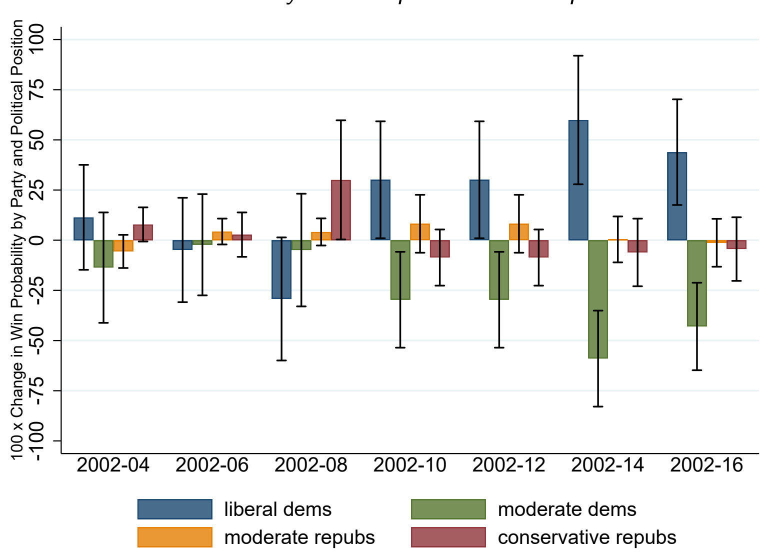
votes in Congress, which yields similar outcomes. In unreported results, we find that a unit trade shock raises the election probability of conservative Republicans increases by 36.8 percentage points over the 2002 to 2010 period, with the demise of moderate Democrats accounting for 66% of that effect, and a smaller decline of moderate Republicans accounting for the remainder while the election probability of liberal Democrats is unchanged.

Figure 8: Exposure to Chinese Import Competition and Change in Ideological Position of Election Winner, 2002-2004/2016. Heterogeneity by Initial Local Racial Composition. Dependent Variable: 100 x Change in Indicators for Election of Politician by Party and Political Position.

a. Counties with Majority Non-Hispanic White Population in 2000



b. Counties with Minority Non-Hispanic White Population in 2000



Notes: Estimates of equation (6), with full controls for initial economic conditions, political conditions, and demographic characteristics. Each bar represents a separate regression in which the dependent variable is the change in the likelihood that a given type of candidate wins the election. The four categories—liberal Democrat, moderate Democrat, moderate Republican, conservative Republican—are exhaustive and mutually exclusive. Whiskers indicate 90% confidence intervals. Panel A presents estimates for counties with majority non-Hispanic white populations in 2000, while panel B presents estimates for the complementary set. Observations are weighted by a county-district cell's share in the total year-2000 voting age population of a district, so that each district has a total weight of one. Standard errors are two-way clustered on counties and congressional districts. Regression details appear in Appendix Table A11.

These results suggest a possible resolution to the apparent paradox in Figure 6, in which greater trade exposure helps GOP candidates in majority white districts but does not hurt them in majority non-white districts. In both sets of districts, candidates advantaged by adverse trade shocks pull support from moderate Democrats. In majority white districts, the GOP conservatives who gain at moderates' expense pull support across party lines and thereby increase the likelihood of a GOP win. In majority non-white districts, liberal Democrats pull support from moderates within their own party, such that there is no net impact in which party takes congressional seats. Combining these two sets of impacts, the GOP benefits on net from greater import competition.

These findings, in which congressional districts diverge in their political responses to trade shocks based on their initial racial composition, are consistent with an emerging political economy literature documenting a connection between voter responses to trade shocks that cleaves according to economic, racial, or ethnic identity (Grossman and Helpman, 2018; Gennaioli and Tabellini, 2019). Our results go beyond these regularities to show that economic shocks related to trade have a causal impact on political partisanship that separates according to race. In combination with our earlier findings that trade shocks raise both voter turnout and individual-level campaign contributions, these results are consonant with a recent literature that emphasize the role of identity in shaping both partisan affiliation and policy outcomes. In the model of Glaeser et al. (2005), opportunistic politicians deploy strategic extremism to spur participation among core supporters, which amplifies partisan identification and raises turnout.[57] Models by Grossman and Helpman (2018) and Gennaioli and Tabellini (2019) provide clear rationales as to why adverse economic shocks should catalyze this process—fostering divergent shifts in group identification and potentially leading to abrupt shifts in trade policy. Our results appear strongly supportive of this line of applied theory.

6 Impact of Trade Shocks on Presidential Elections

Our findings indicate that trade exposure contributed to a net rightward shift in the ideology of elected legislators and a gain in House seats for the Republican party despite not having a sizable impact on party vote shares. Because each Congressional district chooses among a disparate set of candidates, votes cast for a candidate from the same political party in different districts are not necessarily votes cast in favor of a legislator with the same ideological position relative to local alternatives. Presidential elections by contrast provide a setting in which all localities simultaneously choose among the same candidates. Since the time-varying geographic structure of congressional districts is not relevant in this context, we can analyze county-level changes of party vote shares in

[57] The Glaeser et al. (2005) model does not, however, feature a specific role for economic (or other) shocks to intensify strategic extremism.

presidential elections for a longer time period that is not constrained by re-districting.

In Table 6, we estimate the impact of trade exposure on the change in the county-level GOP vote share between the 2000 and 2008 and the 2000 and 2016 presidential elections. These highly competitive elections bracket the time period of our analysis of congressional elections. The three years considered—2000, 2008, and 2016—correspond to elections in which a two-term incumbent (Bill Clinton, George W. Bush, Barack Obama, respectively) was stepping down from office, and thus represent common positions in the political cycle. Our measure of trade exposure is that used in equation (6), now defined over the period 2000 to 2008, while the instrumentation strategy continues to follow that in section 3.1.[58]

Table 6: Exposure to Chinese Import Competition and Presidential Elections, 2000-2008 and 2000-2016, 2SLS Estimates. Dependent Variable: Change in Percentage of Two-Party Vote Obtained by Republican Candidate, 2008 (McCain) or 2016 (Trump) vs 2000 (Bush)

	(1)	(2)	(3)	(4)	(5)
	A. Δ Net Republican Vote Share 2000-2008				
Δ CZ Import Penetration 2000-2008	1.54	5.60	2.38	1.75	1.59
	(0.73)	(1.41)	(1.24)	(0.86)	(0.85)
	B. Δ Net Republican Vote Share 2000-2016				
Δ CZ Import Penetration 2000-2008	3.86	3.98	1.72	1.99	1.71
	(1.48)	(1.69)	(1.71)	(0.97)	(0.90)
Estimation Method	2SLS	2SLS	2SLS	2SLS	2SLS
F-statistic First Stage	63.7	50.2	46.4	48.1	48.0
2000 Ind/Occ Controls		yes	yes	yes	yes
Census Division Dummies			yes	yes	yes
2000 Demography Controls				yes	yes
1992/1996 Election Controls					yes

Notes: N=3,107 counties, excluding Alaska and Hawaii. The mean change in net Republican vote share is -3.50 (s.d. 5.69) between 2000 and 2008 and is -0.74 (s.d. 9.95) between 2000 and 2016. Industry and occupation controls in column 2 are measured at the CZ level and comprise the fraction of CZ employment in the manufacturing sector and the Autor and Dorn (2013b) routine share and offshorability index of a CZ's occupations. Census division dummies in column 3 allow for different time trends across the nine geographical Census divisions. Demographic controls in column 5 comprise the percentage of a county's population in nine age and four racial groups, as well as the population shares that are female, college-educated, foreign-born, and Hispanic. Election controls in column 5 comprise the Republican two-party vote share in the presidential elections of 1992 and 1996, measured at the county level. Observations are weighted by counties' total votes in the 2000 presidential election, and standard errors are clustered by CZ.

[58]The sequentially added control variables follow the specification used for congressional elections, except that we lag controls for electoral outcomes in presidential elections by an additional four years to avoid a mechanical correlation with the outcome variables.

The 2SLS estimates reported in panel A of Table 6 find a positive and marginally statistically significant impact ($t = 1.87$) of rising Chinese import competition on the share of votes going to the GOP presidential candidate between 2000 and 2008. The point estimate for the column 6 regression, which includes full controls, implies that the Republican two-party vote share rose by nearly a full percentage point for an interquartile range increase in import penetration ($1.59 \times 0.58 = 0.91$). Panel B indicates that the trade-induced shift in party vote share persisted after 2008. Counties that had been more exposed to import competition during the Chinese import boom continued to favor the Republican candidate in the 2016 election, where the impact with full controls in column 6 is larger in magnitude (1.71) and slightly more precisely estimated ($t = 1.90$) than for the 2000-2008 period.[59] These results on presidential elections corroborate our finding from congressional elections that greater trade exposure induces a net shift in favor of candidates on the right.

7 Concluding Remarks

The polarization of national politics has been one of the defining developments of American discourse of the last several decades. The coincidence of intensifying political partisanship and rising income inequality has led many to conjecture that economic changes are at least partly responsible for greater political divisiveness. Indeed, political actors have frequently suggested a connection between changes in the U.S. economy and the growing ideological divide in Congress. In the 2016 U.S. presidential campaign, candidates from both parties singled out China's rise as an international competitor as a principal cause of U.S. economic malaise. Yet, there is a paucity of evidence that substantiates a causal impact of specific economic shocks on political polarization.

Our contribution in this paper is to show that this vitriolic campaign rhetoric is indicative of underlying economic pressures that find voice in electoral contests. Growing import competition from China has contributed to a shift to the right in the media-viewing habits and political beliefs of U.S. adults as well as more competitive congressional elections, greater polarization in the ideological orientation of campaign contributors, and net gains in the number of conservative GOP representatives, which come largely at the cost of moderate Democrats. During the two most recent

[59] In a related research note (Autor et al., 2017), we also find a significant positive impact on the change in GOP presidential county vote shares over 2000 to 2016 for a CZ-level trade shock that extends from 2000 to 2014, the last year for which we have trade data (where most of the increase in Chinese import penetration occurred by 2008). We calculate that a 50.0% ceteris paribus reduction in the China trade shock between 2000-14 would have tipped the narrow Republican voter majority in the states of Pennsylvania, Wisconsin, and Michigan, leading to an Electoral College victory for Hillary Clinton, instead of a victory for Donald Trump. This notional exercise highlights the relevance of a trade-induced shift in party vote shares in presidential elections, which are more closely contested than most congressional elections. It however corresponds to a restrictive scenario where local exposure to the China shock affects the 2016 U.S. presidential general election exclusively through its effect on the local Republican two-party vote share. Our results above show that the China shock altered the ideological composition of the House prior to 2016, and those representatives' political activities may have subsequently contributed to the 2016 election outcome.

non-incumbent presidential elections, 2008 and 2016, trade shocks also appeared to differentially increase the vote share of the Republican candidate.

It may be unsurprising that negative impacts of trade on U.S. manufacturing have engendered an intense political response. Less expected is that the valence of this response depends non-monotonically on the initial racial composition of a congressional district. In majority-white districts, the political beneficiaries of these economic forces are Republicans, particularly from the far right, whereas majority-minority districts experience shifts to the left end of the spectrum. In both majority-white and majority-minority districts, these polarizing ideological shifts come primarily at the electoral expense of moderate Democrats, meaning that the net gains in seats accrues primarily to the Republican party. The paradox of converging popular beliefs about the source of economic challenges accompanied by diverging beliefs about appropriate political responses is consistent with theoretical models that connect economic adversity to in-group/out-group identification, as motivated in part by group-based resource competition or opportunistic use of political extremism.

What may distinguish trade in terms of its impact on political outcomes is that its disruptive effects are so concentrated demographically and geographically. The loss of manufacturing jobs has represented a major contraction in high-wage earning opportunities, especially for less-educated males (Autor et al., 2019). Further, whereas exposure to technological change in the labor market has affected both wealthy cities populated by white-collar professionals and factory towns populated by blue-collar workers, rising import penetration from low-wage countries disproportionately bears on local labor markets that historically specialized in labor-intensive manufacturing. The combination of these features enhances the salience of the labor-market impacts of trade and therefore their political resonance (Margalit, 2011). While it would be unwarranted to conclude that the China trade shock is the original or fundamental cause of three decades of growing U.S. political polarization, our analysis of the China trade shock highlights a nuance masked by aggregate data: the connection between economic and political polarization may arise not entirely from overarching secular changes in the U.S. economy that affect skill demands nationally, but also from shocks whose disruptive force falls heavily on an identifiable set of voters who in turn respond with concentrated vehemence at the polls.

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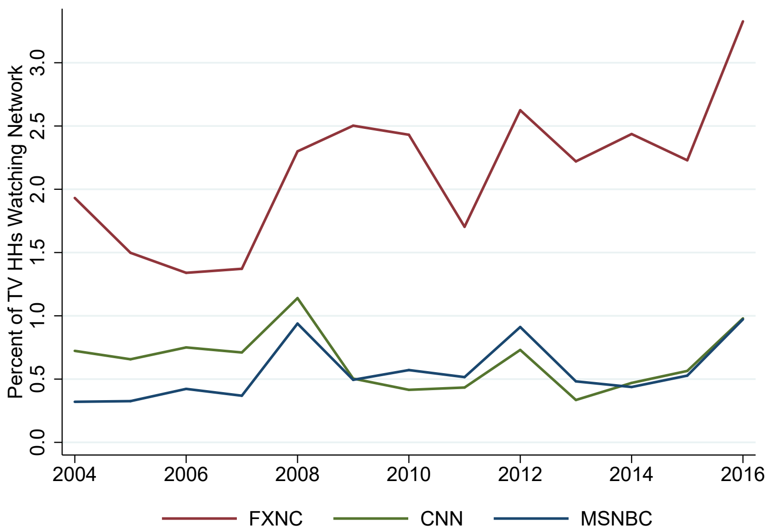
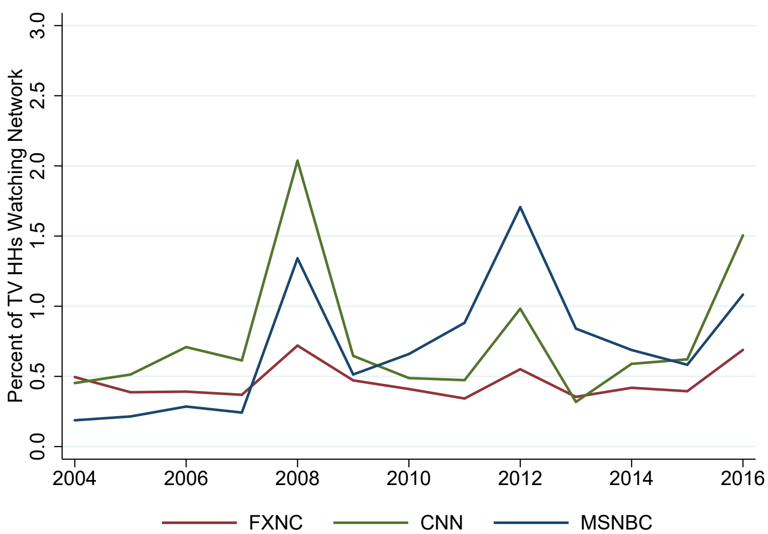
Appendix Figures and Tables

Figure A1: Nielsen Rating for Cable News Networks, by Race of Household Head, 2004 to 2016

a. Households Headed by Non-Hispanic Whites



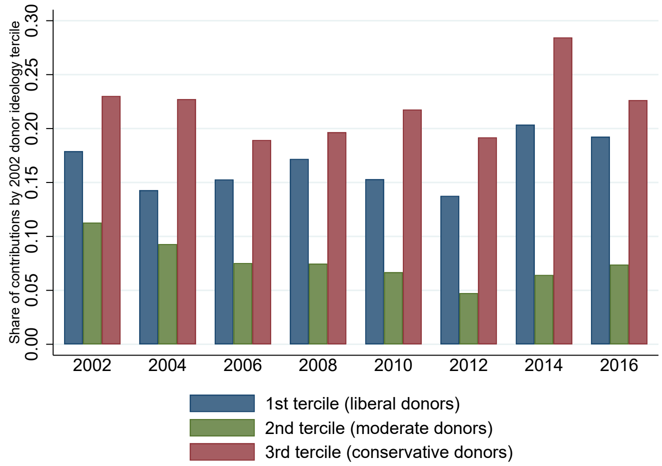
b. Households Headed by Hispanics, Non-Whites



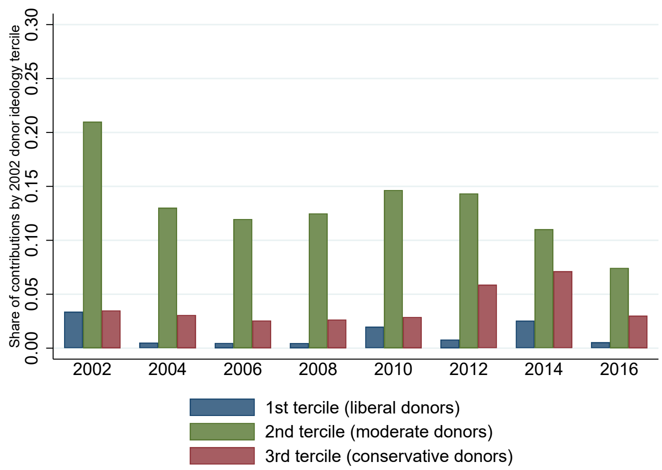
Notes: Nielsen ratings indicate the fraction of all TV-owning households that are tuned to a particular program at a particular time. Figure plots average ratings for the 5pm to 11pm time-slot, Monday through Friday, during the month of November for years 2004 through 2016. Sample size for each November estimate ranges from $99,000$ to $119,000$ households. Panels A and B present estimates for Nielsen households split according to the race and ethnicity of the household head.

Figure A2: Polarization in Campaign Finance Scores by Type of Campaign Donor

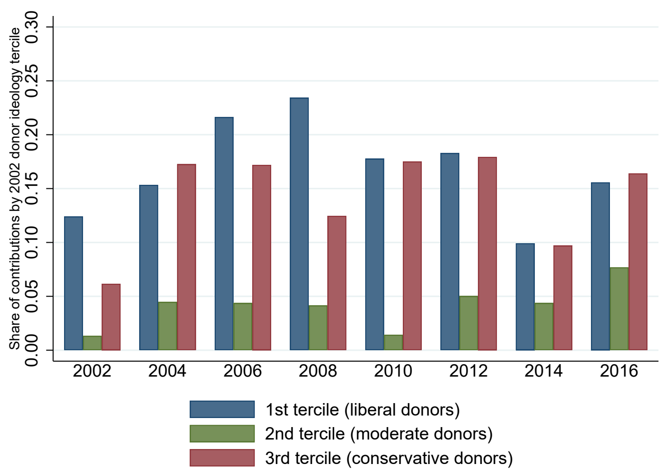
a. Campaign Contributions by Individuals



b. Campaign Contributions by Corporations

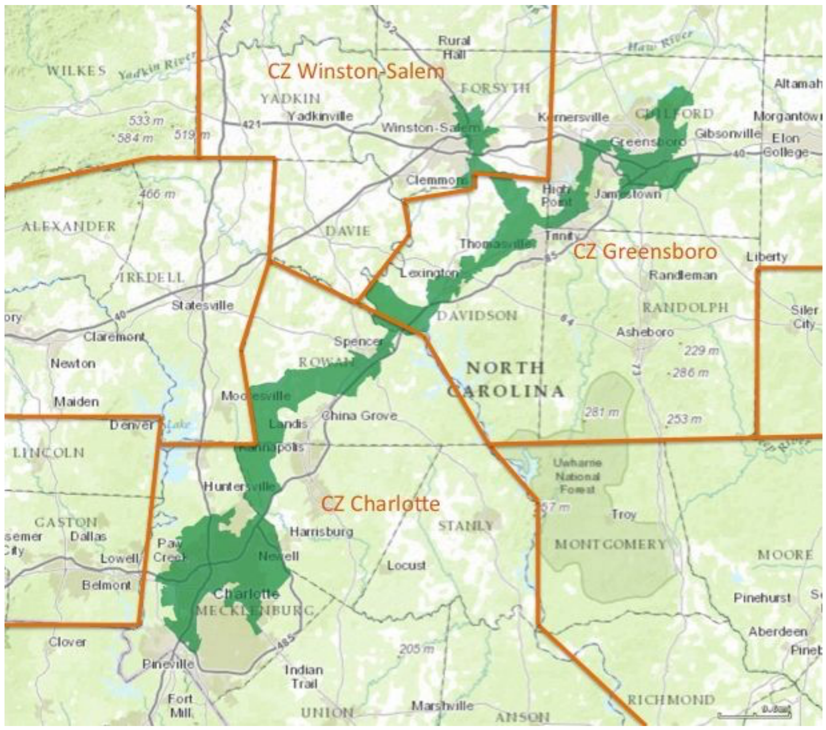


c. Campaign Contributions by Non-Corporate Organizations



Notes: Calculations based on Data Database on Ideology, Money in Politics, and Elections database (DIME; Bonica, 2013). Donor ideology is divided into ideology terciles based on campaign contributions in 2002 ranked by dollar-weighted CF scores. The first tercile comprises the most liberal donors, while the third tercile comprises the most conservative donors. The height of each bar in each reported year reflects the share of all contributions (in dollars) falling within each 2002 ideology tercile by donor types, which are individuals, corporations, and non-corporate organizations in panels A, B, and C respectively.

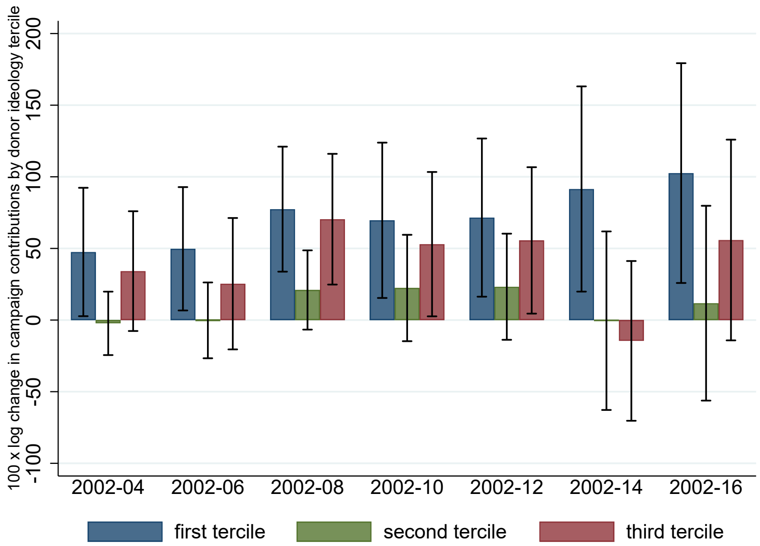
Figure A3: County-District Cells for the 12th Congressional District of North Carolina for the 111th Congress.



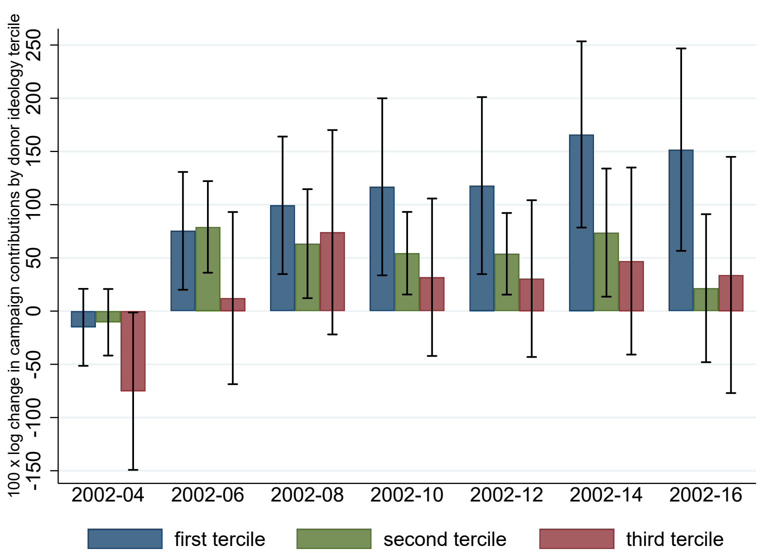
Notes: Figure depicts the geography of North Carolina Congressional District 12, which crosses three Commuting Zones as defined for the 111th Congress.

Figure A4: Exposure to Chinese Import Competition and Campaign Contributions, 2002-2004/2016. Dependent Variable: Relative Change in Contributions (in log points)

a. Counties with Majority Non-Hispanic White Population in 2000



b. Counties with Minority Non-Hispanic White Population in 2000



Notes: Figure reports estimates of equation (6) for the relationship between changes in China trade exposure between 2002 and 2010 and 100× log changes in campaign contributions within ideology terciles (based on 1992 contributions, as per Figure (2)) within county-district cells across designated year pairs. Each bar represents a coefficient from a separate regression while whiskers indicate 90% confidence intervals. Panel A presents estimates for counties with majority non-Hispanic white populations in 2000, while panel B presents estimates for the complementary set. All regressions include the full vector of control variables from column 5 of Table 4. Observations are weighted by a county-district cell's share in the total year-2000 voting age population of a district, so that each district has a total weight of one. Standard errors are two-way clustered on counties and congressional districts. Full regression results are reported in Appendix Table A7.

Table A1: Questions Comprising the Ten Item Pew Ideological Consistency Scale

	Conservative Position (1)	*Liberal Position* (2)
1	Government regulation of business usually does more harm than good	Government often does a better job than people give it credit for
2	Government is almost always wasteful and inefficient	Government regulation of business is necessary to protect the public interest
3	Poor people today have it easy because they can get government benefits without doing anything in return	Poor people have hard lives because government benefits don't go far enough to help them live decently
4	The government can't afford to do much more to help the needy	The government should do more to help needy Americans, even if it means going deeper into debt
5	Blacks who can't get ahead in this country are mostly responsible for their own condition	Racial discrimination is the main reason why many black people can't get ahead these days
6	Immigrants today are a burden on our country because they take our jobs, housing and health care	Immigrants today strengthen our country because of their hard work and talents
7	Most corporations make a fair and reasonable amount of profit	Good diplomacy is the best way to ensure peace
8	Stricter environmental laws and regulations cost too many jobs and hurt the economy	Business corporations make too much profit
9	The best way to ensure peace is through military strength	Stricter environmental laws and reulations are worth the cost
10	Homosexuality should be discouraged by society	Homosexuality should be accepted by society

Notes: Pew Ideological Consistency Scale, administered 1994 though present. Individual questions were recoded as “-1” for a liberal response, “+1” for a conservative response, “0” for other (don’t know/refused/volunteered) responses. Scores on the full scale range from -10 (liberal responses to all 10 questions) to +10 (conservative responses to all 10 questions). Documentation available at http://www.people-press.org/2014/06/12/appendix-a-the-ideological-consistency-scale/ (accessed 11/23/2017)

Table A2: Share of Population by Pew Ideology Score and by Race/Ethnicity

Year	Mean Score (1)	% Liberal (2)	% Moderate (3)	% Conservative (4)
		A. Non-Hispanic Whites		
2004	-0.63	30.9	46.9	22.2
2011	0.39	27.1	38.7	34.1
2014	0.02	30.9	36.0	33.1
2015	0.09	31.0	33.9	35.0
Δ2004-15	0.71	0.1	-13.0	12.9
		B. Hispanics and Non-Whites		
2004	-1.65	37.3	53.5	9.2
2011	-1.83	40.1	49.7	10.2
2014	-1.92	42.3	46.5	11.3
2015	-1.97	44.0	44.8	11.1
Δ2004-15	-0.32	6.7	-8.6	1.9

Notes: The Pew Ideology score ranges from -10 (most liberal to +10 (most conservative). Columns 2-4 define liberals as those with scores of -10 to -3, moderates as those with scores from -2 to 2, and conservatives as those with scores from 3 to 10. Sample sizes for survey participants who reside in the 48 mainland states and who have complete demographic information are 1,994 in 2004, 3,016 in 2011, 9,868 in 2014, and 5,907 in 2015. Observations are weighted by survey weights.

Table A3: Summary Statistics: Import Shocks.

	2002-2010 (1)	2000-2008 (2)
Mean	0.71	0.90
25th Percentile	0.40	0.53
75th Percentile	0.90	1.11
P75 - P25	0.49	0.58

Notes: The 2002-2010 import shock in column 1 is used for the main analysis, and weights commuting zones by their adult voting-age population in 2000. The 2000-2008 import shock in column 2 is used for the analysis of presidential elections since 2000, and weights commuting zones by their number of votes in the 2000 presidential election.

Table A4: Congressional Redistricting, 2002 to 2016

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
			A. Prevalence of Redistricting				
Redistricted States	ME, PA, TX	GA, part of TX	none	none	42 States	none	FL, NC, VA
No. Districts w/o Changes	380	414	432	432	7	432	383
No. Districts w/ Changes	52	18	0	0	425	0	49
			B. Frequency of Party Change				
in Districts w/o Changes	2.9%	7.0%	7.4%	16.2%	0.0%	4.3%	1.4%
in Districts w/ Changes	19.4%	15.7%	n/a	n/a	15.2%	n/a	16.3%
End Year	2004	2006	2008	2010	2012	2014	2016

Notes: The three congressional districts of Alaska and Hawaii are excluded. Panel B indicates the population-weighted fraction of county-district cells that change party in an election, reported separately for districts without and with boundary changes.

Table A5: Exposure to Chinese Import Competition and Cable News Viewership, 2004 to 2016. Dependent Variable: Nielsen TV Rating or Nielsen Market Share

	2004-2008	2004-2012	2004-2016
	(1)	(2)	(3)
	A. Combined Ratings of TV News Networks		
Δ CZ Import Penetration x	-0.01	0.03	0.42
[t>2007]	(0.43)	(0.41)	(0.79)
	B. Market Share FOX News		
Δ CZ Import Penetration x	8.83	10.48	10.53
[t>2007]	(4.28)	(5.30)	(3.91)
	C. Market Share CNN		
Δ CZ Import Penetration x	-2.55	-4.18	-3.89
[t>2007]	(4.57)	(3.44)	(3.54)
	D. Market Share MSNBC		
Δ CZ Import Penetration x	-6.28	-6.28	-6.65
[t>2007]	(4.30)	(3.72)	(2.76)

$N = 6,813,\ 6,923, 6,890$ CZ-year-age-race cells in columns 1 through 3 respectively. All regressions include the full vector of control variables from column 6 in Table 2. Observations are weighted by Nielsen's estimate of the number of TV households in each cell, and standard errors are clustered on Commuting Zones.

Table A6: Exposure to Chinese Import Competition and FOX News Ratings, 2004 to 2016. Dependent Variable: Nielsen Market Share for FOX News by Age-Race Groups

	2004-2008	2004-2012	2004-2016
	(1)	(2)	(3)
	Market Share Fox News		
Δ CZ Import Penetration x	13.16	12.45	12.35
[t>2007] x N-H White 18-34	(5.88)	(6.90)	(5.79)
Δ CZ Import Penetration x	7.33	12.93	13.49
[t>2007] x N-H White 35-54	(5.25)	(6.34)	(4.98)
Δ CZ Import Penetration x	10.77	11.41	12.57
[t>2007] x N-H White 55+	(4.67)	(5.42)	(4.47)
Δ CZ Import Penetration x	10.93	15.32	4.37
[t>2007] x Other Group 18-34	(6.49)	(8.70)	(6.54)
Δ CZ Import Penetration x	6.60	7.76	7.32
[t>2007] x Other Group 35-54	(5.19)	(6.38)	(4.44)
Δ CZ Import Penetration x	1.78	-0.75	3.27
[t>2007] x Other Group 55+	(4.76)	(5.66)	(5.66)

Notes: $N = 5,110$, $5,097$, $5,037$ in columns 1, 2, and 3. In November 2011, the combined Nielsen rating of FOX News, CNN and MSNBC was 0.9/2.0/4.9 for young/middle-aged/older non-Hispanic whites, and 0.5/0.9/2.1 for young/middle-aged/older Hispanics and non-whites. The market share of FOX News was 63/66/61 and 49/44/37 percent in the six groups. All regressions include the full vector of control variables from column 5 in Table 1. Observations are weighted by Nielsen's estimate of the number of TV households in each cell, and standard errors are clustered on Commuting Zones.

Table A7: Exposure to Chinese Import Competition and Campaign Contributions, 2002-2004/2016. Dependent Variable: Relative Change in Contributions (in log points)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	A. Left-Wing Contributions (1st Tercile of Donor CF Score)						
Δ CZ Import Penetration	41.45	43.61	73.31	71.00	72.40	108.89	111.76
	(23.83)	(23.56)	(25.43)	(31.06)	(31.49)	(41.00)	(43.08)
	B. Moderate Contributions (2nd Tercile of Donor CF Score)						
Δ CZ Import Penetration	-3.08	3.05	18.56	23.60	23.97	13.79	13.13
	(11.79)	(13.97)	(15.24)	(19.63)	(19.58)	(33.97)	(36.72)
	C. Right-Wing Contributions (3rd Tercile of Donor CF Score)						
Δ CZ Import Penetration	17.88	12.18	57.65	46.05	47.12	-6.18	52.59
	(23.31)	(25.19)	(25.08)	(27.15)	(27.59)	(30.65)	(37.86)
End Year	2004	2006	2008	2010	2012	2014	2016
Controls	yes	yes	yes	yes	yes	yes	yes

Notes: $N = 3,772$ county-district cells. Panels A through C indicate the over-time change in contributions from donors whose CF score falls into the first, second, and third tercile of the dollar-weighted distribution of donor ideology in 2002. All regression estimated with 2SLS; the first-stage F-statistic is 29.2. All regressions estimated with the full set of controls. Industry and occupation controls are measured at the CZ level and comprise the fraction of CZ employment in the manufacturing sector and the Autor and Dorn (2013b) routine share and offshorability index of a CZ's occupations. Census division dummies allow for different time trends across the nine geographical Census divisions. Demographic controls comprise the percentage of a county's population in nine age and four racial groups, as well as the population shares that are female, college-educated, foreign-born, and Hispanic. Election controls comprise the Republican two-party vote share in the presidential elections of 1992 and 1996, measured at the county level. Observations are weighted by a county-district cell's share in the total year-2000 voting age population of a district, so that each district has a total weight of one.

Table A8: Exposure to Chinese Import Competition and Electoral Results, 2002-2010. Dependent Variables: Change in Republican Two-Party Vote Share, Change in Republican Win Probability (in % pts)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	A. Change in Republican Win Probability						
Δ CZ Import Penetration	-0.27	-3.96	-2.54	24.08	24.08	26.05	27.07
	(5.07)	(7.12)	(8.13)	(12.07)	(12.07)	(12.19)	(12.37)
	B. Change in Republican Vote Share						
Δ CZ Import Penetration	4.85	-1.41	-5.35	-1.08	-1.05	-2.94	-6.39
	(4.22)	(5.45)	(6.23)	(5.98)	(5.98)	(7.27)	(6.31)
End Year	2004	2006	2008	2010	2012	2014	2016
Controls	yes	yes	yes	yes	yes	yes	yes

Notes: N = 3772 county-district cells. The dependent variable is the over-time change in probability of a Republican candidate winning the election (panel A) and the change in the Republican share of the two-party vote (panel B). The base year is 2002. All regression estimated with 2SLS; the first-stage F-statistic is 29.2. All regressions estimated with the full set of controls. Industry and occupation controls are measured at the CZ level and comprise the fraction of CZ employment in the manufacturing sector and the Autor and Dorn (2013b) routine share and offshorability index of a CZ's occupations. Census division dummies allow for different time trends across the nine geographical Census divisions. Demographic controls comprise the percentage of a county's population in nine age and four racial groups, as well as the population shares that are female, college-educated, foreign-born, and Hispanic. Election controls comprise the Republican two-party vote share in the presidential elections of 1992 and 1996, measured at the county level. Observations are weighted by a county-district cell's share in the total year-2000 voting age population of a district, so that each district has a total weight of one.

Table A9: Exposure to Chinese Import Competition and Electoral Results, 2002-2016. Dependent Variables: Change in Republican Win Probability (in % pts)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	A. Majority White Counties						
Δ CZ Import Penetration	-0.20	-5.29	-2.73	28.78	28.78	30.98	31.23
	(5.95)	(8.14)	(9.21)	(14.65)	(14.65)	(14.90)	(15.09)
	B. Majority Nonwhite Counties						
Δ CZ Import Penetration	2.27	7.10	34.18	-0.45	-0.45	-5.69	-5.69
	(3.45)	(4.37)	(17.32)	(9.17)	(9.17)	(10.97)	(10.97)
End Year	2004	2006	2008	2010	2012	2014	2016
Controls	yes	yes	yes	yes	yes	yes	yes

Notes: $N = 3,772$ county-district cells. Panels A and B indicate the over-time change in probability of a Republican candidate winning the election. The base year is 2002. All regression estimated with 2SLS; the first-stage F-statistic is 29.2. All regressions estimated with the full set of controls. Industry and occupation controls are measured at the CZ level and comprise the fraction of CZ employment in the manufacturing sector and the Autor and Dorn (2013b) routine share and offshorability index of a CZ's occupations. Census division dummies allow for different time trends across the nine geographical Census divisions. Demographic controls comprise the percentage of a county's population in nine age and four racial groups, as well as the population shares that are female, college-educated, foreign-born, and Hispanic. Election controls comprise the Republican two-party vote share in the presidential elections of 1992 and 1996, measured at the county level. Observations are weighted by a county-district cell's share in the total year-2000 voting age population of a district, so that each district has a total weight of one. Standard errors are two-way clustered on counties and congressional districts.

Table A10: Exposure to Chinese Import Competition and Change in Ideological Position of Election Winner , 2002-2010. Dependent Variable: 100 x Change in Indicators for Election of Politician by Party and Political Position

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
				A. Liberal Democrats			
Δ CZ Import Penetration	3.72	-0.53	5.75	-8.43	-7.91	-1.12	-2.52
	(5.96)	(7.94)	(9.66)	(8.79)	(8.78)	(10.68)	(10.27)
				B. Moderate Democrats			
Δ CZ Import Penetration	-3.45	4.49	-3.21	-15.64	-16.16	-25.34	-23.05
	(7.74)	(8.60)	(10.17)	(13.92)	(13.95)	(16.04)	(15.91)
				C. Moderate Republicans			
Δ CZ Import Penetration	-4.55	-6.70	-10.55	-5.83	-5.83	-6.22	-1.38
	(6.99)	(9.01)	(10.79)	(12.17)	(12.17)	(13.01)	(12.35)
				D. Conservative Republicans			
Δ CZ Import Penetration	4.29	2.74	8.01	29.88	29.88	28.82	26.84
	(8.26)	(9.38)	(12.42)	(15.89)	(15.89)	(14.96)	(15.61)
End Year	2004	2006	2008	2010	2012	2014	2016
Controls	yes	yes	yes	yes	yes	yes	yes

Notes: $N = 3,772$ county-district cells. Panels indicate the over-time change in ideology of election winners by party and ideology for liberal Democrats, moderate Democrats, moderate Republicans, and conservative Republicans. The base year is 2002. All regression estimated with 2SLS; the first-stage F-statistic is 29.2. All regressions estimated with the full set of controls. Industry and occupation controls are measured at the CZ level and comprise the fraction of CZ employment in the manufacturing sector and the Autor and Dorn (2013b) routine share and offshorability index of a CZ's occupations. Census division dummies allow for different time trends across the nine geographical Census divisions. Demographic controls comprise the percentage of a county's population in nine age and four racial groups, as well as the population shares that are female, college-educated, foreign-born, and Hispanic. Election controls comprise the Republican two-party vote share in the presidential elections of 1992 and 1996, measured at the county level. Observations are weighted by a county-district cell's share in the total year-2000 voting age population of a district, so that each district has a total weight of one. Standard errors are two-way clustered on counties and congressional districts.

Table A11: Exposure to Chinese Import Competition and Change in Ideological Position of Election Winner, 2002-2004/2016. Heterogeneity by Initial Local Racial Composition. Dependent Variable: 100 x Change in Indicators for Election of Politician by Party and Political Position

I. Congressional Districts with Majority Non-Hispanic White Population in 2000

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
				A. Liberal Democrats			
Δ CZ Import Penetration	4.82	2.07	11.98	-11.34	-10.45	-8.91	-10.17
	(6.83)	(8.96)	(10.51)	(9.51)	(9.45)	(10.87)	(10.22)
				B. Moderate Democrats			
Δ CZ Import Penetration	-4.62	3.21	-9.24	-17.44	-18.33	-22.28	-19.11
	(8.72)	(9.59)	(11.74)	(16.50)	(16.55)	(18.49)	(18.08)
				C. Moderate Republicans			
Δ CZ Import Penetration	-4.90	-5.78	-11.48	-6.09	-6.09	-7.48	-0.78
	(7.89)	(10.15)	(12.34)	(14.48)	(14.48)	(15.27)	(14.68)
				D. Conservative Republicans			
Δ CZ Import Penetration	4.70	0.49	8.74	34.91	34.91	34.28	30.15
	(9.33)	(10.78)	(14.61)	(19.14)	(19.14)	(17.74)	(18.75)
End Year	2004	2006	2008	2010	2012	2014	2016
Controls	yes	yes	yes	yes	yes	yes	yes

II. Congressional Districts with Minority Non-Hispanic White Population in 2000

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
				A. Liberal Democrats			
Δ CZ Import Penetration	11.40	-4.86	-29.27	30.12	30.12	59.90	43.89
	(15.90)	(15.80)	(18.64)	(17.71)	(17.71)	(19.47)	(16.00)
				B. Moderate Democrats			
Δ CZ Import Penetration	-13.67	-2.24	-4.91	-29.67	-29.67	-59.00	-42.98
	(16.72)	(15.34)	(17.06)	(14.51)	(14.51)	(14.54)	(13.24)
				C. Moderate Republicans			
Δ CZ Import Penetration	-5.59	4.33	4.12	8.19	8.19	0.40	-1.28
	(5.02)	(3.91)	(4.10)	(8.77)	(8.77)	(6.96)	(7.25)
				D. Conservative Republicans			
Δ CZ Import Penetration	7.86	2.77	30.06	-8.64	-8.64	-6.09	-4.41
	(5.17)	(6.73)	(18.05)	(8.50)	(8.50)	(10.24)	(9.64)
End Year	2004	2006	2008	2010	2012	2014	2016
Controls	yes	yes	yes	yes	yes	yes	yes

Notes: $N = 3,772$ county-district cells. Panels indicate the over-time change in ideology of election winners by party and ideology for liberal Democrats, moderate Democrats, moderate Republicans, and conservative Republicans, for counties that were majority non-Hispanic white in 2000 (panel I) and those that were minority non-Hispanic white in 2000 (panel II). The base year is 2002. All regression estimated with 2SLS; the first-stage F-statistic is 29.2. All regressions estimated with the full set of controls. Industry and occupation controls are measured at the CZ level and comprise the fraction of CZ employment in the manufacturing sector and the Autor and Dorn (2013b) routine share and offshorability index of a CZ's occupations. Census division dummies allow for different time trends across the nine geographical Census divisions. Demographic controls comprise the percentage of a county's population in nine age and four racial groups, as well as the population shares that are female, college-educated, foreign-born, and Hispanic. Election controls comprise the Republican two-party vote share in the presidential elections of 1992 and 1996, measured at the county level. Observations are weighted by a county-district cell's share in the total year-2000 voting age population of a district, so that each district has a total weight of one.