Importing Political Polarization?

The Electoral Consequences of Rising Trade Exposure[*]

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Abstract

Has rising trade integration between the U.S. and China contributed to the polarization of U.S. politics? Analyzing outcomes from the 2002 and 2010 congressional elections, we detect an ideological realignment that is centered in trade-exposed local labor markets and that commences prior to the divisive 2016 U.S. presidential election. Exploiting the exogenous component of rising trade with China and classifying legislator ideologies by their congressional voting record, we find strong evidence that congressional districts exposed to larger increases in import competition disproportionately removed moderate representatives from office in the 2000s. Trade-exposed districts initially in Republican hands become substantially more likely to elect a conservative Republican, while trade-exposed districts initially in Democratic hands become more likely to elect either a liberal Democrat or a conservative Republican. Polarization is also evident when breaking down districts by race: trade-exposed locations with a majority white population are disproportionately likely to replace moderate legislators with conservative Republicans, whereas locations with a majority non-white population tend to replace moderates with liberal Democrats. In contrast with much previous work in political science, we find limited impacts of economic shocks on the probability of party turnover (an anti-incumbency effect) or on the electoral vote shares of the major parties (a party realignment effect). Focusing on legislator behavior rather than on party vote counts, we find that trade exposure abets the replacement of incumbents from both parties with more ideologically strident successors.

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1 Introduction

The 2016 U.S. presidential election has been as divisive as any in recent memory. The rancor on the campaign trail mirrors the partisan divide in Congress, a split that has been widening since the 1980s. DW-Nominate scores (Poole and Rosenthal, 1985 and 1991), which rank congressional representatives on a liberal-conservative scale according to their roll-call votes, show that the ideological gap between the parties is now at historic highs (McCarty, Poole, and Rosenthal, 2006). This polarization is due to a substantial rightward shift among congressional Republicans and a modest leftward shift among congressional Democrats, such that few centrists remain in either party. In the mid 1970s, the voting of more-liberal congressional Republicans overlapped with that of more-conservative representatives in the Democratic Party; today, even the most liberal House Republican is more conservative than the most conservative House Democrat.[1]

In this paper, we examine whether the exposure of local labor markets to increased foreign competition has exacerbated partisan divisions in the U.S. Congress. We estimate the impact of rising manufacturing imports from China on congressional elections and voting by congressional representatives over the period 2002 to 2010. Our interest is in seeing whether adverse economic shocks related to international trade may cause voters or their elected legislators to support positions that lean towards political extremes on the left or right.

As an indication of the prominence of globalization in U.S. political debates, presidential candidates from both major parties have agreed on the need to take a much more aggressive line on international trade. On the right, Republican candidate Donald Trump pledged to impose a 45% tariff on U.S. imports from China;[2] on the left, Democratic hopeful Bernie Sanders categorically rejected supporting any trade deal, including the Trans-Pacific Partnership, which had been negotiated by a Democratic president.[3] Of course, using foreign trade to stoke voter anxiety is nothing new in American politics. During the 1992 U.S. presidential campaign, candidate Ross Perot famously quipped that if approved, the North American Free Trade Agreement would create a "giant sucking sound" as U.S. manufacturing factories

[1]Members of the two parties last overlapped in their voting behavior in the 2001-2003 congress, see Figure 3 below.

[2]See "Donald Trump Says He Favors Big Tariffs on Chinese Exports," *New York Times*, Jan. 7, 2016.

[3]See http://www.ontheissues.org/2016/Bernie_Sanders_Free_Trade.htm.

were pulled into Mexico.

Whatever the merits or demerits of populist trade policies, there is now abundant evidence linking international trade to the decline of U.S. manufacturing jobs. From the 1950s to the 1980s, manufacturing allowed U.S. workers without a college degree to attain a middle-class lifestyle. Such opportunities have largely disappeared, leaving the U.S. economy more partitioned between workers in highly paid managerial and professional occupations and workers in low-wage service occupations (Autor and Dorn, 2013). Manufacturing industries more exposed to import competition from China have seen higher rates of plant exit (Bernard, Jensen, and Schott, 2006), larger contractions in employment (Pierce and Schott, 2016; Acemoglu, Autor, Dorn, Hanson and Price, 2016), and lower lifetime incomes for affected workers (Autor, Dorn, Hanson, and Song, 2014). The local labor markets that are home to more-exposed industries have endured substantial employment reductions and persistent increases in rates of unemployment, non-participation in the labor force, and uptake of government transfers (Autor, Dorn, and Hanson, 2013). The political implications of the decline in U.S. manufacturing are not yet well understood.

To structure our analysis, we consider three mechanisms through which trade shocks may affect regional political divisions. One is the well-known *anti-incumbent effect*. A large literature, beginning with Fair (1978), has found that economic downturns are bad for sitting politicians and their parties.[4] Voters punish incumbents at the polls for negative economic outcomes, including those caused by greater import competition (Margalit, 2011; Jensen, Quinn, and Weymouth, 2016). We verify in our data that adverse trade shocks diminish vote shares for the party initially in power. However, such patterns cannot explain greater partisanship, as on their own they imply that regions suffering a persistent trade-induced decline in manufacturing would simply alternate their support for political parties or churn through elected legislators.

A second mechanism through which trade shocks may affect political cohesion is through a *realignment effect*. Previous literature documents that when individuals perceive actual or expected declines in economic opportunity, their political preferences shift in favor of redistribution (Alesina and La Ferrara, 2005; Bruner, Ross, and Washington, 2011; Giuliano

[4]See Lewis-Beck and Stegmaier (2000) and Anderson (2007) for reviews of the literature on how economic conditions affect electoral outcomes.

and Spilimbergo, 2014). In response to adverse trade shocks, these patterns would indicate greater support for Democrats, who tend to favor progressive taxation and generous social spending. Che, Lu, Pierce, Schott, and Tao (2015) present evidence that over the period 1998 to 2010, U.S. counties exposed to greater import competition had larger increases in vote shares for Democratic candidates, though it is unclear how vote-share changes at the county level translate into electoral outcomes at the congressional-district level. In Europe, the exact opposite response has been observed. French and German regions that have been more exposed to trade with low-wage countries have seen larger increases in vote shares for extreme-right parties (Malgouyres, 2014; Dippel, Gold, and Heblich, 2015).[5] These impacts, while striking, mainly affect parties outside of the political mainstream that appear to have limited chance of gaining national power. In our data, we do not find that adverse trade shocks induce a realignment in vote shares toward either major political party.

A third mechanism, which has received less attention in the literature and for which we find strong support, is a *polarization effect*, in which a negative economic shock increases the electoral success of both non-centrist left-wing and right-wing politicians. Holding constant political conditions in 2002—including the party in power, the vote share of the winning party, and the DW-Nominate score of the initial office holder—districts exposed to larger increases in import competition are substantially less likely to elect a moderate legislator in 2010. More trade-exposed districts see larger moves away from the political center, as measured by the absolute change in DW-Nominate scores of elected legislators. This shift is due not to changes in the voting behavior of existing office holders but to the election of more extreme legislators, especially on the right. Greater trade exposure makes districts initially in Republican hands substantially more likely to elect a conservative Republican, and, in some specifications, initially Democratic districts more likely to elect a liberal Democrat. A polarization effect is also manifest when separating locations by race rather than by party. In response to an increase in import competition, trade-affected regions with an initial population majority of non-Hispanic whites are more likely to elect a conservative Republican, whereas trade-affected districts with white minorities are more likely to elect a liberal Democrat.

Why might trade shocks differentially affect political outcomes across regions? One answer

[5] Similar outcomes obtained during the Great Depression, when European countries that had more prolonged downturns saw greater support for far-right political movements (de Bromhead, Eichengreen, and O'Rourke, 2013).

is that voters differ in their political preferences and over time have become more segregated geographically, such that party affiliation is now more unevenly distributed across space (Bishop, 2009). Yet, evidence that geographic political segregation has increased is weak (Abrams and Fiorina, 2012). While there has long been regional heterogeneity in cultural and political attitudes, this heterogeneity does not appear to have become more pronounced in recent decades (Glaeser and Ward, 2006; Ansolabehere, Rodden, and Snyder, 2008; Fiorina and Abrams, 2008). Alternatively, voters may differ not in their preferences—e.g., how to trade off economic efficiency against social equity—but in their beliefs about how the economy operates, such as whether we live in a "liberal" world in which government intervention is helpful or a "conservative" world in which it is harmful. Dixit and Weibull (2007) provide a theory that accounts for how economic shocks, be they related to trade or other events, may induce divergence in beliefs across groups of otherwise similar individuals. Suppose that two groups (Democrats, Republicans) agree on objectives (to minimize the loss from government policy) but differ in their prior beliefs about the state of the world (whether the world is liberal or conservative). These groups may respond to the same signal—say, a change in aggregate income—by updating their beliefs in opposite directions, with liberals becoming more convinced that the world is "liberal" and conservatives feeling similarly more confident that the world is "conservative."[6] While such polarization does not last forever, convergence of the two groups to common posterior beliefs need not occur quickly or monotonically. Acemoglu, Chernozhukov, and Yildiz (2015) show that with uncertainty about the distribution of the signal, convergence need not occur at all.

Our findings do not support the hypothesis that trade shocks have made red states redder or blue states bluer, in the narrow sense of increasing the regional dominance of a single party, which is often how political polarization is characterized (e.g., Gelman, Park, Shor, Bafumi, and Cortina, 2008). These results highlight the importance of studying the subsequent voting behavior of winning candidates, as captured by DW-Nominate scores, rather than the share of the vote that winning candidates garner. By favoring those at the extremes of their respective parties, and in particular of conservative Republicans, greater trade exposure does not help

[6] The technical condition needed for this outcome to occur is failure of the monotone likelihood ratio property for the probability density of the observed outcome (e.g., aggregate income) conditional on the policy enacted (e.g., income redistribution) and the state of the world (e.g., liberal or conservative). See Andreoni and Mylovanov (2012) on polarization of opinions in lab experiments and Rodrik (2014) on how worldviews shape policy choices.

candidates to win lopsided electoral victories, but rather abets the electoral ascendance of more ideologically strident politicians.

In the empirical analysis, our task is complicated by the fact that local labor markets, which we take to be commuting zones (CZs), do not map one-to-one into congressional districts. Whereas CZs are aggregations of contiguous counties, gerrymandering creates districts that may combine bits and pieces of multiple counties and span several commuting zones. Our solution is to divide the continental U.S. into county-by-congressional-district cells, attach each cell to its corresponding CZ, and weight each cell by its share of total votes in its congressional district. This approach maps trade exposure at the commuting-zone level to political outcomes at the congressional-district level. Much other work takes the unit of analysis to be the county rather than the county-district intersection, which complicates studying which parties win elections.

To measure regional trade exposure, we follow Autor, Dorn, and Hanson (2013) and use the change in import competition from China at the level of four-digit manufacturing industries, weighting each industry by its share of CZ employment in a pre-sample period. Commuting zones differ not just in the overall importance of manufacturing for their local economies but also in the pattern of specialization within manufacturing, which creates strong regional variation in exposure to import competition. Our use of initial-period values for regional industry employment ensures that these specialization patterns are not themselves the consequence of contemporaneous trade exposure. We isolate the component of U.S. import growth that is driven by export-supply growth in China, and not by U.S.-specific product-demand shocks, using the identification strategy in Autor, Dorn, and Hanson (2013). This approach instruments for the change in U.S. industry import growth using growth in industry imports from China in high-income economies other than the U.S.

The time span for our analysis, which stretches from the first mid-term election of the George W. Bush administration in 2002 to the first mid-term election of the Barack Obama administration in 2010, is dictated by the periodic nature of congressional redistricting. Because boundaries for congressional districts are redrawn after each decennial census, we choose to analyze the longest recent time period for which district boundaries remain unchanged in most states. A side benefit of our period of study is that 2002 and 2010 are not presidential election years, in which voting may be influenced as much by the popularity of candidates at the head of the ticket as by candidates for congressional office. Helpfully, the first year of our sample follows China's accession to the World Trade Organization in 2001, which contributed to a dramatic surge in trade that lifted the country's share of world manufacturing exports from 4.8% in 2000 to 15.1% in 2010. The last year in our sample coincides with the rise of the Tea Party movement and the nomination of strongly conservative candidates for elected office (Madestam, Shoag, Veuger, and Yanagizawa-Drott, 2013).

Our paper provides the first evidence that connects adverse trade shocks to political polarization.[7] It contributes to a growing literature on political partisanship, which has yet to reach a consensus on its origins (Fiorina and Abrams, 2008; Gentzkow, Shapiro, and Taddy, 2015). Bartels (2010) and Hacker and Pierson (2010) contend that changes in tax and regulatory policies in the 1970s and 1980s shifted income in favor of the wealthy, who then expanded support for politicians favoring their interests. It is readily apparent that greater political polarization in Congress tracks rising income inequality.[8] However, the evidence linking the concentration in income induced by policy reform to changes in the ideological orientation of the two major parties remains circumstantial. McCarty, Rosenthal, and Poole (2006) suggest that political polarization is due in part to immigration, which has increased the fraction of the poor who are ineligible to vote, thereby reducing electoral support for redistribution. Popular treatments, including Frank (2004), look instead to culture, portraying wealthy elites as manipulating blue-collar workers in the South and Midwest to vote against their economic interests by diverting attention to social issues, such as abortion, gun rights, and gay marriage. Gelman, Park, Shor, Bafumi, and Cortina (2008) argue against either immigration or cultural attitudes as causing greater partisanship, finding that regional voting patterns support neither hypothesis.

Other related work includes the many studies of how congressional representatives vote on trade legislation.[9] Specifically on the impact of import competition from China, Feigenbaum

[7]Other factors that may contribute to political polarization include greater partisanship in the media (DellaVigna and Kaplan, 2007; Levendusky, 2013; Prior, 2013) and stronger ideological sorting of voters by party (Levendusky, 2009). The structure of primary elections, rule changes in Congress, and gerrymandering appear unable to explain the phenomenon (McCarty, Poole and Rosenthal, 2009; Barber and McCarty, 2015). Parker and Barreto (2013) address the rise of the Tea Party, arguing that it is grounded in racial politics and fueled by opposition to President Obama.

[8]See, e.g., Figure 1 in Barber and McCarty (2015).

[9]See, e.g., Bailey and Brady (1998), Baldwin and Magee (2000), Beaulieu (2002), Hiscox (2002), Fordham and McKeown (2003), and Milner and Tingley (2011).

and Hall (2015) and Che, Lu, Pierce, Schott, and Tao (2015) find that support for protectionist trade bills is stronger among politicians from more trade-exposed districts. Similarly, Kleinberg and Fordham (2013) and Kuk, Seligsohn, and Zhang (2015) find that representatives from congressional districts harder hit by the China trade shock are more likely to support foreign-policy legislation that takes a hard line against China. Our work shows that the impacts of trade exposure extend far beyond U.S. policy initiatives and affect the ideological composition of Congress itself.

In section 2, we describe our data on congressional elections, congressional voting patterns, and how we match congressional districts to commuting zones, followed by section 3, which summarizes our data on trade shocks and local labor market conditions. In section 4, we present our main empirical results on the impacts of trade shocks on voting outcomes. Section 5 considers proximate mechanisms, including changes in political parties versus within-party changes in the identity of representatives versus within-representative changes in voting behavior, as well as demographic predictors of changes in the ideology of elected officials. Section 6 concludes.

2 Measuring Electoral Outcomes in Congressional Districts

In a first step of data construction, we combine electoral outcomes for congressional districts with DW-Nominate scores of elected representatives in the congressional terms that succeed the first and last years of the sample period. In a second step, we match these data to contemporaneous economic conditions in commuting zones, including the exposure of these local labor markets to import competition from China. In combination, the full data allow us to analyze the impact of commuting-zone-level trade shocks on congressional-district-level political results.

Our geographic unit of analysis is the county-by-congressional-district cell. The functional building blocks of congressional districts are census tracts, whose amalgamation allows officials to construct districts that meet the constitutional requirement of each holding approximately $1/435^{th}$ of the U.S. population. The area that constitutes a district must be contiguous and lie within the boundaries of a state but may combine sections of multiple counties. Counties, in turn, are the building blocks of CZs, which are clusters of adjoining locations that have the commuting structure of a local labor market (Tolbert and Sizer, 1996; Dorn, 2009). In the empirical analysis, we ascribe to each county-district cell the CZ-level shock that corresponds to the county. We then weight each cell by its share of votes 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 vote-share-weighted average of values in these CZs.

From the full sample of 435 congressional districts, we omit Alaska's one congressional district and Hawaii's two congressional districts because commuting zones are difficult to define for these states. We also omit the one district for Vermont, whose elected congressional representative over the sample period is an independent and thus is attached to neither major political party. In the remaining territory, the two states of Texas and Georgia carried out intercensal redistricting during the 2000s. As a consequence, a set of county-district cells that sum to 15 congressional districts cannot be continuously observed over time, and have to be omitted from the analysis.[10] The resulting set of 3,503 county-district cells covers 416 congressional districts, or approximately 96% of the U.S. population, over the period 2002 to 2010. Table A1 in the Appendix summarizes these details.

Data on election outcomes in county-district cells are from Dave Leip's Atlas of U.S. Presidential Elections.[11] These data track the number of votes received by Democratic, Republican, and other candidates for Congress and for other major offices in each county, in each congressional district, and in each election year. We use these data to tabulate the shares of votes won by Democratic and Republican congressional candidates in each county-by-district cell in 2002 and the change in these shares between 2002 and 2010. We also note whether the winning candidates in 2002 or 2010 ran unopposed, and whether the winner of the 2002 election remained in office after the 2010 election. In part of the analysis, we also utilize county-level votes in the 2000 U.S. presidential election as a measure of initial party orientation by location. As this election was closely contested between two candidates—with Democrat Al Gore winning the popular vote and Republican George W. Bush winning the electoral-college vote—it provides a convenient metric of how votes in the pre-sample period

[10] Nine of these districts are in Texas and five are in Georgia, corresponding to about one third of all districts in each state.

[11] See http://uselectionatlas.org/.

separate along the Democratic-Republican divide.

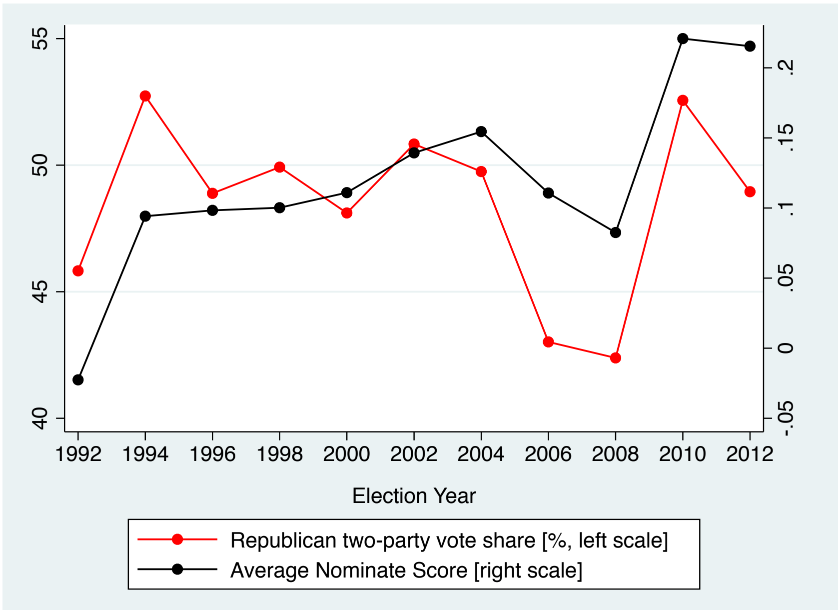
To measure variation in the political orientation of congressional representatives, we use Poole-Rosenthal DW-Nominate scores (Poole and Rosenthal, 1985; McCarty, Rosenthal, and Poole, 2006), which are widely applied in political science and are the foundation for analysis of political polarization in Congress. DW-Nominate uses roll-call (or recorded) votes in the U.S. House of Representatives and the U.S. Senate to categorize elected officials on an ideological scale from liberal to conservative. This score is based on a multidimensional scaling technique, in which one assumes that each piece of legislation can be represented by two points (one for a yea vote, one for a nay vote) in Euclidean space, that each legislator has a well-behaved utility function defined over this space, and that each legislator chooses her vote non-strategically to maximize her static utility, such that one can use a static random utility model to characterize each legislator's yea-or-nay choice. If one observes a common set of legislators voting on many bills, one can estimate the parameters of the utility function and rank legislators in each of the Euclidean dimensions.[12]

Because DW-Nominate is estimated using roll-call votes for all 113 U.S. Congresses, each of which contain a large number of overlapping members from one Congress to another, parameters are comparable across time. For most of U.S. congressional history, DW-Nominate scores exhibited little gain in explanatory power when allowing for more than two Euclidean dimensions; since the early 1980s, by which time the post-1964 realignment of southern conservatives from the Democratic to the Republican Party was complete, there is only modest gain in going beyond *one* dimension (McCarty, Rosenthal, and Poole, 2006). The DW-Nominate score that we use is the ranking of legislators along the primary dimension, which Poole and Rosenthal (1997) describe as a measure of liberal-conservative ideology. Hence-

[12]Each legislator is assumed to have an ideal point in two-dimensional Euclidean space. Each piece of legislation is described by two points in this space, one for a yea outcome and another for a nay outcome. Utility to a legislator from a particular outcome has a non-stochastic component, which is an exponential function of the distance between that outcome and the legislator's ideal point, and a normally distributed iid stochastic component. The non-random component of utility has three sets of parameters: a scalar proportional to the variance of the stochastic component of utility, the coordinates of the legislator's ideal point in the two-dimensional plane, and a pair of weighting parameters in utility, one for each coordinate; the first and third parameters are common across legislators. The coordinate of the first dimension for the legislator's ideal point is interpreted as the liberal-conservative index; the coordinate of the second dimension has been interpreted as party loyalty. Legislators ideal points are allowed to change over time according to a linear time trend. Since the post-civil rights movement realignment of the two major parties, most of the explanatory power of the model comes from the first dimension. See Poole and Rosenthal (1997, 2001) and McCarty, Rosenthal, and Poole (2006).

forth, we refer to the first dimension of DW-Nominate as simply the Nominate score.[13]

Figure 1: Republican Vote Shares and Average Nominate Scores for Congressional Districts



Over the 113 U.S. congresses, the Nominate score in the House is roughly centered on zero (mean= 0.02, standard deviation= 0.38), where the average value over time for each legislator is constrained to lie between 1 (most conservative) and −1 (most liberal). Figure 1 shows average Nominate scores for 1992 to 2012. Our convention is to define the year to be the calendar year in which representatives are elected, which precedes the two-year congressional term on which Nominate scores are based. For example, we use 2002 to represent the 108th Congress, such that the Nominate scores we ascribe to 2002 are based on roll-call votes between 2003 and 2005. This convention is in keeping with the Poole-Rosenthal interpretation of Nominate scores as describing the ideology of legislators. Roll-call votes subsequently reveal these ideologies, which were presumably known to the candidates and to voters at the time of the preceding election.

[13]If a representative is replaced in a special election during a Congressional term, we use only the Nominate score of the winner of the *initial* general election. In the small number of cases where an elected representative casts too few votes to allow for the computation of a Nominate score (typically because the legislator died or quit shortly after being reelected), we are able to approximate the representative's political position using her Nominate score based on her voting record in the preceding Congress. We also adopt the convention of the DW-Nominate data of defining a legislator's identity as a combination of person and party, such that in the rare cases where a representative changes party, this is treated as the replacement of the one legislator with another from the newly adopted party.

Average Nominate scores rise over time from -0.02 in 1992 for the 103th Congress (1993-1995) to 0.22 in both 2010 and 2012 for the 112th (2011-2013) and 113th (2013-2015) Congresses. During our sample period of 2002 to 2010, there is a noticeable jump in Nominate scores at the end of the time span, associated with the election of several dozen strongly conservative Tea Party Republicans in the final year. The progressive rise in Nominate scores over time corresponds with the strengthening of the Republican majority in the House of Representatives. Whereas the House was in Democratic hands for all but two congresses between 1931 and 1995, the Republican Party took control of the House in the 1994 election, following the success of Newt Gingrich's "Contract with America," and has held the chamber for all but two terms since.[14]

In Figure 1, it appears that the average vote share for the Republican party and the average Nominate score track each other over time.[15] The time-series co-movement in these means does not imply, however, that electoral vote shares have increased for legislators who are more conservative; higher Republican vote shares simply mean that Republicans are winning more elections. Figure 2 separates districts by whether the winning candidate was a Democrat or a Republican. For districts won by Republicans (Figure 2a), average Nominate scores have risen over time, whereas since the mid 1990s, Republican average vote shares have fallen, though of course they remain well above 50%. While Republicans have nominated more conservative candidates and these candidates have succeeded in winning elections, the party's win margins are smaller than in the past. The party's more conservative candidates may be repelling some voters, resulting in narrower electoral victories. An alternative interpretation would be that as Republicans have picked up seats in the House they have won in more closely competitive districts, with correspondingly smaller victory margins. This logic would also imply that Democrats should be winning their races by *larger* vote margins, which Figure 2b reveals is not the case. Democratic legislators have on average moved to the left, as indicated by increasing liberalness in the inverted Nominate scores in Figure 2b, but the magnitude of the change is much smaller than among Republicans. Figure 2 shows that in the 1990s there was polarization both in vote shares—in that winning candidates won

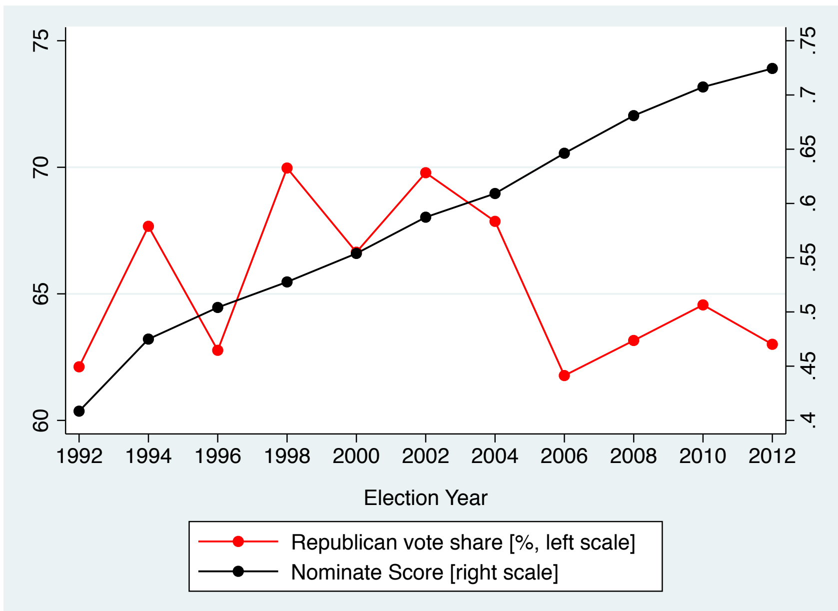
[14]The two House terms under Republican control between 1931 and 1995 were the 80th (1947-1949) and 83rd (1953-1955); the two terms under Democratic control since 1995 were the 110th (2007-2009) and 111th (2009-2011).

[15]The Republican two-party vote share in Figure 1 corresponds to the ratio of Republican votes to the sum of Republicans and Democratic votes.

by larger margins—and in Nominate scores. In the 2000s, polarization is only evident in Nominate scores, which highlights the difficulty in using vote shares alone to study changes in partisanship.

Figure 2: Vote Shares and Nominate Scores by Winning Party

(a) Congressional Districts Won by Republicans



(b) Congressional Districts Won by Democrats

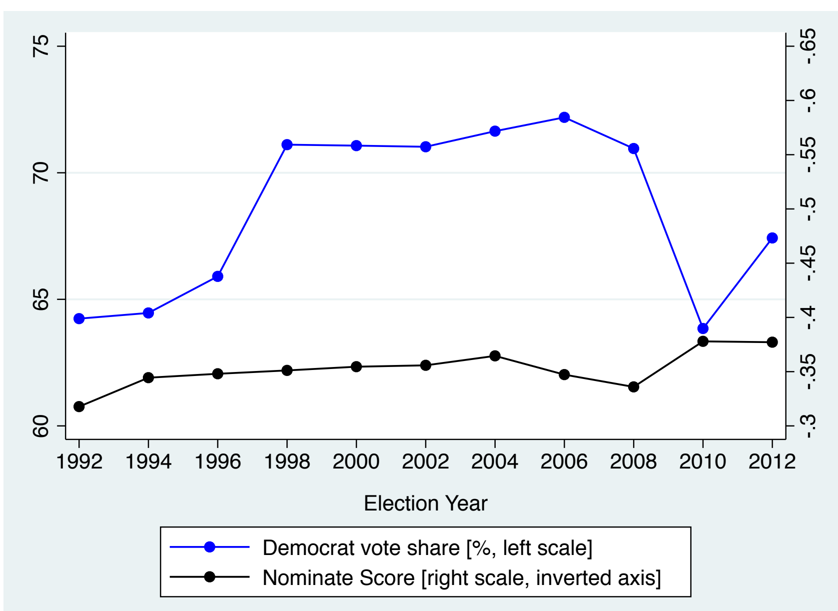
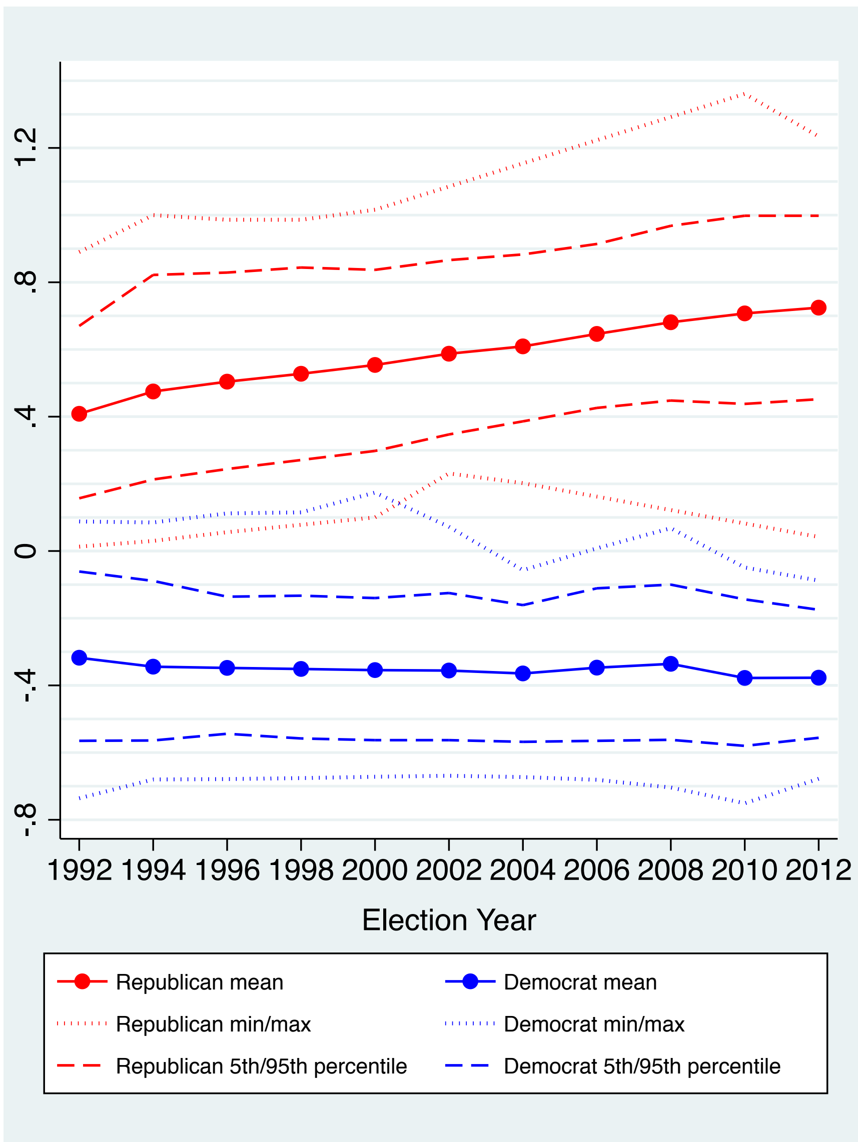


Figure 3 shows the central tendency and spread of Nominate scores for Democratic and Republican representatives, which illustrates the widening partisan divide in Congress. Republican legislators have become markedly more conservative, with their average Nominate score rising from 0.41 in 1992 to 0.72 in 2010, a change equal to 0.82 of the Nominate standard deviation for all 113 Congresses. Democrats, for their part, have become somewhat more liberal, with average Nominate scores falling from -0.32 to -0.38 over the same period, a 0.16 standard-deviation change. Notably, the ideological dispersion of Republican representatives has risen dramatically, with the gap between the least and most conservative (and their distance from the mean Republican) growing substantially after 2004; we do not observe a similar pattern for Democrats. Despite the widening ideological range of elected Republicans, there is *no* ideological overlap after 2000 between the most conservative Democrat and the most liberal Republican.

Figure 3: Polarization in Nominate Scores



3 Measuring Local Labor Market Exposure to Trade

In the empirical analysis, we follow the specification of local trade exposure derived by Autor, Dorn, Hanson and Song (2014) and Acemoglu, Autor, Dorn, Hanson and Price (2016). Our measure of the local-labor-market shock is the average change in Chinese import penetration in a CZ's industries, weighted by each industry's share in initial CZ employment:

$$\Delta IP_{i\tau}^{cu} = \sum_j \frac{L_{ijt}}{L_{it}} \Delta IP_{j\tau}^{cu}. \tag{1}$$

In this expression, $\Delta IP_{j\tau}^{cu} = \Delta M_{j\tau}^{cu}/(Y_{j0} + M_{j0} - X_{j0})$ is the growth of Chinese import penetration in the U.S. for industry j between the start and end of period τ. Following Acemoglu, Autor, Dorn, Hanson and Price (2016), it is computed as the growth in U.S. imports from China during the outcome period 2002-2010, $\Delta M_{j\tau}^{cu}$, divided by initial absorption (U.S. industry shipments plus net imports, $Y_{j0} + M_{j0} - X_{j0}$) in the base period 1991, near the start of China's export boom. The fraction L_{ijt}/L_{it} is the share of industry j in CZ i's total employment, as measured in County Business Patterns data prior to the outcome period in the year 2000.

In equation (1), the difference in ΔIP_{it}^{cu} across commuting zones stems entirely 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. Importantly, differences in manufacturing employment shares are not the primary source of variation. In a bivariate regression, the start-of-period manufacturing employment share explains less than 40%of the variation in ΔIP_{it}^{cu}. In all specifications, we control for the start-of-period manufacturing share within CZs so as to focus on variation in exposure to Chinese imports stemming from differences in industry mix within local manufacturing sectors.[16]

[16]For our sample period, the correlation between the CZ initial manufacturing employment share and the county-level change in the Republican two-party vote share is essentially zero (-0.01).

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 increased 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 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, Krishna, and Ma, 2015), dismantling of the constraints associated with central planning (Naughton, 2007; Hsieh and Song, 2015), and accession to the WTO (Pierce and Schott, 2016) have contributed to a massive increase in the country's manufacturing capacity and a concomitant rise in the country's manufacturing exports (Hsieh and Ossa, 2015).[17]

To identify the supply-driven component of Chinese imports, we instrument for growth in Chinese imports to the U.S. using the contemporaneous composition and growth of Chinese imports in eight other developed countries.[18] Specifically, we instrument the measured import-exposure variable ΔIP_{it}^{cu} with a non-U.S. exposure variable ΔIP_{it}^{co} that is constructed using data on industry-level growth of Chinese exports to other high-income markets:

$$\Delta IP_{it}^{co} = \sum_j \frac{L_{ijt-10}}{L_{uit-10}} \Delta IP_{j\tau}^{co}. \tag{2}$$

This expression for non-U.S. exposure to Chinese imports differs from the expression in equation (1) in two respects. In place of computing industry-level import penetration with U.S. imports by industry ($\Delta M_{j\tau}^{cu}$), it uses realized imports from China by other high-income markets ($\Delta M_{j\tau}^{co}$), and in place of start-of-period employment levels by industry and region, L_{ijt}/L_{it}, it uses employment levels from the prior decade in order to mitigate any simultane-

[17] See Bloom, Draca, and Van Reenen (2016) and Pierce and Schott (2016) for alternative instrumentation strategies.

[18] 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.

ity bias. As documented by Autor, Dorn and Hanson (2016), 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 manufacturing industries. Moreover, cross-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 diminishing trade and tariff costs in these surging sectors are a root cause.

A potential concern about our analysis is that we largely ignore U.S. exports to China, focusing instead on trade flows in the opposite direction. This is for the simple reason that our instrument, by construction, has less predictive power for U.S. exports to China. Nevertheless, to the extent that our instrument is valid, our estimates will correctly identify the direct and indirect effects of increased import competition from China. We note that imports from China are much larger—approximately five times as large—as manufacturing exports from the U.S. 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.

The exclusion restriction underlying our instrumentation strategy requires that the common component of import growth in the U.S. and in other high income countries derives from factors specific to China, associated with its rapidly evolving productivity and trade costs. Any correlation in product demand shocks across high income countries would represent a threat to our strategy, possibly contaminating both our OLS and IV estimates.[19] To check robustness against correlated demand shocks, Autor, Dorn, and Hanson (2013) develop an alternative estimation strategy based on the gravity model of trade. They regress China exports relative to U.S. exports to a common destination market on fixed effects for each importing country and for each industry. The time difference in residuals from this regression captures the percentage growth in imports from China due to changes in China's productivity and foreign trade costs *vis-a-vis* the U.S. By using China-U.S. relative exports, the gravity approach differences out import demand in the purchasing country, helping isolate supply

[19] Note that positive correlation in product demand shocks across high-income economies would make the impact of trade exposure on labor-market outcomes appear smaller than it truly is.

and trade-cost driven changes in China's exports.[20] These gravity-based estimation results are quite similar to the IV approach that we employ in this paper.[21]

Data on international trade for 1991 to 2010 are from the UN Comtrade Database, which gives bilateral imports for six-digit HS products.[22] To concord these data to four-digit SIC industries, we first apply the crosswalk in Pierce and Schott (2012), which assigns 10-digit HS products to four-digit SIC industries (at which level each HS product maps into a single SIC industry), and aggregate up to the level of six-digit HS products and four-digit SIC industries (at which level some HS products map into multiple SIC entries). To perform this aggregation, we use data on U.S. import values at the 10-digit HS level, averaged over 1995 to 2005.[23] All dollar amounts are inflated to dollar values in 2015 using the Personal Consumption Expenditure 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 A2 summarizes our trade exposure measures for county-by-district cells. On average, congressional districts saw an increase in Chinese import penetration by 0.71 percentage points between 2002 and 2010, and this rise was almost identical on average among districts that were won by Republicans and those won by Democrats in 2002. In the analysis that follows, 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 interquartile range is 0.49 percentage points across the full set of districts, and differs only modestly between districts won by Democrats and those won by

[20] The gravity-based approach addresses a second threat to identification, as well. It allows for the possibility that U.S., rather than Chinese, productivity shocks may be driving growth in U.S. imports from China. Because the gravity-model residuals summarize the change in China's comparative advantage relative to the U.S., the measure subsumes changes in U.S. productivity and thereby broadens the interpretation of the estimated coefficient from capturing the impact of supply shocks in China to capturing the impact of China-U.S. relative supply shocks. Despite this change in interpretation, China's much more rapid productivity growth makes it likely that its supply shocks, rather than those specific to the U.S., are the primary drivers of the country's export surge. Brandt, van Biesebroeck and Zhang (2012) estimate that over 1998 to 2007, China had average annual TFP growth in manufacturing of 8.0%, compared to the Bureau Labor Statistics' estimate of 3.9% for the United States (http://www.bls.gov/mfp/).

[21] See Autor, Dorn and Hanson (2013) and Autor, Dorn, Hanson and Song (2014) for further discussion of possible threats to identification using our instrumentation approach.

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

[23] The crosswalk assigns HS codes to all but a small number of SIC industries. We therefore slightly aggregate the four-digit SIC industries so that each of the resulting 397 manufacturing industries matches to at least one trade code, and none is immune to trade competition by construction.

Republicans in 2002 (0.53 and 0.49, respectively).

4 Main Results

We examine the political consequences of exposure to import competition from China in three stages, beginning with changes in party orientation (the likelihood that there is a change in party for a congressional district, and the change in district vote shares for each party), then considering changes in the ideological positioning of elected representatives (the nominal and absolute changes in Nominate scores for elected legislators, the likelihood that a more-liberal or a more-conservative legislator is elected), and finally evaluating how outcomes vary according to initial conditions in congressional districts (initially Republican versus initially Democratic districts, counties that in 2000 voted for George Bush versus those that voted for Albert Gore, and counties with a majority versus minority share of white voters). The first set of results allows us to examine the anti-incumbent effect and the party-realignment effect; the second and third sets of results allow us to assess the evidence for a political polarization effect and to characterize how this effect materializes.

In our primary specification, we estimate an equation of the form:

$$\Delta Y_{jkt} = \gamma_d + \beta_1 \Delta IP^{cu}_{jt} + X^{'}_{jkt}\beta_3 + Z^{'}_{jt}\beta_2 + e_{jkt}. \tag{3}$$

Here, the dependent variable ΔY_{jkt} is the change in electoral outcome between 2002 and 2010 that corresponds to county j and congressional district k. The main variable of interest is the contemporaneous change in import exposure ΔIP_{jt} in the commuting zone to which county j belongs. The first set of control variables X_{jkt} measure start-of-period political conditions corresponding to county-district cell jk. These controls include a dummy for the districts that elected a Republican legislator in 2002, the vote share of the winning party and a dummy for whether a candidate ran unopposed, each of again measured for district k in 2002 and interacted with a dummy for whether a Republican legislator won the 2002 election. Additional political controls include the 2003 to 2005 Nominate score of the winning candidate in 2002, also interacted with a dummy for Republican control.

Equation (3) further includes a vector of dummies γ_d corresponding to the Census geographic division to which each county j belongs, and a vector of control variables Z_{jt} measuring start-of-period economic conditions and demographic characteristics that pertain to the commuting zone encompassing county j. These include the share of manufacturing in CZ employment, the Autor and Dorn (2013) routine-task-intensity index and offshorability index for CZ occupations, the CZ population shares for nine age and four racial groups, and the shares of the CZ population that are female, college educated, foreign born, and Hispanic, where each of these variables is measured in the year 2000. All regressions are weighted by the 2002 share of county j in votes cast in congressional district k, which ensures that each district has equal weight in the analysis. If a congressional district spans multiple commuting zones, the weighting structure produces averages across these CZs, where weights are based on initial county electoral participation in the district. Standard errors are two-way clustered on the CZ and the congressional district. Following our strategy outlined above, equation (3) is estimated using two-stage least squares, with the import-exposure variable instrumented by contemporaneous changes in Chinese imports to other non-U.S. high-income countries as in (2).

4.1 Anti-Incumbent and Party Realignment Effects

Given the voluminous literature on the impact of economic conditions on incumbency, a logical starting point for our analysis is to assess how greater trade exposure impacts continuity in party control over a congressional district. In Table 1, we report estimation results for equation (3), in which the dependent variable is an indicator for whether there was a change in party controlling a district between 2002 and 2010. The first two columns show results, first from OLS and then from 2SLS models, with no additional control variables included in the estimation. In column (1), the OLS coefficient is positive but imprecisely estimated. When moving to 2SLS, the estimate increases in magnitude but remains statistically insignificant. The fact that the 2SLS point estimate exceeds its OLS counterpart is consistent with findings in Autor, Dorn, and Hanson (2013), showing that the exogenous component of rising China trade penetration generates substantially more negative local labor market impacts than does the observed (endogenous) trade measure, which likely comprises a mixture of Chinese supply shocks and domestic demand shocks. The column (2) estimate of 8.48 indicates that for two districts, one at the 25^{th} percentile of the increase in trade exposure (a 2002-2010 increase of 0.40 percentage points in import penetration) and another at the 75^{th} percentile of the increase in trade exposure (a 2002-2010 increase of 0.89 percentage points in import penetration), the more exposed district would be 4.2 $(8.48 \times (0.80 - 0.49))$ percentage points more likely to vote a new party in power between 2002 and 2010. Relative to the mean probability of a party change between 2002 and 2010 of 12.6 percentage points, this effect is non-trivial. The estimate falls slightly short of statistical significance at the 10% level, however, and precision falls further as we add the full set of covariates.

Table 1: Import Exposure and Congressional Election Outcomes 2002 - 2010. (Dependent Variable: 100 × Indicator for Change in Party)

	Change in Party, Election 2010 vs 2002							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Δ CZ Import Penetration	4.64	8.48	8.42	8.73	7.54 ~	8.40	8.73	7.71
	(2.89)	(5.40)	(5.38)	(5.38)	(4.00)	(9.09)	(8.37)	(8.26)
Estimation:	OLS	2SLS	2SLS	2SLS	2SLS	2SLS	2SLS	2SLS
F-statistic first stage		33.47 **	33.45 **	34.93 **	35.04 **	13.76 **	13.67 **	11.87 **
Control Variables:								
2002 Elected Party			yes	yes	yes	yes	yes	yes
2002 Election Controls				yes	yes	yes	yes	yes
2002 Nominate Controls					yes	yes	yes	yes
2000 Ind/Occ Controls						yes	yes	yes
2000 Demography Controls							yes	yes
Census Division Dummies								yes

N=3503 County*District cells. The states of AK, HI and VT are omitted, as well as country*district cells that did not exist in both 2002 and 2010 due to rezoning in TX and GA. The model in column 3 adds a dummy for the election of a Republican candidate in 2002. The additional election controls in column 4 comprise for vote share of the winning party and a dummy for elections that are unopposed, each interacted with the dummy for a Republican election victory. The Nominate controls in column 5 comprise the score of the 2002 election winner on the Nominate liberal-to-conservative scale, based on votes cast during the 2002-2004 congressional period, again interacted with a dummy for a Republican victory. Industry and occupation controls are measured at the CZ level and comprise the share of manufacturing in total employment (from 2000 County Business Patterns data), as well as the routine share and offshorability among occupations (based on 2000 Census data). Demographic controls comprise the percentage of a county's population in 9 age and 4 racial groups, as well as the population shares that are female, college-educated, foreign-born, and Hispanic. Census division dummies allow for different time trends across the 9 geographical Census divisions. Observations are weighted by a cell's fraction in total votes of its district in 2002, so that each district has an equal weight in the regression, and standard errors are two-way clustered on CZs and Congressional Districts. ~ $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$.

In columns (3) to (5), we include as regressors measures of electoral conditions in 2002 (indicator of initial winning party, vote share of that party, and indicator for unopposed election), and Nominate scores for the legislator elected in 2002. By interacting these variables with an indicator for the party elected in 2002, we flexibly control for the initial presence of a liberal, moderate, or conservative Democratic or Republican office holder, as well as any time trends in party vote shares or political polarization. The addition of these election variables modestly attenuates the impact of trade exposure on the likelihood of a change in the party in power, though the point estimate attains marginal significance in column (5). The remaining columns (6) to (8) of Table 1 add controls for initial economic conditions pertaining to a county-district cell, including CZ industry and occupation controls in 2000, CZ demographic controls in 2000, and dummies for the Census geographic division. These covariates do not greatly affect the point estimate but they substantially erode precision.

One reason for the imprecise anti-incumbent effect detected in Table 1 may be that the time period we consider spans two presidential elections, including one in 2008 in which a two-term incumbent president stepped down from office. Such transitions typically give a boost to the opposition party (Fair, 1996; Lewis-Black and Stegmaier, 2000), in this case the Democrats. Barack Obama's 2008 election victory created a coattail effect (Calvert and Ferejohn, 1983), which allowed the Democrats to retake the House of Representatives, after six terms in Republican hands. Following a coattail election, there is often regression to the mean, in which the party that loses House seats in a presidential-election year regains many of these seats in the first subsequent mid-term congressional election (Erikson, 1988). A mid-term correction occurred in 2010, which is the end year for our analysis. The 2002 to 2010 period thus includes some back-and-forth transitions in party control over House seats, which may add volatility to the dependent variable in Table 1.[24]

Column (1) of Table 2 provides an alternative measure of the impact of trade exposure on incumbency, in which we examine the change in the county-district vote share between 2002 and 2010 for the party that held the district in the initial year. The 2SLS regression includes the full set of political and economic controls used in column (8) of Table 1, which we take as our baseline specification. There is a negative and precisely estimated impact of trade exposure on vote shares for the party initially in power. The column (1) coefficient estimate of -6.98 (t-value 2.54) indicates that when comparing congressional districts at the 75^{th} versus 25^{th} percentiles of trade exposure, the more exposed district would have have a 3.4 (-6.98×0.49) percentage-point lower share of the 2010 vote going to the party that was in power in 2002, where the mean vote share of the winning party in 2010 is 62.1% and the

[24] Given that the time period for our analysis spans the Great Recession, it's also possible that any anti-incumbent effect of greater trade exposure is swamped by the effect of broader economic conditions.

mean 2002-2010 vote-share change is -8.48%. Congressional districts containing commuting zones subject to larger increases in import competition see a diminution of support for the party initially in office. This finding is consistent with the well-known result that voters punish parties that preside during bad economic times.

Table 2: Import Exposure and Congressional Election Outcomes 2002-2010. (Dependent Variables: Change in Percentage of Vote Obtained by 2002 Winning Party; Change in Percentage of Vote Obtained by Republican, Democrat, and Other Parties; 100 $\times$ Change in Probability that 2010 Race is Unopposed, or Is Won by $>$ 75% of Vote)

	Change in Voting Outcomes 2002-2010							
	% Vote for Party that Won in 2002		Republican Vote Share	Democrat Vote Share	Other Vote Share	Pr(Winner gets >75% of Vote)	Pr(Winner is Unopposed)	
	(1)		(2)	(3)	(4)	(5)	(6)	
Δ CZ Import Penetration	-6.98	*	1.97	0.31	-2.28	-13.49	-13.04	~
	(2.75)		(2.71)	(2.88)	(1.79)	(12.23)	(6.68)	

N=3503 County*District cells. All regression include the full set of control variables from Table 1. Observations are weighted by a cell's fraction in total votes of its district in 2002, so that each district has an equal weight in the regression, and standard errors are two-way clustered on CZs and Congressional Districts. ~ p ≤ 0.10, * p ≤ 0.05, ** p ≤ 0.01.

We next consider evidence for party-realignment effects, under which trade-induced changes in vote shares skew systematically in favor of one party or the other. In columns (2) to (4) of Table 2, we estimate the impact of an increase in CZ import competition on the change in the county-district vote shares of the Democratic Party, the Republican Party, and other parties. In all regressions, there is a null effect. Distinct from the findings in Che, Lu, Pierce, Schott, and Tao (2015), there is no generalized positive impact of trade exposure on the vote share for Democrats, and distinct from Dippel, Gold, and Heblich (2015) there is no across-the-board beneficial effect of trade exposure on voting for the most conservative party.[25] Increased exposure to China does not appear to foster a unidimensional realignment in favor of either major party or of third parties.

[25] Our results differ from Che, Lu, Pierce, Schott, and Tao (2015) in multiple respects. Most notably, we use county-by-congressional district cells as the unit of analysis, which allows us to examine who wins elections, whereas they focus on vote shares at the county level; and we examine the period 2002 to 2010, for which the definition of congressional districts remains constant, whereas they study the period 1998 to 2010, which spans congressional redistricting in 2002. To allow for closer comparison of our Table 2 results with their results, we also estimated the column (2) to column (4) regressions including state fixed effects, which in Che, Lu, Pierce, Schott, and Tao (2015) prove necessary to find a significant impact of import competition on party vote shares. When including state fixed effects, we continue to observe a null effect of trade exposure on the vote shares for either major party or for other parties.

The combination of results in column (1) of Table 2, which indicate that over the sample period trade exposure shifts votes away from the party initially in power in 2002, and of the results in columns (2) to (4), which reveal that no one party benefits disproportionately from changes in vote shares, suggests that greater import competition may have helped make congressional elections more competitive. We address this possibility directly in the final two columns of Table 2. Column (5) shows that in congressional districts subject to greater trade exposure in the 2000s, there is a sizable but statistically insignificant reduction in the probability that the winning candidate earns more than 75% of the vote (an outcome that occurred in 14.7% of districts in 2010). And in column (6), we see that districts subject to greater import competition become less likely to have candidates run unopposed (which occurred in 6.2% of districts in 2010), an impact that is significant at the 10% level. More competitive elections could potentially be the consequence of parties running more centrist candidates against each other who realize narrower electoral margins because they are competing for similar groups of voters. However, Figure 2 showed that during the outcome period, electoral victories actually narrowed as more politically extreme candidates gained office. The next set of results will show that greater trade exposure has contributed to this phenomenon by abetting the electoral ascendancy of legislators who hold more extreme positions, as revealed by their voting behavior. This move toward the extremes is an embodiment of political polarization.

4.2 The Political Polarization Effect

Table 3 presents our first results that document a direct impact of rising import competition from China on the widening partisan divide in the U.S. Congress. In column (1), the dependent variable is the 100 times the change in the Nominate score, which is increasing in the conservative positioning of a candidate based on her roll-call votes. In panel A, this 2002 to 2010 change compares the voting behavior of legislators elected in 2002, whose roll-call votes are observed in the 108^{th} Congress (2003 to 2005), against the voting behavior of legislators elected in 2010, whose roll-call votes are observed in the 112^{th} Congress (2011 to 2013). Greater trade exposure predicts an increase in the Nominate score in a district, indicating that on net districts subject to larger increase in import competition from China shift more strongly toward legislators who are further to the right on the political spectrum. Again comparing districts at the 75^{th} and 25^{th} percentiles of trade exposure, the more-exposed district would have an increase in the Nominate score that is 0.18 standard deviations higher.[26] From Table 3, we see that the trade-induced shift in favor of more conservative legislators does not arise because Republican candidates generally receive higher vote shares. It results, instead, from the election of more conservative Republicans.

Table 3: Import Exposure and Change in Political Position of Election Winner 2002-2010. (Dependent Variables: 100 $\times$ Change Nominate or Absolute Nominate Score of Winner)

	2002-2010 Change in Political Position				Decomposition of Change in Absolute Nominate Score			
	Nominate Score		Absolute Nominate Score		Shift to Right		Shift to Left	
	(1)		(2)		(3)		(4)	
	A. Between and Within Person Change of Nominate Score							
Δ CZ Import Penetration	18.41	*	14.15	*	10.69	*	3.46	
	(7.93)		(6.09)		(5.30)		(2.32)	
	B. Between Person Change of Nominate Score Only							
Δ CZ Import Penetration	20.13	**	15.61	**	12.14	*	3.47	
	(7.86)		(5.95)		(5.15)		(2.33)	

N=3503 County*District cells. The outcome in column 1 is the Nominate score times 100 (with negative values for liberals and positive values for conservatives), while the outcome in column 2 is the absolute value of that score, corresponding to the distance of a politician from the political center of the Nominate scale. Columns 3 and 4 decompose the change in absolute Nominate score into a shift to the right (higher Nominate values) and a shift to the left (lower Nominate values). Panel B replaces the Nominate scores of the 2010 election winners with their Nominate score from the 108th (2003-2005) congress or the first subsequent congress to which they were elected. This eliminates a within-person change in the Nominate score for districts that elected the same representative in 2002 and 2010. All regressions include the full set of control variables from Table 1. Observations are weighted by a cell's fraction in total votes of its district in 2002, so that each district has an equal weight in the regression, and standard errors are two-way clustered on CZs and Congressional Districts. ~ $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$.

The net positive impact of trade exposure on the Nominate score is not necessarily informative about *absolute* changes in the ideology of elected legislators. It could reflect a conservative shift among both Democratic and Republican representatives, with Democrats moving closer to the center and Republicans moving further to the right, or it could reflect movements away from the center in both parties, with Republican shifts being larger than those among Democrats. We address gross changes in Nominate scores in column (2) of

[26] For 2002 (the 108^{th} Congress), the standard deviation in 100 x Nominate score is 49, and the estimated interquartile effect is thus $(0.49 \times 18.41)/49 = 0.18$.

Table 3, in which the dependent variable is the change between 2002 and 2010 of the 100 times the absolute value of the distance between a legislator's Nominate score and the political center, which we take to be zero.[27] Under this metric, a one-unit shift to the right in the Nominate score is accorded the same value as a one-unit shift to the left. Column (2) reveals that greater trade exposure leads to a large and precisely estimated move away from the political center. Comparing districts at the 75^{th} versus 25^{th} percentile of trade exposure, the more-exposed district would see a relative increase of 0.36 ($0.49 \times 14.15/19$) standard deviations in its distance from the political center.[28] Columns (3) and (4) decompose the absolute change in Nominate scores, shown in column (2), into a rightward shift, defined as the absolute value of $\max\left[0, Nom_{d,2010} - Nom_{d,2002}\right]$, and a leftward shift, defined as the absolute value of $\min\left[0, Nom_{d,2010} - Nom_{d,2002}\right]$, where Nom_{dt} is the Nominate score for the legislator in district d in year t. Greater trade exposure induces a large and statistically significant rightward shift in the voting behavior of elected legislators, as seen in column (3), and a smaller and imprecisely estimated leftward shift, as seen in column (4). The point estimates suggest that about three quarters of the movement away from the political center induced by trade is the result of increasing conservativeness among elected legislators, while one quarter is due to increasing liberalness.

The positive effect of trade exposure on Nominate scores seen in the panel A regressions of Table 3 could be the result of sitting legislators changing their voting behavior to become more conservative over time or it could be the result of more-conservative legislators replacing less-conservative legislators during the 2000s. To resolve this ambiguity, in panel B we redefine the dependent variable to be the Nominate score for the winning legislator in 2010 as observed in the first Congress in which she served during the 2002 to 2010 window. If the same legislator remains in office during the entire period, the Nominate score is that for 2002, whereas if the legislator who won in 2010 was first elected in 2004, the Nominate score is the value for this earlier year, and so forth. The panel B regressions thus capture the impact of trade exposure on the *between-legislator* change in Nominate scores, as the outcome variable with have a value of zero for districts that elected the same representative in both 2002 and 2010. In all columns, coefficients in the panel B regressions—the between-legislator effect—are very

[27] As noted above, the mean Nominate score over all 113 congresses is very close to zero.

[28] For 2002, the standard deviation of 100 $\times$ the absolute value of distance from the political center (i.e., the absolute value of the Nominate score) is 19.

similar to those in the panel A regressions, which capture the between-plus-within-legislator effect. This reveals that changes in the Nominate score primarily result from the election of new legislators, rather than from changing voting patterns of incumbents.

Appendix Table A3 explores this pattern in greater detail by documenting how Nominate scores change within districts as representatives are variously reelected, replaced with same-party representatives, or displaced by members of the opposing party. Logically, when a house seat changes party hands, the Nominate score of the officeholder swings sharply (panel B), averaging +95 points for Democrat-to-Republican party changes, and −73 points for Republican-to-Democrat party changes (where reported Nominate scores are multiplied by 100). Republican-to-Republican swaps are also associated with substantial rightward movements in Nominate scores, averaging +15 points, while Democrat-to-Democrat swaps are associated with comparatively modest leftward movements, averaging −3 points. Among officeholders of either party who persist between 2002 and 2010, there is little change in observed ideology, though there is noticeably more rightward drift among Republicans incumbents (+6 points) than leftward drift among Democrat incumbents (−1.5 points).

Panel C of Appendix Table A3 documents the contribution of each margin of adjustment—party change, within-party representative change, and changes in Nominate scores among incumbents—to the overall within-district change in the Nominate score, accounting both for the probability of each of the six possible outcomes and the conditional mean change Nominate score change associated with each outcome.[29] Between 2002 and 2010, the average Nominate score change across the districts in our sample was 7.4, a sizable rightward shift. The two largest contributory factors to this shift are a net increase in the number of seats held by Republicans (contributing $6.8 - 3.8 = 3.0$ points by summing over the first two columns of Panel C) and the replacement of Republicans elected in 2002 with other Republican politicians who were more conservative on average (contributing 3.7 points). Countervailing against these trends was a small movement leftward following Democrat-to-Democrat swaps (−0.3 points). Voting behavior of incumbent Republicans and Democrats contributed to polarization as well, but its net effect was modest. On average, shifts in voting behaviors

[29]The contribution of each margin to the total is simply the fraction of districts falling into each category (panel A) multiplied by the mean change in the nominate score conditional on each outcome occurring. For example, Democrat-to-Republican transitions, which occur in 30 of 416 districts in our sample and are associated with a +94.75 swing in the Nominate score contribute $6.83 = 94.75 \times (30/416)$ points to the mean Nominate score change between 2002 and 2010 of 7.4 (obtained by summing the six entries in panel C).

of incumbent Republicans added 1.4 to Nominate scores while shifts in voting behavior of incumbent Democrats moved Nominate scores by -0.4 in the opposite direction.

Panels D and E of Appendix Table A3 carry out a parallel analysis for changes in the Republican percentage of the two-party vote between 2002 and 2010, averaged across Congressional districts. While it has to be the case that the Republican vote increases in districts that switch from Democrat to Republican control (column 1 of Panel D) and falls in districts that move from Republican to Democrat (column 2), it is not clear ex ante how vote shares change in districts that were won by the same party in both 2002 and 2010. Columns 3 and 4 of Panel D reveal that both the replacement of a Republican with another Republican and the replacement of a Democrat with another Democrat were associated with an average vote share loss by more than 10 percentage points for the successful party. Even in districts that reelected the same legislator in 2002 and 2010, the winning party's vote share fell over that period (columns 5 and 6 of Panel D). In districts that were held by the same party at the start and end of the outcome period, Nominate scores and vote shares thus moved in opposite directions: While Republican districts became more conservative and Democrat districts became more liberal (columns 3 to 6 of Panel B), this polarization of political positions was accompanied by more narrow election victories for the two parties. The districts that remained with the same party also account for a majority of the overall rightward shift in Nominate score, while they contribute little to the 1.3 percentage points average increase in Republican vote share across all districts.[30]

The movement away from the political center seen in Table 3 reflects the much-discussed demise of moderates in Congress (see, e.g., Layman, Carsey, and Horowitz, 2006). Table 4 examines the fortunes of centrists directly. The dependent variable in column (1) is the 2002-2010 change in an indicator for whether a "moderate" candidate is elected, where a moderate is defined to be a legislator whose Nominate score falls between the 20^{th} and 80^{th} percentiles of Nominate scores in the 107^{th} Congress (2001-2003).[31] Districts subject to larger increases in import competition from China are substantially less likely to elect a moderate

[30]By summing over columns 3 to 6 in Panels C and E, one obtains that districts held by the same party contributed 4.4 points to the 7.4 point average shift in Nominate score, while accounting for only 0.1 points of the 1.3 point gain in Republican vote share.

[31]This definition of moderates is, of course, somewhat arbitrary. Over all 113 congresses, the 20th and 80th percentile of Nominate scores (multiplied by 100) are -35.3 to 38.6. For the 107^{th} Congress (2001-2003), it is -38.9 to 59.0, indicating that many right-of-center legislators we are calling moderate would be decidedly conservative by historical standards.

legislator, an effect that is statistically significant (t-value 2.70). Comparing more and less trade-exposed districts, the more-exposed district would become 18.5 percentage points less likely to have a centrist in power between 2002 and 2010. To put this magnitude in context, over the 2002 to 2010 time period, the fraction of "moderates" in the House declines to 37.1% from a baseline of 56.8%.

Table 4: Import Exposure and Change in Political Position of Election Winner 2002-2010. (Dependent Variables: 100 $\times$ Change in Indicators for Election of Politician by Party and Political Position)

	Change in Probability 2002-2010 that Winner has Given Political Orientation									
	Moderate		Liberal Democrat	Moderate Democrat		Moderate Republican	Conserv-ative Republican		Tea Party Member	
	(1)		(2)	(3)		(4)	(5)		(6)	
Δ CZ Import Penetration	-37.66	**	0.27	-23.69	**	-13.97	37.38	**	24.44	~
	(13.95)		(7.11)	(8.72)		(9.58)	(14.04)		(12.77)	

N=3503 County*District cells."Liberal Democrats", "Moderates" and "Conservative Republicans" are defined as politicians whose Nominate scores would respectively put them into the bottom quintile, middle three quintiles, or top quintile of the Nominate score in the 107th (2001-2003) congress that preceeds the outcome period. A Tea Party Member is defined as a representative who was a member of the Tea Party or Liberty Caucus during the 112th (2011-2013) Congress. These two caucuses which are often associated with the Tea Party movement were first established in 2010 and 2011, respectively. All regressions include the full set of control variables from Table 1. Observations are weighted by a cell's fraction in total votes of its district in 2002, so that each district has an equal weight in the regression, and standard errors are two-way clustered on CZs and Congressional Districts. ~ $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$.

Subsequent columns of Table 4 examine how trade shocks reallocate House seats according to party and ideology. We examine the change in the likelihood of a district electing a legislator who positions herself as a liberal Democrat (column 2), a moderate Democrat (column 3), a moderate Republican (column 4), or a conservative Republican (column 5). In each regression, the dependent variable is the change over 2002 to 2010 in an indicator for whether a legislator of a particular type is elected. A "liberal" is a legislator whose Nominate score falls in the below the 20^{th} percentile for the 107^{th} Congress, a moderate continues to indicate a legislator whose Nominate score is in the 20^{th} to $80th$ percentile range of the 107^{th} Congress , and a "conservative" is a legislator whose Nominate score is above the 80^{th} percentile for the 107^{th} Congress. Because the four categories in columns (2) through (5) are exhaustive and mutually exclusive, the coefficients sum to zero across these columns and therefore indicate how trade exposure changes the political orientation of congressional districts.

Districts subject to greater import competition become substantially less likely to elect a moderate Democrat and substantially more likely to elect a conservative Republican, with both of these effects precisely estimated (t-values of 2.72 and 2.66, respectively). A more versus less trade-exposed district would become 11.7 (-23.69×0.49) percentage points less likely to have a moderate Democrat in power and 18.3 (35.24×0.49) percentage points more likely to be represented by a conservative Republican. Thus, a substantial fraction of the gains accruing to right-wing Republicans come at the expense of centrist Democrats. But this apparent rightward shift is far from monotone: trade exposure substantially reduces the electoral success of moderate Democrats but has no measurable effect on the prospects of liberal Democrats;[32] conversely, large gains among conservative Republicans in trade-exposed districts are accompanied by economically large, albeit imprecisely estimated, losses among moderate Republicans.

The regression model in column (6) of Table 3 additionally tests whether the trade shock affects the likelihood that a district elected a member who was or became affiliated with the congressional Tea Party Caucus or the Liberty Caucus. These two organizations, founded in 2010 and 2011, were the first congressional caucuses that have widely been characterized as being part of the Tea Party movement (Parker and Barreto, 2013). Tea Party membership thus provides us with an alternative outcome measure that captures the success of right-wing Republicans without relying on Nominate scores.[33] Consistent with the column (5) results for conservative Republicans, the column (6) estimate indicates that trade exposure raises the probability of the election of a Tea Party member to office, with an effect size of 12.0 (24.44×0.49) percentage points in a more versus less trade-exposed district (t=1.91).

[32]The fraction of districts represented by liberal Democrats increases from 19.9 percent in 2002 to 22.6 percent in 2010

[33]The average Nominate score of a Tea Party member in the 112th Congress is 0.80, which just slightly exceeds the average score of 0.78 for all legislators who are classified as conservative Republicans in Table 4. Members of the Tea Party and Liberty Caucuses make up over 40% of the conservative Republicans elected in 2010, though it is certainly possible that support of the Tea Party movement may extend beyond the formal members of these caucuses.

Table 5: Import Exposure and Change in Political Position of Election Winner 2002-2010: Alternative Specifications (Dependent Variable: 100 × Level or Change of Ideological Affiliation of Office-Holder)

	Liberal Democrat		Moderate Democrat		Moderate Republican		Conservative Republican	
	(1)		(2)		(3)		(4)	
			A. Outcomes in First Differences 2002-2010					
A. No Nominate 2002 Control	0.54		-24.06	*	-14.53		38.05	**
	(6.78)		(11.57)		(10.59)		(14.61)	
B. Linear Nominate	0.54		-24.08	*	-14.50		38.03	*
	(6.79)		(9.58)		(10.07)		(15.68)	
C. Quadratic Nominate	0.70		-24.28	**	-14.78		38.36	**
	(7.03)		(8.84)		(9.55)		(14.31)	
D. Linear Nominate x Party (Primary Spec)	0.27		-23.69	**	-13.97		37.38	**
	(7.11)		(8.72)		(9.58)		(14.04)	
E. Quadratic Nominate x Party	1.48		-23.54	**	-14.56		36.62	**
	(6.73)		(8.58)		(9.69)		(13.22)	
F. 4 Nominate Categories	4.63		-29.50	**	-8.39		33.26	*
	(6.31)		(8.95)		(7.59)		(13.55)	
G. Linear x Party + 4 Categories	7.45		-29.47	**	-7.93		29.95	*
	(5.14)		(8.87)		(7.49)		(11.73)	
			B. Outcomes in 2010 Levels					
A. No Nominate 2002 Controls	9.03		-32.55	**	-5.81		29.33	*
	(9.61)		(9.44)		(7.55)		(14.06)	
B. Linear Nominate	9.01		-32.54	**	-5.82		29.35	*
	(7.90)		(9.63)		(7.50)		(11.59)	
C. Quadratic Nominate	8.60		-32.19	**	-5.68		29.26	*
	(5.81)		(9.95)		(7.32)		(11.87)	
D. Linear Nominate x Party (Primary Spec)	9.78	~	-33.20	**	-6.12		29.53	*
	(5.63)		(10.17)		(7.35)		(11.94)	
E. Quadratic Nominate x Party	9.36	~	-31.43	**	-6.69		28.75	**
	(5.66)		(9.04)		(7.46)		(10.83)	
F. 4 Nominate Categories	4.63		-29.50	**	-8.39		33.26	*
	(6.31)		(8.95)		(7.59)		(13.55)	
G. Linear x Party + 4 Categories	7.45		-29.47	**	-7.93		29.95	*
	(5.14)		(8.87)		(7.49)		(11.73)	

N=3503 County*District cells. Classifications of candodate ideology is as in Table 4. Dependent variables: the change in ideological category of office-holder 2002 - 2010 (panel A) and category of office-holder in 2010 (panel B). All models include the full set of controls in Table 1 *except* the Nominate score of the 2002 office-holder, which is included as specified in each row headings. Observations are weighted by a cell's fraction in total votes of its district in 2002. Standard errors are two-way clustered on CZs and Congressional Districts. ~ $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$. ~ $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$.

The specification utilized in Table 4 has as the dependent variable the change in an electoral outcome for a congressional district (e.g., an indicator for whether a moderate Democrat was elected in 2010 minus an indicator for whether a moderate Democrat was elected in 2002). The controls include the initial Nominate score and its interaction with the party initially in power.[34] These specification choices are not derived from a specific model of electoral outcomes and, hence, we could have just as easily specified the dependent variable as the 2010 outcome, rather than the change in outcome, as a function of initial conditions. Similarly, we could potentially control for the initial level of the Nominate score in a variety of ways that differ from our linear-in-party primary specification.

In Table 5, we examine the sensitivity of the results in Table 4 to the choice of specification. Panel A reports regressions using the first difference in the outcome measure as the dependent variable (our baseline approach), whereas panel B reports regressions using the 2010 electoral result as the outcome. We also vary specifications according to whether we exclude the initial Nominate score from the control vector, include the initial score or its quadratic without interaction with the party initially in power, include the initial Nominate score interacted with the party initially in power (the baseline specification), or include the interaction of the party initially in power with the quadratic of the Nominate score or with four bin-size dummies for the Nominate score. These alternative specifications thus allow for varying assumptions regarding how initial conditions affect later outcomes and regarding the distribution of the error term. In all specifications, we continue to find a strong and significant negative impact of trade exposure on the likelihood of a moderate Democrat being elected in a district, and a more modest and statistically insignificant negative effect on the election success of moderate Republicans. The coefficient estimates for conservative Republicans and liberal Democrats are consistently positive, and are always significant for the former outcome while being marginally significant in some specifications for the latter. As a further sensitivity test, Appendix Table A4 re-estimates these models while classifying politicians' ideological affiliations based on the cardinal values of their Nominate scores rather than their percentile rankings in the empirical distribution.[35] These results are substantially identical to those in

[34]Since the initial Nominate score is used to categorize the ideological affiliation of the 2002 legislator, the 2002 Nominate score features in the construction of both the left and right-hand side variables.

[35]Appendix Table A4 classifies as moderates all legislators whose Nominate score is between -0.5 and +0.5 on the -1 to +1 scale. Under this alternative classification, most Democrats would be considered moderate while most Republicans would be considered non-moderate.

Table 4, again underscoring the robust impact of import competition on political polarization.

5 How and Where does Polarization Occur

Does polarization of officeholders occur because political parties change, because the identity of representatives changes across within-party transitions, or because incumbent representatives change their voting behavior? It may in principle be the case that greater trade exposure pushes incumbent politicians to become more conservative or more liberal in their voting in order to stave off political challengers on the right or left. However, our results across Panels (A) and (B) of Table 3 suggest that political realignment stems primarily from between-person rather than within-person changes in voting behavior, a pattern that is consistent with strong persistence in Nominate scores across time among sitting legislators (Poole and Rosenthal, 1997). Alternatively,trade shocks may induce turnover of politicians, with more extreme candidates replacing more centrist ones. Over the sample period, the potential for turnover may have been enhanced by the Bush-to-Obama transition in 2008. The strong coattail effect in this election, which resulted in many Democrats replacing Republicans, set the stage for a strong correction in 2010, in which Republicans regained many of the seats they lost in 2008, plus a good measure more.[36] A more-conservative Republican could thus have sought office in 2010, not by pursuing the generally unwelcome tactic of challenging a fellow Republican, but rather by taking on a Democrat who recently replaced a less-conservative GOP legislator.[37]

The regressions in Table 6 explore the electoral turnover channel by estimating the impact of trade exposure on transitions between parties and candidates. The dependent variable in the first column is an indicator for a change in party between 2002 and 2010. In the second column, it is an indicator for a new representative being elected in 2010 from the same party that won in 2002, and in the third column it is an indicator for the same representative being elected in 2002 and 2010. Because these outcomes are exhaustive and mutually exclusive,

[36]The Republican Party lost many House seats and control of the chamber in 2006, in an election that was distinguished by voter dissatisfaction with then President Bush and the unpopularity of the Iraq War.

[37]It is extremely rare for an incumbent to lose her Congressional seat due to a primary election defeat. . Of the 435 representatives elected to the House in 2002, only 1 percent had lost office cumulatively due to a primary election defeat by 2010. It is possible, however, that an embattled incumbent would step down prior to her party's primary election rather than risking defeat.

the coefficients sum to zero across the three columns in each panel. In Panel A, the column (1) result replicates the regression model in the final column of Table 1: trade exposure does not on average significantly increase the probability of a party transition. Panels B and C reveal that this net null effect aggregates across countervailing responses among initially Democratic and initially Republican districts. In initially Democratic districts, trade exposure substantially raises the probability of party turnover (standardized interquartile effect size of 14.5 percentage points), while in initially Republican districts, trade exposure substantially reduces the probability of a change in party (standardized effect size of −7.2 percentage points).[38] Both effects are economically large, though only modestly significant. Over the sample period, greater trade exposure contributed to the resulting stronger Republican majority in the House, allowing the Republicans to capture seats on net from the Democrats.

The positive effect of trade exposure on the ability of Republicans to pick up House seats might also translate into a higher likelihood of sitting Republicans retaining their seats. Strikingly, we find the opposite. More-trade-exposed districts that were initially in Republican hands are much more likely to have a *new* Republican in power in 2010. Comparing more and less trade-exposed districts, the column (2) results in panel C show that more-exposed districts initially held by Republicans are 21.7 ($40.96 \times (0.95 - 0.42)$) percentage points more likely to have a different Republican in office as of 2010 (on outcome that occurred in 47.6 percent of initially Republican districts). Turnover of elected legislators thus appears to be a key mechanism behind rise of more conservative Republicans in the House. Conversely, among initially Democratic districts, there is also a steep though imprecisely estimated trade-induced decline in the probability of an incumbent maintaining office (panel B). But this effect is equal and opposite to the impact of trade exposure on the probability of a change in the party of the officeholder. Thus, incumbent Democrats in trade-exposed districts are more likely to lose office to Republican challengers rather than being succeeded by other Democratic representatives.

[38]The level and dispersion of trade competition is very similar across districts that were initially held by Democrats and those initially held by Republicans (see Appendix Table A2). The indicated standardized outcomes multiply the impact coefficient by the interquartile range for initially Democrat districts of 0.49 in Panel B and for initially Republican districts of 0.53 in Panel C.

Table 6: Import Exposure and Congressional Election Outcomes 2002-2010. (Dependent Variables: 100 $\times$ Dummy for Change in Party, Change in Representative within Same Party, or No Change in Representative)

	Change in Party		No Change in Party			
			Different Rep		Same Rep	
	(1)		(2)		(3)	
			A. All Districts			
Δ CZ Import Penetration	7.71		15.94		-23.65	*
	(8.26)		(11.45)		(10.63)	
			B. Initially Democratic District			
Δ CZ Import Penetration	29.88	~	-5.21		-24.67	
	(17.82)		(18.05)		(18.37)	
			C. Initially Republican District			
Δ CZ Import Penetration	-13.95	~	41.12	*	-27.17	~
	(7.69)		(16.31)		(13.91)	

N=3,503 County*District cells in Panel A, N=1,234 in Panel B, N=2,269 in Panel C. All regression include the full set of control variables from Table 1. Observations are weighted by a cell's fraction in total votes of its district in 2002, so that each district has an equal weight in the regression, and standard errors are two-way clustered on CZs and Congressional Districts. ~ $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$.

The Table 6 results suggest that the impact of trade exposure on the ideological affiliation of officeholders may be asymmetric between initially Democratic and initially Republican districts. Voters in different districts may interpret news about trade-induced manufacturing decline in disparate ways, as suggested by the Dixit and Weibull (2007) model. In majority Republican districts, conservative newspapers or grass-roots organizations may reinforce beliefs that falling employment is the fault of government intervention or failed trade policy. Majority Democratic districts, in contrast, may see liberal media or advocacy groups convey the message that economic decline is the consequence of unfettered markets and argue for government intervention.

Table 7 tests for heterogeneous effects of trade exposure on the ideological affiliation of elected representatives by reestimating the Table 4 regressions separately for districts that were represented by a Democrat versus those represented by a Republican in 2002. Column (2) shows that in both initially-Democratic and initially-Republican districts, trade exposure makes the election of a moderate in 2010 much less probable, consistent with the results in

Table 4. Perhaps in part because initially Democratic districts are more likely to have a moderate legislator in power after the 2002 election, the negative impact of trade exposure on the likelihood of electing a moderate is larger in these districts. In these districts, losses by moderate Democrats accrue to both liberal Democrats and conservative Republicans. If we take two initially Democratic districts at opposing quartiles of trade exposure, the more-exposed district becomes 22.3 (-45.48×0.49) percentage points less likely to have a moderate Democrat in office in 2010, 7.7 (15.61×0.49) percentage points more likely to have a liberal Democrat in office, and 14.7 (30.07×0.49) percentage points more likely to have a conservative Republican in office. While the trade-induced decline in moderate Democrat officeholders is statistically significant (t-value 2.17), the offsetting gains among liberal Democrats and conservative Republicans are not individually significant. Summing over these margins, trade shocks in initially Democratic districts predict a net rightward shift in the ideology of office holders. This is also seen in the estimated positive impact of trade exposure on the Nominate score (column 1) and on the probability of electing a Tea Party-affiliated representative (column 7), though neither effect is precisely estimated.

Table 7: Import Exposure and Change in Political Position of Election Winner 2002-2010. (Dependent Variables: $100 \times$ Change in Nominate Score of Winner, $100 \times$ Change in Indicators for Election of Politician by Party and Political Position.)

		Change in Probability 2002-2010 that Winner has Given Political Orientation					
	Nominate Score	Moderate	Liberal Democrat	Moderate Democrat	Moderate Republican	Conserv-ative Republican	Tea Party Member
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
				A. Initially Democratic District			
Δ CZ Import Penetration	17.13	-45.67 *	15.61	-45.48 *	-0.19	30.07	31.46
	(15.06)	(21.04)	(18.85)	(18.96)	(6.61)	(19.14)	(23.35)
				B. Initially Republican District			
Δ CZ Import Penetration	12.19 ~	-34.64 *	0.00	-13.95 ~	-20.69	34.64 *	16.53
	(7.11)	(17.54)	.	(7.69)	(14.62)	(17.54)	(15.67)

N=1,234 County*District cells in Panel A, 2,269 County*District cells in Panel B. All regression include the full set of control variables from Table 1. Observations are weighted by a cell's fraction in total votes of its district in 2002, so that each district has an equal weight in the regression, and standard errors are two-way clustered on CZs and Congressional Districts. ~ $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$.

The shift away from moderates is the primary outcome that initially Democratic and initially Republican districts have in common. In initially Republican districts, the subsequent misfortune of centrist Democrats and Republicans however brings electoral gains only for conservative Republicans. Comparing two initially Republican districts, one at the 75^{th} percentile of trade exposure and the other at the 25^{th} percentile, the results in columns (2) and (6) of panel B in Table 7 indicate that the more-exposed district becomes 17.0 (-34.64×0.49) percentage points less likely to have a moderate in office (an outcome that occurred in 33.6% of initially Republican districts). The more-exposed district is in turn 17.0 percentage points more likely to have a conservative Republican in office in 2010 (an outcome that occurred in 66.0 percent of initially Republican districts), since there are no instances in which a liberal Democrat gained a seat in 2010 in a district that was Republican-held in 2002. Reflecting these electoral shifts, the average Nominate score rises in initially Republican districts. This rise is actually slightly smaller than the corresponding rise in initially Democrat districts, reflecting the fact that the latter had much more headroom to rise than the former.

One concern with the Table 7 results is that, because the samples are split according to the outcome of the 2002 house election, they may be partly driven by mean reversion. Appendix Table A5 addresses this concern by dividing counties across panels according to whether they voted majority Republican or majority Democrat in the 2000 presidential election. Using this pre-2002 sample split, the distinction between Republican-leaning and Democrat-leaning districts is at least equally stark. In trade-exposed locations that supported George W. Bush in the 2000 presidential election, the probability of a moderate of either party holding office falls steeply between 2002 and 2010, with a standardized effect size of -25.1 (-51.20×0.49) percentage points. These losses among moderates accrue in their entirety to gains among conservative Republicans, and Nominate scores of office-holders in Bush-supporting districts rise sharply. In counties that supported Albert Gore during the 2000 presidential election, trade exposure is estimated to reduce the probability that a seat goes to a moderate Democrat between 2002 and 2010. Here, offsetting gains accrue roughly evenly to liberal Democrats, moderate Republicans, and conservative Republicans, though none of these effects is precisely estimated.

Taken together, the results in Tables 6 and 7 suggest that the consequence of an increase in import competition from China is a substantial rightward shift accompanied by increased political polarization across local labor markets. Initially Republican districts move toward more conservative legislators; initially Democratic districts move towards both more liberal and more conservative legislators. The primary casualty of these political shifts are moderate Democratic politicians, with moderate Republicans paying a smaller though non-negligible electoral toll. This evidence suggests that it would be hard to back out the changes in political views represented in Congress from looking at changes in vote shares only.

Our analysis of the electoral consequences of rising trade exposure has so far only differentiated among congressional districts according to whether they were initially held by a Democrat or Republican representative in 2002, or voted majority Democratic or Republican in the 2000 presidential election. Table 8 offers a complementary perspective on political polarization by asking whether we observe divergent political consequences of trade exposure across counties subdivided along obvious demographic fault lines. Since the realignment of the two major parties following the passage of the Civil Rights Act in 1964, African-American voters have leaned strongly toward Democrats and southern whites have been equally supportive of Republicans, making race an important predictor of voting behavior. 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. Perhaps contrary to popular perception, the lion's share of U.S. county-district cells have a majority non-Hispanic white population: $3,241$ of $3,503$ cells, corresponding to 350 of the 416 electoral districts (84.1 percent) that are used in our analysis.[39]

This demographic split is, of course, correlated with the political affiliation of elected representatives: 59.2 percent of district-county cells with majority non-Hispanic white populations were represented by moderate or conservative Republicans in 2002; conversely, 79.0 percent of counties with a majority-minority voting-age populations were represented by moderate or liberal Democrats in 2002. The ideological affiliation of elected representatives from these districts are not wholly distinct, however. In 2002, 24.7 percent of representatives from majority non-Hispanic white counties were moderate Democrats, and 9.6 percent of representatives from majority-minority districts were moderate Republicans, with another 11.4 percent conservative Republicans. Our regression model statistically accounts for these initial differences, as above, by taking county-district level first differences in each outcome variable between 2002 and 2010—thus sweeping out a county-district effect—as well as by

[39]Our sample comprises $2,892$ counties, of which $2,719$ 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.

directly controlling for the party affiliation of the 2002 election winner, the vote shares and Nominate scores of incumbents, and the industry/occupation, race, gender, age, ethnicity, and nativity distribution of county residents.[40]

The upper panel of Table 8 shows that trade exposure catalyzed remarkably strong movements towards conservative Republicans between 2002 and 2010 in counties with majority non-Hispanic white populations. Scaling by the interquartile range of trade exposure, our point estimates imply that a more-versus-less-trade-exposed county would on average exhibit a 21.0 percentage point (42.87×0.49) increase in the probability that a conservative Republican takes office, with these gains coming at the expense of moderate Democrats (-13.3 points) and moderate Republicans (-7.7 points). The net effect of these movements is a very large rightward shift in the Nominate score in trade-impacted counties that have majority non-Hispanic white populations, with a standardized effect size of a $+10.3$ point increment to the Nominate score (t-value of 2.41).

Table 8: Import Exposure 2002-2010 and Political Position of 2010 Election Winner. (Dependent Variables: $100 \times$ Change in Nominate Score of Winner, $100 \times$ Change in Indicators for Election of Politician by Party and Political Position)

		Change in Probability 2002-2010 that Winner has Given Political Orientation				
	Nominate Score	Liberal Democrat	Moderate Democrat	Moderate Republican	Conservative Republican	Tea Party Member
	(1)	(2)	(3)	(4)	(5)	(6)
		A. Counties where >1/2 of Voting Age Pop is Non-Hispanic White				
Δ CZ Import Penetration	20.98 *	-0.01	-27.22 **	-15.64	42.87 **	25.28
	(8.69)	(7.94)	(9.88)	(11.61)	(16.17)	(15.47)
		B. Counties where <=1/2 of Voting Age Pop is Non-Hispanic White				
Δ CZ Import Penetration	-7.03	26.88 *	-25.33 *	11.36	-12.91	1.58
	(8.60)	(12.87)	(12.02)	(7.17)	(10.20)	(8.10)

N=3241 County*District cells covering 349.8 weighted districts in Panel A, N=262 County*District cells covering 66.5 districts in Panel B. All regression include the full set of control variables from Table 1. Observations are weighted by a cell's fraction in total votes of its district in 2002, so that each district has an equal weight in the regression, and standard errors are two-way clustered on CZs and Congressional Districts. ~ $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$.

[40]See the notes to Table 1 for additional details.

Focusing attention on the (much smaller) subset of counties with where less than half of the voting-age population is non-Hispanic white (panel B), we find a complementary pattern: liberal Democrats make strong gains in the probability of taking office, with a standardized effect size of 13.2 percentage points (t-value of 2.09). These gains come primarily at the expense of moderate Democrats, though conservative Republicans also appear to lose ground.

6 Discussion

The polarization of national politics has been one the defining events in American society of the last several decades. There is no longer any overlap in the supports of the distributions that describe the ideological positions of elected members of Congress from the two major political parties. 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. Yet, there is paucity of evidence that substantiates a causal impact of economic shocks on political polarization.

Clues for a connection between changes in the U.S. economy and the growing ideological divide in Congress come, fittingly enough, from the politicians themselves. In the 2016 U.S. presidential campaign, candidates from the extremes of both parties have singled out China as a principle cause for U.S. economic malaise. China bashing is now a popular pastime as much among liberal Democrats as among Tea Party Republicans. Our contribution in this paper is to show that this political showmanship is indicative of deeper truths. Growing import competition from China has contributed to the disappearance of moderate legislators in Congress, a shift in congressional voting toward ideological extremes, and net gains in the number of conservative Republican representatives, including those affiliated with the Tea Party movement.

It should perhaps come as no surprise that negative impacts of trade on U.S. manufacturing have engendered an intense political response. What is perhaps less expected is that the nature of this response appears to depend non-monotonically on the initial political orientation of a congressional district. In right-leaning districts, the beneficiaries are overwhelming Republicans from the right end of the spectrum, whereas in left-leaning districts support shifts from centrists both to the left and to the right. This divergence in which types of politicians gain from trade shocks is even more pronounced when separating districts according to initial racial composition. Voters are thus seeking answers to a common source of economic decline from very different types of political actors. 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 of belief formation wherein groups with common objectives but differing worldviews update their beliefs in opposite directions in the face of a common shock. While theory says that such polarization of beliefs should not persist indefinitely, convergence need not occur quickly or monotonically. What theory augurs for the convergence or further divergence of the U.S. political process over the short to medium term is thus highly uncertain.

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Appendix Tables

Table A1: Sample Selection: U.S. Congressional Districts

	No. Districts	% of Total
	(1)	(2)
Total Districts in U.S. Congress	435.0	100%
Excluded States	4.0	1%
AK	1.0	
HI	2.0	
VT	1.0	
Inconsistently Observed Cells	14.7	3%
TX	9.3	
GA	5.4	
Total Districts in Sample	416.3	96%

The sample excludes Alaska and Hawaii due to complications in definining Commuting Zones in those states, and Vermont, whose only district was represented by a congressman without party affiliation during the sample period. It also excludes county-district cells that are not continuously observed over time due to rezoning in the states of Texas and Georgia. The omitted areas correspond to about 1/3 of the districts in each of these states.

Table A2: Summary Statistics: Change in Commuting Zone Import Penetration, 2002 - 2010

	All Districts	District won by R in 2002	District won by D in 2002
	(1)	(2)	(3)
Mean	0.71	0.72	0.71
25th Percentile	0.40	0.42	0.40
Median	0.57	0.62	0.53
75th Percentile	0.89	0.95	0.89
P75 - P25	0.49	0.53	0.49

N=3503 district*county cells in column 1, N=2269 cells in districts that elected Republicans in the 2002 election in column 2, N=1234 cells in districts that elected Democrats in the 2002 election in column 3. Industry import penetration is the growth of annual imports from China 2002-2010, divided by an industry's U.S. domestic market volume in 1991. The Commuting Zone average of import penertration weights each industry according to its 2000 share in total Commuting Zone employment.

Table A3: Mean Changes in Nominate Scores ($\times 100$) by Change in Election Outcome 2002-2010, and Decomposition of Components to Total Observed Change

I. Party Change		II. Representative Change		III. No Change	
Democrat to Republican	Republican to Democrat	Republican to Republican	Democrat to Democrat	Republican Persists	Democrat Persists
(1)	(2)	(3)	(4)	(5)	(6)
A. Number of Districts					
30	22	104	42	95	123
B. Average Change in 100*Nominate Score by Type of District					
94.75	-72.51	14.89	-2.94	6.00	-1.50
C. Contribution to Overall Change in Average Nominate Score					
6.83	-3.80	3.73	-0.30	1.37	-0.44
D. Change in Republican Percentage of Two-Party Vote by Type of District					
29.61	-18.28	-10.31	10.88	-1.18	6.22
E. Contribution to Overall Change in Pct Republican Two-Party Vote					
2.14	-0.96	-2.58	1.10	-0.27	1.84

N=3503 County*District cells. Observations are weighted by a cell's fraction in total votes of its district in 2002. Values in Panel C sum to the average change in Nominate score of 7.39 for the whole sample. Values in Panel E sum to the average change in Republican two-party vote percentage of 1.27 for the whole sample.

Table A4: Import Exposure and Change in Political Position of Election Winner 2002-2010. (Dependent Variables: 100 × Change in Indicators for Election of Politician by Party and Political Position using Alternative Definition)

	Change in Probability 2002-2010 that Winner has Given Political Orientation									
	Moderate		Liberal Democrat		Moderate Democrat		Moderate Republican		Conserv-ative Republican	
	(1)		(2)		(3)		(4)		(5)	
Δ CZ Import Penetration	-29.83	**	-0.06		-23.35	**	-6.48		29.89	*
	(10.91)		(3.44)		(8.80)		(8.26)		(11.73)	

N=3503 County*District cells. The outcome in column 1 is the Nominate score times 100 (with negative values for liberals and positive values for conservatives), while the outcome in column 2 is the absolute value of that score, corresponding to the distance of a politician from the political center of the Nominate scale. "Liberal Democrates" and "Conservative Republicans" are defined as politicians whose Nominate scores are respectively below -0.5 or above +0.5. All regression include the full set of control variables from Table 1. Observations are weighted by a cell's fraction in total votes of its district in 2002, so that each district has an equal weight in the regression, and standard errors are two-way clustered on CZs and Congressional Districts. ~ p ≤ 0.10, * p ≤ 0.05, ** p ≤ 0.01.

Table A5: Import Exposure and Change in Political Position of Election Winner 2002-2010. (Dependent Variables: 100 × Change in Nominate Score of Winner, 100*Change in Indicators for Election of Politician by Party and Political Position)

			Change in Probability 2002-2010 that Winner has Given Political Orientation										
	Nominate Score		Liberal Democrat		Moderate Democrat		Moderate Republican		Conserv-ative Republican		Tea Party Member	
	(1)		(2)		(3)		(4)		(5)		(6)	
			A. County with Democratic Majority in 2000 Presidential Election									
Δ CZ Import Penetration	6.51		5.41		-14.97		4.12		5.44		11.40	*
	(11.72)		(15.64)		(17.66)		(11.07)		(14.21)		(5.70)	
			B. County with Republican Majority in 2000 Presidential Election									
Δ CZ Import Penetration	24.43	**	-1.06		-29.04	**	-22.16	~	52.26	**	28.03	
	(8.47)		(5.11)		(8.92)		(12.10)		(16.76)		(17.60)	

N=942 County*District cells in Panel A, N=2,561 County*District cells in Panel B. All regression include the full set of control variables from Table 1. Observations are weighted by a cell's fraction in total votes of its district in 2002, so that each district has an equal weight in the regression, and standard errors are two-way clustered on CZs and Congressional Districts. ~ p ≤ 0.10, * p ≤ 0.05, ** p ≤ 0.01.