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. We further contrast the electoral impacts of trade exposure with shocks associated with generalized changes in labor demand and with the post-2006 U.S. housing market collapse.

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1 Introduction

The 2016 U.S. presidential election has been as contentious as any in recent memory. The rancor on the campaign trail mirrors the partisan divide in Congress, which has been widening since the 1980s. DW-Nominate scores (Poole and Rosenthal, 1985 and 1991), which rank legislators on a liberal-conservative scale according to their roll-call votes, show that the ideological gap between the parties is at historic highs (McCarty, Poole, and Rosenthal, 2006).[1] 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 Republican legislators overlapped with that of more-conservative Democrats; today, even the most liberal House Republican is more conservative than the most conservative Democrat in the House.[2]

In this paper, we examine whether the exposure of local labor markets to increased foreign competition and other economic shocks has exacerbated partisan divisions in 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. To see whether the political repercussions of trade shocks differ from those of other shocks, we also examine the electoral impacts of overall industry-labor-demand shifts (e.g., Diamond, 2016) and the post-2006 housing-price collapse (e.g., Mian and Sufi, 2009 and 2011). Our interest is in seeing whether adverse shocks related to international trade or other events may cause voters to elect legislators who support positions that lean towards political extremes on the left or right.

The prominence of globalization in U.S. political debate is evident in the aggressive line 2016 presidential candidates have taken on trade. Republican party nominee Donald Trump pledged to impose a 45% tariff on U.S. imports from China,[3] while Democratic counterpart Hillary Clinton promised to block the proposed Trans-Pacific Partnership, which had been negotiated by a Democratic president.[4] Although leading politicians on the left and right share, at least for now, a dim view of trade agreements, the two major parties remain far apart on nearly every other major economic and social issue.[5] Voters thus face stark choices in selecting the legislators who best address their concerns.

[1]In the 1990s and especially the 2000s, greater polarization is also evident in the content of political speech, e.g., the frequency with which ideologically loaded phrasings such as the "death tax" or "tax breaks for the wealthy" are used by Democratic and Republican legislators (Gentzkow, Shapiro, and Taddy, 2016).

[2]Members of the two parties last overlapped in their congressional voting behavior in the 2001-2003 Congress (see Figure 3).

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

[4]See http://www.ontheissues.org/2016/Hillary_Clinton_Free_Trade.htm.

[5]See Dimock, Doherty, Kiley, and Oates (2014).

Voters' unease about international trade is backed by abundant evidence linking 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 professional occupations and low-wage service jobs (Autor and Dorn, 2013). 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 incomes for affected workers (Autor, Dorn, Hanson, and Song, 2014). The local labor markets that are home to more-exposed industries have endured substantial job loss and persistent increases in unemployment, non-participation in the labor force, and uptake of government transfers (Autor, Dorn, and Hanson, 2013). The political implications of manufacturing decline, whether related to trade or other forces, are not yet well understood.

To shed light on this issue, our analysis considers three mechanisms through which trade and other 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.[6] Voters punish incumbents at the polls for negative economic outcomes, including those caused by greater import competition (Margalit, 2011; Jensen, Quinn, and Weymouth, 2016). To align ourselves with the literature, 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. On their own, they imply that trade-exposed regions would simply alternate support between the major parties.

A second mechanism through which trade shocks may affect political cohesion is through a *realignment effect*. When individuals perceive declines in economic opportunity, their political preferences tend to shift in favor of redistribution (Alesina and La Ferrara, 2005; Bruner, Ross, and Washington, 2011; Giuliano and Spilimbergo, 2014). In the case of trade shocks, these patterns would predict greater support for Democrats, who tend to favor progressive taxation and generous social spending. Che, Lu, Pierce, Schott, and Tao (2015) find that over the period 1998 to 2010, U.S. counties exposed to greater import competition had larger increases in vote shares for Democratic candidates. As electoral outcomes are determined by votes at the congressional-district and not county level, such shifts do not necessarily translate into victories for Democratic legislators, however, nor do they indicate how legislators behave once in office. In Europe, the opposite voter response is observed. French and German regions that have been more exposed to trade with low-

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

wage countries have seen larger increases in vote shares for extreme-right parties (Malgouyres, 2014; Dippel, Gold, and Heblich, 2015), though none of these parties has become influential enough to form part of a governing coalition.[7] In our data, which are for congressional districts rather than counties, adverse trade shocks do not realign party vote shares.

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 non-centrist politicians. Holding constant political conditions in 2002—including the party in power, the vote share of the winning party, and the DW-Nominate score for the political orientation 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 candidates, 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, makes initially Democratic districts more likely to elect a liberal Democrat. The polarization effect is especially evident when separating locations by race rather than by initial 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.[8]

Our findings do not support the hypothesis that trade shocks make red states redder or blue states bluer, in the narrow sense of increasing the regional dominance of one major party or the other (Gelman, Park, Shor, Bafumi, and Cortina, 2008). Indeed, by helping more extreme legislators, greater trade exposure leads to fewer lopsided electoral victories, higher voter turnout, and larger individual campaign contributions, all of which indicate tighter races. These results highlight the value of studying the ideological positioning of winning candidates, as captured by DW-Nominate scores, rather than the vote shares that winning candidates garner.

To see whether the political impacts of import competition are distinct from other economic shifts, we incorporate a general Bartik (1991) shock into the analysis, which captures local changes in labor demand as predicted by national shifts in industry employment, and a measure of the change in local housing prices during the post-2006 housing bust (Mian, Sufi, and Trebbi, 2015; Palmer,

[7] During the Great Depression, European countries that had more prolonged downturns also saw greater support for far-right political movements (de Bromhead, Eichengreen, and O'Rourke, 2013).

[8] See Parker and Barreto (2013) on the role of race in the rise of the Tea Party movement.

2015). The Bartik shock, once decomposed into separate shocks for college and non-college workers (Diamond, 2016), helps capture any positive effects of industry export growth on the demand for high-skilled labor. Although predicted generalized contractions in local employment do marginally increase DW-Nominate scores (implying more conservative congressional representatives), these effects are considerably weaker than for trade shocks. Results on housing prices are broadly in line with our trade findings: local labor markets subject to larger post-2006 housing-price drops move away from moderate toward conservative legislators. However, these effects only hold for initially Republican districts.[9]

Why might trade shocks differentially affect political outcomes across regions? One answer is that voters have become more geographically segregated by ideology (Bishop, 2009). Yet, there is little evidence that geographic sorting of individuals according to their political beliefs has become more pronounced in recent decades (Glaeser and Ward, 2006; Ansolabehere, Rodden, and Snyder, 2008; Abrams and Fiorina, 2012). Alternatively, voters on the left and the right may have become more hardened in their beliefs about the world. Dimock, Doherty, Kiley, and Oates (2014) document that during the 2000s, the difference in views on a wide range of topics between those who identify or lean Democratic and those who identify or lean Republican grew substantially further apart.[10] 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, in which government intervention is helpful, or conservative, in which government intervention is harmful). These groups may respond to the same signal—a change in aggregate income—by updating beliefs in opposite directions, with liberals becoming more convinced the world is "liberal" and conservatives feeling similarly more confident the world is "conservative".[11] While such polarization may not last forever, convergence of the two groups to common posterior beliefs need not occur

[9] These findings are consistent with Mian, Sufi, and Trebbi (2014), who show that countries that experience financial crises, which are often related to troubles in housing markets, tend to see voter preferences shift in favor of more ideologically extreme politicians. See Mian, Sufi, and Trebbi (2010) on factors that shape congressional voting on housing-market legislation in the aftermath of the Great Recession.

[10] Grouping those who identify with or lean toward a party divides the large majority of independent voters into one party's camp or the other. Since the early 1990s, the Pew Center has surveyed eligible voters on a consistent set of topics including government regulation and spending, poverty, race relations, immigration, corporate profits, environmental laws, defense policy, and attitudes toward homosexuality. Whereas differences in views on these topics between Democratic and Republican leaners/identifiers were stable in the 1990s and early 2000s, they diverged sharply in the mid to late 2000s.

[11] The 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.

quickly nor monotonically.[12]

In the 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 district population share. This approach maps trade exposure at the CZ level to political outcomes at the congressional-district level, allowing us to examine how economic shocks affect not just party vote shares but which candidates win elections and how they behave once in office.

To measure regional trade exposure, we follow Autor, Dorn, Hanson, and Song (2014) and use the change in import penetration 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 both in the importance of manufacturing for their local economies and in their 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, Hanson, and Song (2014).

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 almost all 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).

[12] Acemoglu, Chernozhukov, and Yildiz (2015) show that with uncertainty about the distribution of the signal, convergence in beliefs need not occur at all.

Our paper provides the first evidence that connects adverse trade shocks to the demise of centrist politicians and an increase in political polarization. It contributes to a growing literature on political partisanship, which has not reached consensus on its origins (Fiorina and Abrams, 2008; Gentzkow, 2016). Bartels (2010) and Hacker and Pierson (2010) contend that changes in tax and regulatory policies in the 1970s and 1980s raised incomes for the wealthy, who then supported politicians favoring their interests. Although greater polarization in Congress does track rising income inequality (Barber and McCarty, 2015; Voorheis, McCarty, and Shor, 2016), the evidence linking these two phenomena is largely 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 cannot vote, thereby reducing support for redistribution. Popular treatments, including Frank (2004), look to culture, portraying elites as manipulating blue-collar workers to vote against their pocketbooks by diverting attention to abortion, gun rights, and gay marriage. Yet, Gelman, Park, Shor, Bafumi, and Cortina (2008) fail to find that immigration or cultural attitudes have increased partisanship.[13] There is also a business-cycle component to polarization. Mian, Sufi, and Trebbi (2014) find that over the 1879-2010 period DW-Nominate scores rise following banking crises and stock-market crashes, though these movements are not permanent.[14]

Other related work includes the many studies of how congressional representatives vote on trade legislation.[15] Specifically on the impact of import competition from China, Feigenbaum and Hall (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 well beyond U.S. trade policy initiatives and affect the overall ideological composition of Congress.

In section 2, we describe our data on congressional elections and voting patterns, and next summarize our data on local labor markets in section 3. In section 4, we present our main empirical results on the impacts of economic shocks on voting outcomes. Section 5 considers proximate mechanisms behind these impacts. Section 6 concludes.

[13] 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).

[14] Over the 1948-2008 period, they also find that financial crises are followed by increases in U.S. voter identification with more extreme ideological positions on the right and the left.

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

2 Measuring 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, as well as national shifts in industry labor demand, and the fallout of the housing-market collapse during the Great Recession. In combination, the full data allow us to analyze the impact of commuting-zone-level economic 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 the adult population in the district, such that each congressional district has equal weight in the analysis. If a district spans multiple CZs, the economic factors that are mapped to the district will be a population-share-weighted average of 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 must be omitted from the analysis.[16] The resulting set of 3,504 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

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

Elections.[17] 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 number and shares of votes won by Democratic and Republican congressional candidates in each county-by-district cell in 2002 and the change in these values between 2002 and 2010. We also note the number of registered voters and voter turnout by county, 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.[18] For additional measures of the competitiveness of congressional elections, we use the Database on Ideology, Money in Politics, and Elections (Bonica, 2013), which tabulates campaign contributions by donor and recipient for all amounts in excess of $200 using reports mandated by the Federal Electoral Commission.

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 analyses 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, each legislator has a well-behaved utility function defined over this space, and 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.[19]

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. history, DW-Nominate scores exhibited little gain in

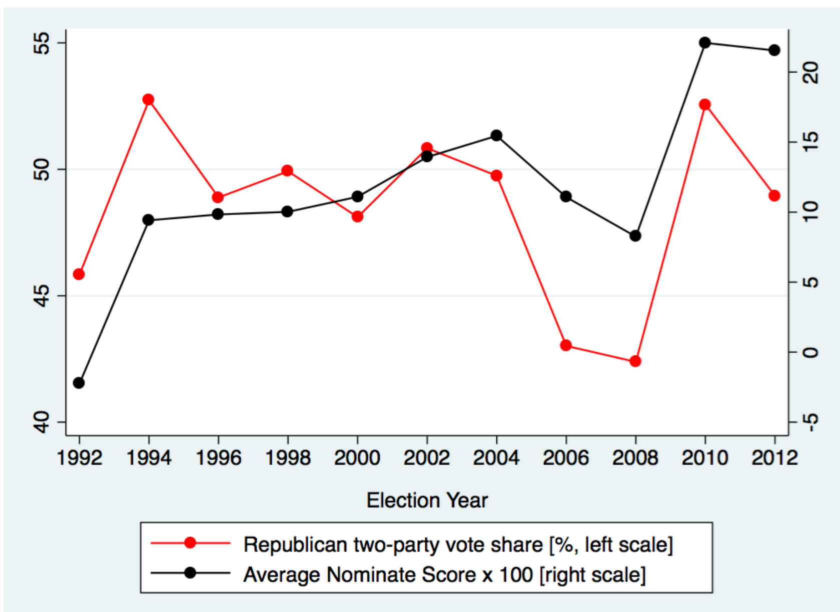
[17] See http://uselectionatlas.org/.

[18] Data on registered voters are missing in 2002 and/or 2010 in Georgia, Mississippi, North Dakota and Wisconsin. These four states are omitted from the corresponding part of the empirical analysis.

[19] 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. See Poole and Rosenthal (1997, 2001) and McCarty, Rosenthal, and Poole (2006).

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.[20] Henceforth, we refer to the first dimension of DW-Nominate as simply the Nominate score. For presentational purposes, we rescale the Nominate score by multiplying by 100 throughout the analysis.

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



Over the 113 U.S. congresses, the scaled Nominate score in the House is roughly centered on zero (mean= 2, standard deviation= 38), where the average value over time for each legislator is constrained to lie between 100 (most conservative) and -100 (most liberal). Figure 1 shows averages of this scaled Nominate score and the share of Republicans in the two-party vote for the 1992 to 2012 elections across all congressional districts.[21] 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

[20] The Gentzkow, Shapiro, and Taddy (2016) measure of speech polarization in Congress (see footnote 1) is an alternative indicator of the ideological positioning of congressional legislators. It is, however, not available for our entire sample period (nor it is available for all House members, as some legislators have insufficient speech recorded in the Congressional Record to estimate a score). For the time period when the Gentzkow, Shapiro, and Taddy measure overlaps with DW-Nominate scores, the two series are highly significantly correlated.

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

the Nominate scores we ascribe to 2002 are based on roll-call votes that occurred between January 2003 and January 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.

Average Nominate scores rise over time from −2 in 1992 for the 103th Congress (1993-1995) to 22 in both 2010 and 2012 for the 112th (2011-2013) and 113th (2013-2015) Congresses, a rise of two-thirds of a standard deviation relative to its distribution over all 113 Congresses.[22] 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 secular rise in Nominate scores 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.[23] Although the time trend in Nominate scores in Figure 1 is clearly positive, movements in the series are not monotonic. There are drops in 2006, a mid-term-election year in which voters punished Republicans for the unpopular Iraq War, and 2008, a presidential-election year in which many Democrats rode the coattails of Barack Obama to victory. The subsequent mid-term election in 2010 more than reversed these Democratic gains, as the rise of the Tea Party reinforced the post-1990 rightward shift in Congress. Such back-and-forth party gains and losses in congressional seats are a common pattern as the country moves through the cycle of mid-term and presidential elections (Calvert and Ferejohn, 1983; Erikson, 1988). Periodic swings in parties' seat totals create within-party turnover in legislators, a factor that will be important for understanding the sources of change in Nominate scores that the regression analysis uncovers.

In Figure 1, it appears that the average vote share for the Republican party and the average Nominate score track each other over time. The time-series co-movement in these means does not imply, however, that vote shares have increased for legislators who are especially conservative; higher Republican vote shares simply mean that Republicans are winning more elections. Figures 2 and 3 separate districts by whether the winning candidate was a Democrat or a Republican, the first in terms of vote shares and the second in terms of mean Nominate scores. Whereas mean Nominate

[22]The average Nominate score in our sample of 416 consistently observed districts increased from 13.9 in 2002 to 21.3 in 2010. These values closely track the average Nominate scores for the entire Congress (13.9 in 2002 and 21.7 in 2010).

[23]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).

scores for Republicans have risen over time (Figure 3), mean vote shares for winning Republicans have fallen during the 2000s (Figure 2), though of course they remain above 50%. Despite Republicans having nominated and gained electoral offices with more conservative candidates, the party's candidates have been winning with decreasing margins. The party's more conservative candidates may be putting off some voters, resulting in narrower victories. An alternative interpretation would be that Republicans have picked up seats in the House in more closely competitive districts, with correspondingly smaller victory margins. This logic would however also imply that Democrats should be winning their races by *larger* vote margins, which Figure 2 reveals is not the case.

Figure 2: Vote Shares by Winning Party

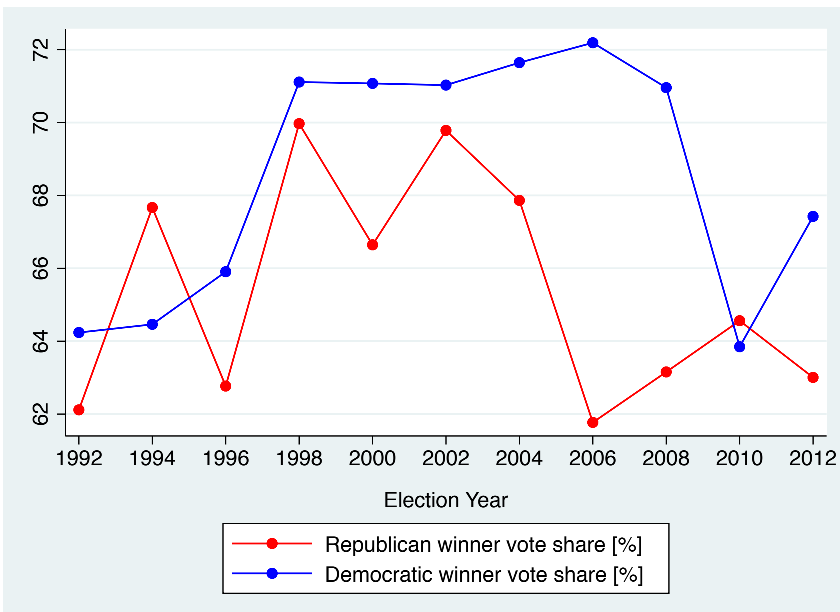
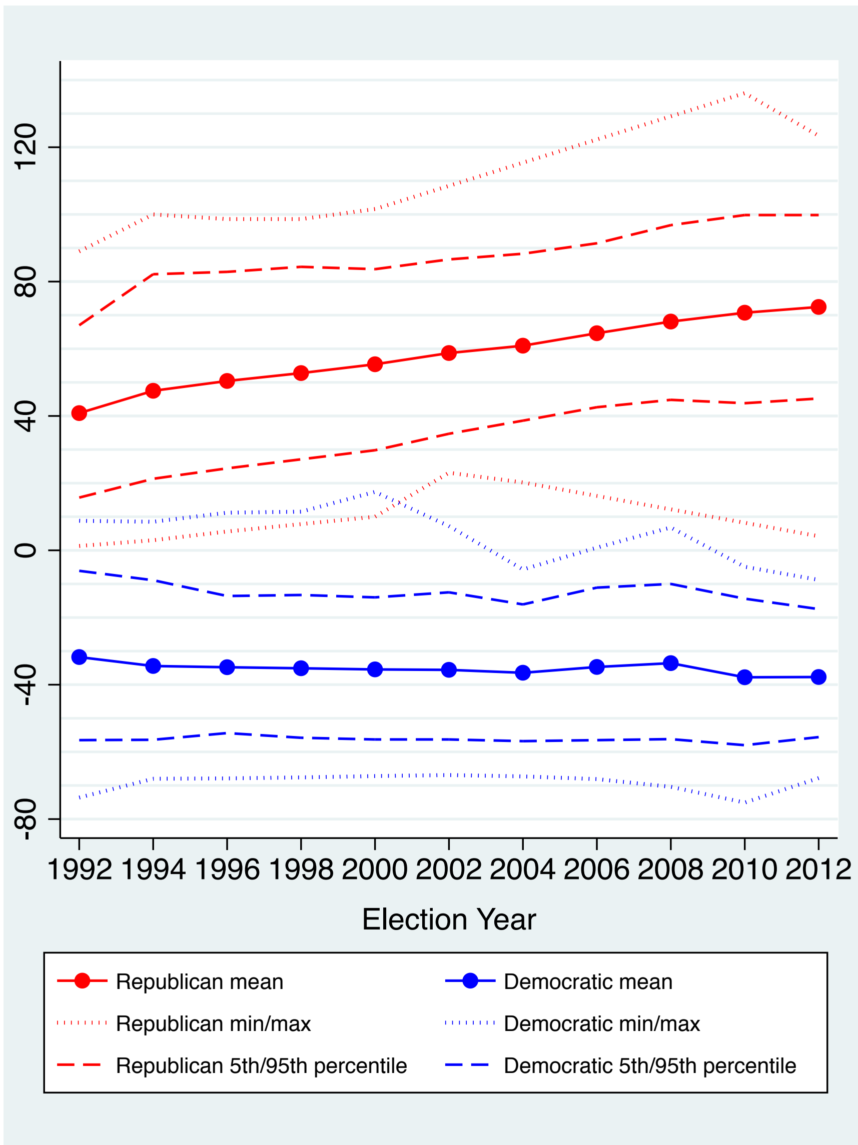


Figure 3 shows both the central tendency and the spread of Nominate scores for Democratic and Republican representatives, and starkly illustrates the widening partisan divide in Congress. Republican legislators have become markedly more conservative in their voting on legislation, with their average Nominate score rising from 41 in 1992 to 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 -32 to -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.

Figures 2 and 3 together show that in the 1990s there was polarization both in vote shares and Nominate scores: winning candidates won by larger margins and legislative voting of Democrats and Republicans moved in opposite ideological directions. In the 2000s, polarization is only evident in Nominate scores, a pattern that highlights the difficulty in using vote shares alone to study changes in partisanship.

Figure 3: Polarization in Nominate Scores



Given that congressional representatives vote on legislation that spans an immense array of subjects, one may wonder whether a scalar measure of ideology is capable of summarizing their voting behavior. To investigate this matter, we estimated Nominate scores from 2002 to 2010 (i.e., from the 108th Congress beginning in 2003 to the 112th Congress beginning in 2011) for seven major issue areas separately: budgetary issues, regulation, domestic social policy, defense and foreign policy, tariffs and trade regulation, immigration and naturalization, and globalization (trade and immigration combined).[24] Table A2 reports correlations across legislators between DW-Nominate scores (which cover all issue areas) and the seven issue-specific Nominate scores that we estimate. We report correlations for level values in 2002 and 2010 and for changes in values over 2002 to 2010. The correlations among Nominate scores in levels are above 0.92 for all issues in both years; the correlations among changes in Nominate scores for 2002 and 2010 are above 0.84 for all issues. This strong continuity across topics in voting and in changes in voting suggests that the scalar DW-Nominate measure of ideology provides a surprisingly comprehensive characterization of the legislative positioning of congressional representatives.

Our focus on how trade shocks affect electoral outcomes may seem to imply that voters choose representatives according to their stance on trade policy. Although trade has taken a central role in the 2016 presidential campaign, in the typical congressional session there are relatively few votes on trade-related topics. Of the 1,218 roll-call votes in the 108th Congress (2003-2005), just 23 were related to tariffs and trade regulation, and of the 1,602 roll-call votes in the 112th Congress (2011-2013), there were only 11 trade-related roll-call votes. Of course, the number of votes on a topic is not necessarily indicative of the prominence of the issue in voter's minds. However, truly consequential votes on trade policy—such as the passage of NAFTA or approval of China's accession to the World Trade Organization—occur perhaps a couple of times per decade.[25] The rarity of consequential votes on trade policy suggest that in selecting legislators, voters may have other issues foremost in mind. We thus view changes in Nominate scores as indicators of overall shifts in legislator ideology and not primarily in legislators' positions on trade.

[24] These issue areas are based on aggregate Peltzman codes in the Nominate data that identify the primary policy area a piece of legislation addresses: budget general and special interest (Peltzman codes 1, 2), regulation general and special interest (Peltzman codes 3, 4), domestic social policy (Peltzman code 5), defense and foreign policy budget and resolutions (Peltzman codes 61, 62, 71, 72), tariffs and trade regulation (issue code 50), and immigration/naturalization (issue code 59). For the results reported in Table A2, we estimate Nominate scores for each major issue area on its own. In this exercise, we exclude legislation on government organization, congressional rules, Indian affairs, and Washington, DC (see http://voteview.com/dw-nominate_textfile.html). To simplify the analysis for this validation exercise, we held Nominate scores constant across legislators over the 2002 to 2010 period and estimated the model using the R routine for W Nominate scores from http://voteview.com/wnominate_in_R.html. Because of the strong persistence over time in Nominate scores for individual legislators and the relatively short time period that we examine, this parameter homogeneity constraint (which applies only to the exercise in Table A2 and not to Nominate scores used in our main analysis) is not strong.

[25] The major pieces of trade legislation considered in the 108th and 112th Congresses were free-trade agreements with Central America, Colombia, and South Korea.

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:

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

In this expression, $\Delta IP^{cu}_{j\tau} = \Delta M^{cu}_{j\tau}/(Y_{j0}+M_{j0}-X_{j0})$ is the growth of Chinese import penetration in the U.S. for industry j over 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, $\Delta M^{cu}_{j\tau}$, 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 (1), the difference in ΔIP^{cu}_{it} 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 percent of the variation in ΔIP^{cu}_{it}. In all specifications, we control for the start-of-period manufacturing share within CZs so as to focus on variation in exposure to trade stemming from differences in industry mix within local manufacturing sectors.

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).

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.[26] 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 it replaces all other variables with lagged values to mitigate any simultaneity bias.[27] 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 of its manufacturing export growth.

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,

[26]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.

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

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 Bartik shock we describe below helps account for export-driven industry labor-demand shocks.

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.[28] 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 and trade-cost driven changes in China's exports.[29] These gravity-based estimation results are quite similar to the IV approach that we employ in this paper.[30]

Data on international trade for 2002 to 2010 are from the UN Comtrade Database, which gives bilateral imports for six-digit HS products.[31] To concord these data to four-digit SIC industries, we first apply the crosswalk in Pierce and Schott (2012), which assigns ten-digit HS products to four-digit SIC industries (at which level each HS product maps into a single SIC industry), and 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 ten-digit HS level, averaged over 1995 to 2005.[32] All dollar amounts are

[28]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 since these shocks would generate rising imports and rising domestic production simultaneously.

[29]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/).

[30]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, and see Bloom, Draca, and Van Reenen (2016) and Pierce and Schott (2016) for alternative instrumentation strategies for the change in industry import penetration.

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

[32]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,

inflated to dollar values in 2015 using the PCE deflator. Data on CZ employment by industry from the County Business Patterns for the years 1990 and 2000 is used to compute employment shares by industry in (1) and (2).

Table A3 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 Republicans in 2002 (0.54 and 0.49, respectively).

Greater exposure to import competition from China is far from the only recent shock to hit local labor markets. Employment shocks may also stem from changes in technology, consumer tastes, and regulatory policies, among myriad other factors. To construct an expansive measure of industry shocks to local labor markets, we use a Bartik (1991) type measure,

$$B_{it} = \sum_j \frac{L_{ijt}}{L_{it}} \frac{\Delta L_{j\tau}^{-i}}{L_{jt}^{-i}}, \tag{3}$$

where L_{ijt}/L_{it} is the share of industry j in the employment of CZ i in 2000 and $\Delta L_{j\tau}^{-i}/L_{jt}^{-i}$ is the relative growth of employment in industry j over 2000 to 2010 for all U.S. regions excluding CZ i.[33] This exposure index leverages the fact that because regions differ in their industry specialization patterns, and employment rises and falls unevenly across industries over time, CZs are differentially exposed to national changes in U.S. industry structure. To the extent that technological advance or expanded global trade are sources of change in industry labor demand, workers may be differentially affected by these shocks depending on whether or not they engage in skill-intensive tasks (Grossman and Rossi-Hansberg, 2008). A shock to an industry—such as the automation of routinized production (Autor and Dorn, 2013), China's accession to the WTO (Pierce and Schott, 2016), or changes in environmental regulations (Walker, 2013)—may increase demand for one skill group even as it decreases demand for others. In further analysis, we follow Diamond (2016) and construct (3) for two skill groups, college-educated workers (with at least some college) and non-college-educated workers

and none is immune by construction to trade competition.

[33]The Bartik shock is constructed based on civilian employment of individuals age 16-64 who do not live in institutional group quarters. The data are taken from the Census 2000 and the pooled ACS 2009-2011 (Ruggles, Genadek, Goeken, Grover, and Sobek, 2015), and converted from Public Use Micro Areas to commuting zones using the methodology described in Dorn (2009).

(with at most a high school degree), thus allowing the impacts of industry shocks to be skill-group specific.[34]

Our time period spans the Great Recession, which was triggered in part by the collapse of the U.S. housing market (Mian and Sufi, 2014). The literature has proposed various explanations for the housing boom and bust, including excessive growth in subprime mortgages (Mian and Sufi, 2009), over-leveraging by existing homeowners (Mian and Sufi, 2011), and heterogeneous expectations about housing prices (Burnside, Eichenbaum, and Rebelo, 2013), each of which may have been abetted by the securitization of mortgages, an abundant supply of global savings, and easily available credit (Keys, Mukherjee, Seru, and Vig, 2010; Bernanke, Bertault, DeMarco, and Kamin, 2011; Favara and Imbs, 2015). Whatever the underlying cause of recent U.S. housing troubles, a commonly used metric for local-area exposure to the housing bust is the change in housing prices from the peak of the market in early 2007 to the market trough in early 2011 (e.g., Mian, Sufi, and Trebbi, 2015; Palmer, 2015). Following the literature, we use the Federal Housing Finance Agency (FHFA) housing price index on repeat sales at the zip-code level to construct the average log change in housing prices at the CZ level between 2007q1 and 2011q1. The average change in nominal housing prices over this period is -19.7 log points, with a standard deviation of 19.6. Counties at the 90th percentile saw housing prices rise by 1.2 log points over the period, whereas counties at the 10th percentile saw housing prices drop by a stunning 45.8 log points.[35] Recognizing that changes in local housing prices may be partly the result of unobserved regional shocks, we do not include housing prices in our baseline specifications. As it turns out, the peak-to-trough change in housing prices at the CZ level is only weakly correlated with the change in CZ exposure to import penetration (correlation coefficient of 0.16). This modest positive correlation between trade exposure and house price changes is the result of more trade-exposed CZs having had smaller run-ups in housing prices during the mid 2000s and therefore smaller run-downs after 2006. Over the longer 1999-2011 interval, this correlation is equal to a de minimis -0.08. Consequently, including the change in housing prices in the analysis has little impact on our findings regarding import competition.

4 Main Results

We examine the political consequences of exposure to import competition from China and other shocks in three stages: beginning with changes in party orientation—the likelihood that there is a

[34]The construction of the skill group-specific labor demand shocks follows the same format as equation (3), except that all employment counts are either restricted to college-educated workers, or to workers without college education.

[35]These figures correspond to changes in house prices CZs in our sample of $3,504$ county-district cells. Where a CZ is included in multiple electoral districts, it contributes multiple observations to this calculation.

change in party for a congressional district, and the change in district vote shares for each party; next, 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 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:

$$\Delta Y_{jkt} = \gamma_d + \beta_1 \Delta IP^{cu}_{jt} + X^{'}_{jkt}\beta_3 + Z^{'}_{jkt}\beta_2 + e_{jkt}. \tag{4}$$

Here, the dependent variable ΔY_{jkt} is the change in an 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. In later specifications, we add the Bartik measure in (3) and the 2007-2011 change in local housing prices as shocks of interest. The first set of control variables X_{jkt} measure start-of-period political conditions corresponding to county-district cell jk. These include the vote share of the winning party in 2002, a dummy for whether the winning candidate ran unopposed in 2002, and the Nominate score of the winning candidate in 2002, each interacted with a dummy for whether a Republican legislator won the 2002 election.

Equation (4) further includes a vector of dummies γ_d corresponding to the Census geographic division to which county j belongs, and a vector of control variables Z_{jkt} measuring start-of-period economic conditions and demographic characteristics, either at the CZ or the county level. These include the share of manufacturing in CZ employment, the Autor and Dorn (2013) routine-task-intensity index and offshorability index for CZ occupations, county population shares for nine age and four racial groups, and the shares of the county population that are female, college educated, foreign born, and Hispanic, where each of these variables is measured in 2000.[36] All regressions are weighted by the 2002 share of county j in the adult population in congressional district k, which ensures that each district has equal weight in the analysis. If a congressional district spans multiple

[36] The manufacturing employment share is based on County Business Pattern data, the Autor-Dorn indices are based on occupational employment shares in the Census combined with task data from the Dictionary of Occupational Titles, and all population composition variables are based on Census short-form tabulations.

commuting zones, the weighting structure produces averages across these CZs, where weights are based on initial county population shares in the district. Standard errors are two-way clustered on the CZ and the congressional district. Following our strategy outlined above, we estimate (4) 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 (4), in which the dependent variable is an indicator for whether there was a change in the party controlling a district between 2002 and 2010. The first two columns show estimates, 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, likely because the latter comprises a mixture of Chinese supply shocks and domestic demand shocks. The column (2) point estimate of 8.78 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 0.90 percentage points in import penetration), the more exposed district would be 4.3 ($8.78 \times [0.899 - 0.405]$) percentage points more likely to vote a new party in power in 2010 compared to 2002. Relative to the mean probability of a party change between 2002 and 2010 of 12.6 percentage points, this magnitude 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	4.45	8.78	8.72	9.03 ~	7.71 ~	8.76	9.07	8.16
	(2.90)	(5.43)	(5.42)	(5.42)	(4.00)	(9.02)	(8.26)	(8.15)
Estimation:	OLS	2SLS	2SLS	2SLS	2SLS	2SLS	2SLS	2SLS
F-statistic first stage		38.01 **	38.03 **	39.32 **	39.65 **	15.12 **	13.67 **	13.09 **
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

Notes: N=3,504 County*District cells. The outcome variable has a mean of 12.45. The 2SLS models in columns 2-8 instrument for the change in Chinese import penetration in the US using the change in other developed countries' imports from China. The model in column 3 includes a dummy for the election of a Republican candidate in 2002. The additional 2002 election controls in column 4 are the vote share of the winning party and a dummy for unopposed elections, each interacted with the dummy for a Republican election victory. The Nominate controls in column 5 comprise the Nominate score of the 2002 election winner based on votes cast during the 2003-2005 Congressional period, again interacted with a dummy for a Republican victory. Industry and occupation controls in column 6 are measured at the CZ level and comprise the share of manufacturing in total employment (from the 2000 County Business Patterns data), as well as routine share and offshorability among occupations (based on Autor and Dorn (2013) and derived from 2000 Census data). Demographic controls in column 7 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 in column 8 allow for different time trends across the 9 geographical Census divisions. Observations are weighted by a cell's share of total district population in 2000, so that each district has equal weight in the regression. 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 columns (4) and (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, county demographic controls in 2000, and dummies for the Census geographic division. These covariates do not affect the point estimate but they do reduce precision.

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 × Change in Probability that 2010 Race is Unopposed, or Is Won by > 75% of Vote; Turnout in Opposed Races; Log Campaign Contributions)

A. Change in Voting Outcomes

	Vote % for Party that Won in 2002 (1)		Republican Two-Party Vote Share (2)		Republican Vote Share (3)		Democrat Vote Share (4)		Other Parties Vote Share (5)
Δ CZ Import Penetration	-7.05	**	0.18		1.60		0.86		-2.45
	(2.69)		(2.84)		(2.62)		(2.84)		(1.76)
Mean Outcome	-8.5		1.2		1.2		-1.3		0.1
Level in 2002	70.6		50.6		48.8		48.1		3.1

B. Change in Indicators of Competitiveness

	Opposition		Turnout in Opposed Races				Log Campaign Contributions		
	Pr(R+D Compete) (1)		100 × Ln(Votes) (2)		% of Registered Voters (3)		Individual Donors (4)		Corporate Donors (5)
Δ CZ Import Penetration	11.94	~	7.00	~	5.89	**	79.31	**	4.58
	(6.43)		(3.78)		(2.02)		(30.58)		(24.49)
Mean Outcome	12.3		13.8		3.1		86.0		111.1
Level in 2002	81.6		1079.1		47.2		602.6		610.8

Notes: N=3,504 County*District cells, except N=2,620 in column B2 and N=2,363 in column B3. The Republican two-party vote percentage in column A2 corresponds to the percentage of Republican votes in the total of Republican plus Democratic votes; the impact of the trade shock on the Democratic two-party vote share corresponds to the column A2 coefficient multiplied by minus one. Column B1 measures the likelihood that both a Republican and a Democratic candidate are competing in a district. Columns B2 and B3 measure turnout only for races in which candidates of several parties obtained votes, thus omitting unopposed races. Data on registered voters for column B3 is missing for the states of Georgia, Mississippi, North Dakota and Wisconsin. Columns B4 and B5 measure the log point change of one plus observed campaign contributions in $1,000. All regressions include the full set of control variables from Table 1. Observations are weighted by a cell's share of total district population in 2000, and standard errors are two-way clustered on CZs and Congressional Districts. ~ $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$.

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. The coattails of Barack Obama in 2008 helped the Democrats to retake the House, after six terms in Republican hands. Following a coattail election, there is often reversion to the mean, in which the party that loses seats in the presidential election regains seats in the subsequent mid-term congressional election (Erikson, 1988). A mid-term correction occurred in 2010, which is the end year for our analysis. The back-and-forth transitions of the 2002 to 2010 period may add volatility to the dependent variable in Table 1.[37]

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 -7.05 (t-value 2.63) 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 3.5 (-7.05×0.49) 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 mean 2002 – 2010 vote-share change is -8.5%. Congressional districts containing commuting zones subject to larger increases in import competition thus 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.

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 (5) of Table 2, we estimate the impact of an increase in CZ import competition on the change in the county-district shares of the Republican Party, out of two-party votes cast, and of the Democratic Party, the Republican Party, and other parties, out of all votes cast. 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 Malgouyres (2014) and Dippel, Gold, and Heblich (2015) there is no across-the-board beneficial effect of trade exposure on voting for the most conservative party.[38] Increased exposure to China does not appear to foster a unidimensional realignment in favor of either major party or of third parties.

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

[38]Our approach differs 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 makes our results representative at the district level and allows us to examine who wins elections, whereas Che et al. focus on vote shares at the county level that can pertain to different elections across multiple districts. Further, we examine the period 2002 to 2010, for which the definition of congressional districts remains constant, whereas Che et al. study the period 1998 to 2010, which spans intercensal congressional redistricting in 2002.

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 (5), which reveal that no one party benefits disproportionately from changes in vote shares, suggests that greater import competition may make congressional elections more competitive. We examine this possibility in the second panel of Table 2. Column (1) of panel B reveals that districts subject to greater import competition become more likely to have both major parties contest an election (which occurred in 81.6% of districts in 2002), an impact that is significant at the 10% level. In columns (2) and (3) of the second panel, we see that greater import competition produces an increase in votes cast (for which the mean was $48,533$ in 2002) and in the percentage of registered voters who cast ballots (which had a mean value of 47.2% in 2002) in a county-district cell, where the first effect is significant at the 10% level and the second at the 1% level.

In more-competitive races, the value of a marginal dollar of campaign spending may be higher. Consistent with this reasoning, column (4) of panel B shows that campaign contributions by individual donors grow by more in districts subject to larger increases in import competition, where the effect is precisely estimated. Comparing more and less trade-exposed districts, growth in log individual contributions is 39 log points higher in more-trade-exposed locations (where the mean log change in individual campaign contributions across districts over 2002 to 2010 is 86 log points). The same effect does not hold for contributions made by corporate entities, for which the estimated coefficient in column (5) of panel B is close to zero (in 2002 individual contributions account for an average of 48.0% of total contributions across districts). In the campaign-finance literature, individual contributions tend to account for a larger share of total contributions in closer political races (Ansolabehere, de Figueiredo, and Snyder, 2003), a finding with which our results are broadly consistent.

More competitive elections could potentially be the consequence of parties running more centrist candidates against each other who, because they are competing for similar groups of voters, realize narrower electoral margins. However, Figure 2 showed that during the period of study, electoral victories actually narrowed as more politically extreme candidates gained office. The next set of results will demonstrate 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 on the floor of the House. 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 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).[39] Greater trade exposure predicts an increase in the Nominate score in a district, indicating that on net, districts subject to larger increases 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.[40] 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 representatives.

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 Table 3, in which the dependent variable is the change between 2002 and 2010 of 100 times the absolute value of the distance between a legislator's Nominate score and the political center, which we take to be zero.[41] 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 ($= 0.49 \times 13.99/19$) standard deviations in its distance from the political center.[42] Columns (3) and (4) decompose the

[39] Nominate scores are missing for three representatives who were elected in 2010 but participated in very few votes during the 112^{th} Congress. For representatives Giffords (D-AZ) and Lee (R-NY), we measure political positioning based on their Nominate score for the 111^{th} Congress, and for representative Boehner (R-OH) who rarely voted during his first term as a speaker of the House, we use the average of his Nominate scores in the 111^{th} and 113^{th} Congresses.

[40] For 2002 (the 108^{th} Congress), the standard deviation of 100 x Nominate score is 49, and the estimated interquartile effect (based on an interquartile range of exposure of 0.49) is thus $(0.49 \times 18.13)/49 = 0.18$.

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

[42] For 2002, the standard deviation of 100 $\times$ the absolute value of distance from the political center (i.e., the absolute

absolute change in Nominate scores, shown in column (2), into a rightward shift, defined as the absolute value of max $[0, Nom_{d,2010} - Nom_{d,2002}]$, and a leftward shift, defined as the absolute value of min $[0, Nom_{d,2010} - Nom_{d,2002}]$, 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 less precisely 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, whereas one quarter is due to increasing liberalness.

Table 3: Import Exposure and Change in Ideological Position of Election Winner 2002-2010. (Dependent Variables: 100 × 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	18.13	*	13.99	*	10.83	*	3.16	
	(7.91)		(6.12)		(5.32)		(2.22)	
Mean Outcome	7.4		7.6		10.8		-3.2	
	B. Between Person Change of Nominate Score Only							
Δ CZ Import Penetration	19.69	*	15.30	**	12.17	*	3.13	
	(7.82)		(5.96)		(5.18)		(2.24)	
Mean Outcome	6.2		5.9		9.0		-3.0	

Notes: N=3,504 County*District cells. The outcome in column 1 is the Nominate score times 100 (with negative values for more liberal and positive values for more conservative views), 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 share of total district population in 2000, and standard errors are two-way clustered on CZs and Congressional Districts. ~ $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$.

Does the positive effect of trade exposure on Nominate scores seen in the panel A regressions of Table 3 reflect sitting legislators changing their voting behavior to become more conservative, or is

value of the Nominate score) is 19.

it due largely to more-conservative legislators replacing less-conservative legislators? To answer this question, panel B redefines the dependent variable as the difference in political positions between a district's representatives in 2002 and 2010, where each legislator's Nominate score for 2010 is replaced by the Nominate score 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 2010 Nominate score is replaced with 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 will have a value of zero for the 52% of districts in our sample 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 similar to those in the panel A regressions, which capture the between-plus-within-legislator effect. Thus, changes in the Nominate score stem primarily from the election of new, more ideologically conservative legislators, rather than from rightward movements in the voting patterns of incumbents.

Table A4 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. Over the 2002 to 2010 period, panel A shows that 30 House seats shift from a Democratic to a Republican legislator, whereas 22 seats change hands in the other direction. These numbers are dwarfed by within-party transitions. There are 42 seats that go from one Democrat to another and an impressive 104 seats that go from one Republican to another. Within-party transitions in legislators over 2002 to 2010 are primarily *not* the result of intra-party challenges, in which, say, a Tea Party candidate takes on a sitting Republican in a primary election. Turnover in legislators from such intra-party battles are exceedingly rare.[43] More common is a Republican-to-Republican transition to result from the 2002 incumbent losing a subsequent election to a Democrat with a fellow Republican winning the seat back in a later year. As mentioned, the 2006 and 2008 elections saw gains in the House by Democrats and 2010 saw Republicans more than recoup these gains. The resulting turnover in legislators creates the potential for substantial changes in ideological composition within parties over relatively short horizons.

Logically, when a House seat changes party hands, the Nominate score of the officeholder swings sharply, as seen in panel B of Table A4, averaging +94.8 points for Democrat-to-Republican party changes, and −72.5 points for Republican-to-Democrat party changes. Republican-to-Republican

[43] 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 risk defeat.

swaps are also associated with substantial rightward movements in Nominate scores, averaging +14.9 points, while Democrat-to-Democrat swaps are associated with comparatively modest leftward movements, averaging −2.9 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 (+6.0 points) than leftward drift among Democratic incumbents (−1.5 points).

Panel C of Table A4 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.[44] 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). The voting behavior of incumbent Republicans and Democrats contributed to polarization as well, but its net effect was modest. Shifts in voting of incumbent Republicans added 1.4 to the average Nominate score in Congress while shifts in voting behavior of incumbent Democrats moved Nominate scores by −0.4.

Panels D and E of Table A4 carry out a parallel analysis for changes in the Republican percentage of the two-party vote between 2002 and 2010, averaged across Congressional districts. Although 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. Notably, 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

[44] 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 7.39 (obtained by summing the entries in the first six columns of panel C).

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 1.2 percentage points average increase in Republican vote share across all districts.[45]

Table 4: Import Exposure and Change in Ideological Position of Election Winner 2002-2010. (Dependent Variables: 100 × 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	-35.96	**	0.17		-22.91	**	-13.04	35.79	**	24.30	~
	(13.35)		(7.01)		(8.56)		(9.02)	(13.54)		(12.65)	
Mean Outcome	-19.7		2.6		-4.6		-15.0	17.0		11.7	
Level in 2002	56.8		19.9		27.0		29.8	23.3		6.1	

Notes: N=3,504 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 share of total district population in 2000, 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 movement away from the political center seen in Table 3 reflects the much-discussed demise of moderates in Congress (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), which immediately precedes our sample period.[46] Districts subject to larger increases

[45]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 1.2 point gain in Republican vote share.

[46]This definition of moderates is, of course, somewhat arbitrary. Over all 113 congresses, the 20th and 80th percentile

in import competition from China are substantially less likely to elect a moderate legislator, an effect that is statistically significant (t-value 2.69). Comparing more and less trade-exposed districts along the interquartile range, the more-exposed district would become 17.6 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 from 56.8% to 37.1%.

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 categories in columns (2) to (5) are exhaustive and mutually exclusive, the coefficients sum to zero across columns and therefore indicate how trade exposure changes the ideological composition of Congress.

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 2.68 and 2.64, respectively). A more versus less trade-exposed district would become 11.2 (-22.91×0.49) percentage points less likely to have a moderate Democrat in power and 17.5 (35.79×0.49) percentage points more likely to be represented by a conservative Republican. Although a substantial fraction of the gains accruing to right-wing Republicans come at the expense of centrist Democrats, the rightward shift is not monotone. Trade exposure substantially reduces the electoral success of moderate Democrats but has no measurable effect on the prospects of liberal Democrats.[47] Conversely, gains among conservative Republicans in trade-exposed districts are accompanied by large, albeit imprecisely estimated, losses among moderate Republicans.

The regression in column (6) of Table 4 tests additionally whether the trade shock affects the likelihood that a district elects 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,

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. We examine below the robustness of our results to an alternative definition of moderate affiliation based on cardinal values of the Nominate score.

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

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.[48] 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 11.9 (24.30×0.49) percentage points in a more versus less trade-exposed district (t=1.92).

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.[49] These specification choices are not derived from a specific model of electoral outcomes and, hence, we could alternatively 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 A5, 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 allow for varying assumptions regarding how initial conditions affect later outcomes and 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

[48]The average Nominate score of a Tea Party member in the 112th Congress is 78, and happens to equal the average score of 78 for all legislators who are we 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 the effective support of the Tea Party extends beyond the formal membership of these caucuses. The Tea Party movement in Congress remains fluid. Although the Tea Party Caucus is now largely inactive, the Liberty Caucus remains active and the more recently formed House Freedom Caucus has grown equal in size (DeSilver, 2015).

[49]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.

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, Table A6 re-estimates these models while classifying politicians' ideological positions based on cardinal values of Nominate scores rather than percentile rankings in the empirical distribution.[50] The results are nearly identical to those in Table 4, underscoring the robust impact of import competition on political polarization.

4.3 Discussion: Economic Liberalism versus Anti-Globalization

It may seem paradoxical that trade shocks induce net gains in congressional seats for conservative Republicans, as these legislators belong to a party that since the 1950s has advocated liberalizing trade (Deslter, 1995). However, support for trade among Republicans, and especially among those in the right-wing of the party, is more equivocal than economic doctrine would imply. No less a conservative stalwart than Senator Barry Goldwater opposed the Trade Expansion Act of 1962, which granted the president the authority to negotiate tariff reductions as part of the Kennedy Round of the General Agreement on Trade and Tariffs.[51] Goldwater's opposition to GATT at the time is emblematic of a recurrent cognitive dissonance among conservatives regarding trade: conservative legislators tend to support free trade as a concept but to be skeptical of the international treaties that lower trade barriers. Equivocation on trade is also evident in U.S. legislative voting patterns. Milner and Tingley (2011) show that from the 1970s to the early 2000s, congressional representatives with higher Nominate scores (indicating a more conservative ideological position) have favored trade agreements in some congressional sessions but opposed them in others, with no evident time trend. In terms of public opinion, Republican voters are far from a uniformly pro-globalization bloc. A 2016 Pew Research Center survey of registered voters shows that those viewing free-trade agreements as having been a "bad thing for the U.S." included 53% of voters who identify or lean Republican and just 34% of voters who identify or lean Democrat.[52]

Globally, skepticism about trade is common among far-right political movements, including the Tea Party, which tend to have nationalist and populist tendencies (Judis, 2016). In Europe, successful right-wing populist parties—such as the Freedom Party of Austria, the Dutch Party for Freedom, the National Front in France, and the UK Independence Party (UKIP)—tend to be op-

[50] Table A6 classifies as moderates all legislators whose Nominate score is between −50 and +50 on the −100 to +100 scale. Under this alternative classification, most Democrats would be considered moderate while most Republicans would be considered non-moderate.

[51] See Patrick J. Buchanan, "Free Trade vs. the Republican Party," *American Conservative*, May, 20, 2016.

[52] Those who view free trade agreements as having been a "good thing for the U.S." include 38% of Republican identifiers/leaners and 56% of Democrat identifiers/leaners (Doherty et al., 2016).

posed to borders that are freely open either to immigration or to trade. Their support is stronger among the less educated, males, older cohorts, and members of ethnic majorities (Inglehart and Norris, 2016). UKIP supported Great Britain's exit from the European Union, as did many prominent figures of the country's mainstream Conservative Party, while in France the National Front argues for raising duties on manufacturing imports and renegotiating all EU treaties.[53] Put in this context, the anti-globalization rhetoric of Donald Trump's 2016 presidential campaign is in tune with modern expressions of right-wing populism. Voters choosing to support Republicans on the far right may therefore not be expecting to elect ardent free traders but rather politicians who take a less cosmopolitan view of globalization and exhibit skepticism of economic openness.

5 How and Where does Polarization Occur

Do the consequences of trade shocks for the ideological positioning of candidates reflect the particular impacts of trade on electoral outcomes or do they embody more-general consequences of adverse economic shocks on partisanship that heretofore have gone unnoticed? In this section, we explore whether the impacts of trade shocks depend on the initial party strength or demographic structure of congressional districts and incorporate into the analysis alternative measures of economic shocks related to generalized changes in industry labor demand (Bartik shocks) and the recent U.S. housing-market bust (the post-2006 change in local housing prices).

5.1 Sources of Heterogeneity in Impacts on Polarization

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 legislative-voting behavior. This pattern that is consistent with strong persistence in Nominate scores across time among sitting legislators (Poole and Rosenthal, 1997), while trade shocks 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. A more-conservative Republican could have sought office in 2010, not by pursuing

[53] See "France's National Front: On the March," *The Economist*, March 29, 2014.

the unwelcome tactic of challenging a fellow Republican, but rather by taking on a Democrat who recently replaced a less-conservative GOP legislator.

Table 5: Import Exposure and Congressional Election Outcomes 2002 - 2010. (Dependent Variables: 100 × 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	8.16		14.27		-22.43	*
	(8.15)		(11.12)		(10.29)	
Mean Outcome	12.4		35.1		52.4	
			B. Initially Democratic District			
Δ CZ Import Penetration	29.82	~	-6.36		-23.46	
	(17.81)		(18.26)		(17.97)	
Mean Outcome	15.4		21.6		63.1	
			C. Initially Republican District			
Δ CZ Import Penetration	-13.23	~	38.66	*	-25.43	~
	(7.32)		(15.38)		(13.34)	
Mean Outcome	9.9		47.2		43.0	

Notes: N=3,504 County*District cells in Panel A, N=1,233 in Panel B, N=2,271 in Panel C. All regressions include the full set of control variables from Table 1. Observations are weighted by a cell's share of total district population in 2000, 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 regressions in Table 5 investigate 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, as in the preceding Table 1. 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, 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. Continuing down column (1), 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 raises the probability of party turnover (standardized interquartile effect size of 14.6 percentage points), while in initially Republican districts, trade exposure reduces the probability of a change in party (standardized effect size of -6.5 percentage points).[54] Both effects are sizable, though only significant at the $p \leq 0.10$ level. 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 have translated 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 were 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 18.9 (38.66×0.49) percentage points more likely to have a different Republican in office as of 2010 (on outcome that occurred in 47.2 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 (column 3, 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 (column 1, panel B). Thus, incumbent Democrats in trade-exposed districts were more likely to lose office to Republican challengers rather than being succeeded by other Democratic representatives.

The Table 5 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. Survey data substantiate the perception that the divide on economy policy between those who identify or lean Democratic and those who identify or lean Republican widened considerably in the 2000s. Dimock, Doherty, Kiley, and Oates (2014) find that between 2004 and 2014 the gap between Republican identifiers/leaners and Democratic identifiers/leaners who agree with the statement "government regulation of business usually does more hard than good" grew from 7 percentage points (45% vs. 38%) to 39 percentage points (68% vs. 29%), with most of the change occurring by 2011. Although similar divergences

[54]As reported in section 3, 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 Table A3).

materialized on views regarding the wastefulness of government spending, the helpfulness of government assistance to the needy, and the consequences of immigration, there was no further divergence in views on military spending or the morality of homosexuality. Overall, whereas 49% of adults identified as moderates (i.e., neither mostly or consistently liberal or conservative) in both 1994 and 2004, this share fell to 39% by 2014. The 2000s, but not the 1990s, were thus a decade in which both voters and their elected representatives were vacating the middle of the political spectrum.

Table 6: Import Exposure and Change in Ideological 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.)

	Nominate Score		Change in Probability 2002-2010 that Winner has Given Political Orientation: Liberal Democrat	Moderate Democrat		Moderate Republican	Conserv-ative Republican		Tea Party Member
	(1)		(2)	(3)		(4)	(5)		(6)
			A. Initially Democratic District						
Δ CZ Import Penetration	16.93		15.30	-45.12	*	-0.26	30.07		31.18
	(14.96)		(18.59)	(18.61)		(6.81)	(19.24)		(23.64)
Mean Outcome	13.0		5.6	-21.0		3.6	11.8		5.4
Level in 2002	-36.4		42.4	57.6		0.0	0.0		0.0
			B. Initially Republican District						
Δ CZ Import Penetration	12.17	~	0.00	-13.23	~	-19.26	32.49	*	16.89
	(6.91)		.	(7.32)		(13.57)	(16.05)		(15.02)
Mean Outcome	2.5		0.0	9.9		-31.5	21.6		17.4
Level in 2002	58.3		0.0	0.0		56.1	43.9		11.6

Notes: N=1,233 County*District cells in Panel A, 2,271 County*District cells in Panel B. All regressions include the full set of control variables from Table 1. Observations are weighted by a cell's share of total district population in 2000, and standard errors are two-way clustered on CZs and Congressional Districts. ~ p ≤ 0.10, * p ≤ 0.05, ** p ≤ 0.01.

Table 6 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 initially 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. The negative impact of trade exposure on the likelihood of electing a moderate is particularly large in initially Democratic districts, where 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 22.1 (-45.12×0.49) percentage points less likely to have a moderate Democrat in office in 2010, 7.5 (15.30×0.49) 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 2.42), 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.

The shift away from moderates is the primary outcome that initially Democratic and initially Republican districts have in common. In initially Republican districts, the removal of centrist Democrats and Republicans brings electoral gains for conservative Republicans exclusively. 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 6 indicate that the more-exposed district becomes 15.9 (32.49×0.54) percentage points less likely to have a moderate in office (an outcome that occurred in 24.6 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, the more-exposed district is in turn 15.9 percentage points more likely to have a conservative Republican in office in 2010 (an outcome that occurred in 65.5 percent of initially Republican districts). 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 6 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. Table A7 addresses this concern by dividing counties across panels according to whether they voted majority Republican or majority Democrat in the 2000 presidential election. 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 separate along the Democratic-Republican divide. 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 either a Democratic or Republican moderate holding office falls steeply between 2002 and 2010, with a standardized effect size of -23.7 ($(-27.84 - 20.54) \times 0.49$) percentage points.

These losses among moderates accrue in their entirety to gains among conservative Republicans, with Nominate scores of office-holders in Bush-supporting districts rise sharply. In counties that supported Al 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 5 and 6 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 is moderate Democratic politicians, with moderate Republicans paying a smaller though non-negligible electoral toll. This evidence further underscores that it would be infeasible to accurately infer the changes in political views represented in Congress from changes in vote shares alone.

Our analysis of the electoral consequences of rising trade exposure has so far only differentiated among congressional districts according to initial party in power. Table 7 offers a complementary perspective on political polarization by testing whether we observe divergent political consequences of trade exposure across counties subdivided along 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.[55]

This demographic split is, not surprisingly, correlated with the political affiliation of elected representatives: 59.2 percent of district-county cells with majority 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 white counties were moderate Democrats, and

[55] 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.

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 directly controlling for initial political orientation, economic conditions, and demographic characteristics.

Table 7: Import Exposure 2002-2010 and Ideological Position of 2010 Election Winner. (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. Counties where >1/2 of Voting Age Pop is Non-Hispanic White					
Δ CZ Import Penetration	21.15 *	0.07	-26.90 **	-14.90	41.73 **	25.31 ~
	(8.63)	(7.86)	(9.65)	(10.78)	(15.37)	(15.31)
Mean	8.5	2.2	-4.2	-17.7	19.7	13.4
Level in 2002	20.1	16.1	24.8	33.5	25.7	6.3
	B. Counties where <=1/2 of Voting Age Pop is Non-Hispanic White					
Δ CZ Import Penetration	-8.28	25.66 *	-22.90 ~	9.99	-12.74	1.12
	(7.87)	(12.59)	(11.82)	(6.81)	(9.95)	(7.56)
Mean	1.8	5.0	-6.8	-1.8	3.5	3.4
Level in 2002	-17.7	39.3	38.4	10.9	11.4	5.3

Notes: N=3,241 County*District cells covering 347 districts in Panel A, N=263 County*District cells covering 69 districts in Panel B. All regressions include the full set of control variables from Table 1. Observations are weighted by a cell's share of total district population in 2000, 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 upper panel of Table 7 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 congressional district would on average exhibit a 20.4 percentage point (41.73×0.49) increase in the probability that a conservative Republican takes office, with these gains coming at the expense of moderate Democrats (-13.2 points) and moderate Republicans (-7.3 points). The net effect of these movements is a 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 +10.4 point increment to the Nominate score (t-value of 2.45).

Focusing attention on the (much smaller) subset of counties 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 12.6 percentage points (t-value of 2.04). These gains come primarily at the expense of moderate Democrats, though conservative Republicans also lose ground. For political cleavages identified along either initial-party or racial lines, the story is comparable. Trade shocks favor non-centrist politicians, with conservatives winning at the expense of moderates in initially Republican or white-majority districts and liberals benefiting from moderates' demise in initially Democratic or majority-minority districts.

5.2 Impacts of Other Economic Shocks on Political Polarization

In Table 8, we extend the analysis in Tables 5 and 6 by adding measures of two economic shocks to our baseline specification, the Bartik measure in (3) for the predicted change in CZ log employment based on national-industry employment changes, and the peak-to-trough log change in local housing prices during the post-2006 housing-market collapse. The first measure captures changes in national-industry economic conditions, inclusive and beyond trade with China, including the effects of technological advance (e.g., Autor and Dorn, 2013), globalization broadly defined (e.g., Ebenstein, Harrison, McMillan, and Phillips, 2014), and federal policies whose impacts may be sector-specific (e.g., Walker 2013). The second measure captures the differential exposure of local labor markets to the U.S. housing-market downturn associated with variation in the expansion of subprime lending (Mian and Sufi, 2009 and 2011; and Mian, Sufi, and Trebbi 2010), supply constraints on new housing construction (Glaeser, Gyourko, and Saiz, 2008), and other factors.

Comparing first coefficients on the change in import penetration in Table 8 with those in Tables 5 and 6, we see that results for the effect of import exposure on electoral outcomes are substantially the same. Whereas in initially Democratic districts greater trade exposure makes a change in party over 2002 to 2010 more likely (column 1, panel A), it has the opposite effect in initially Republican districts (column 1, panel B). And in both initially Democratic and initially Republican districts, changes in trade exposure reduce the likelihood that a moderate legislator is elected, especially when considering centrist Democrats (column 6), and increases the likelihood that a conservative Republican wins office (column 8). Patterns of statistical significance in the two tables are also similar. Precision rises for the estimated impact of import penetration on the change in likelihood that a conservative Republican is elected in an initially Democratic district (column 8, panel A), and falls modestly for estimated impacts of trade exposure on Nominate scores (column 4, panel B) and outcomes for moderate Democrats (column 6, panel B) and conservative Republicans (column 8, panel B) in initially Republican districts. As compared to Table 6, the first coefficient moves from insignificant to marginally significant (t-value 1.64) and the second three coefficients move from significance at the 5% to the 10% level. Overall, our results for the effect of exposure to trade shocks emanating from China on the ideological positioning of election winners are largely unchanged by introducing other economic shocks.

Consider next the electoral implications of exposure to the downturn in the U.S. housing market that accompanied the Great Recession. The results in Table 8 indicate that changes in housing prices are strongly related to changes in electoral outcomes, but only in initially Republican districts. In districts that were in GOP hands in 2002, decreases in housing prices during the post-2006 bust lead to a lower likelihood that there is a change in party (column 1, panel B), a larger increase in Nominate scores (column 4, panel B), a larger decrease in the likelihood that a moderate legislator is elected (columns 6 and 7, panel B), and a larger increase in the likelihood that a conservative Republican wins office (column 8, panel B). These impacts are consequential in their magnitudes. Similar to our standardization of trade shocks, consider two initially Republican districts, one at the 25th percentile of changes in housing prices (a 2007q1 to 2011q1 change of −28.5 log points) and one at the 75th percentile of changes in housing prices (a 2007q1 to 2011q1 change of −4.6 log points). In the district with the larger reduction in housing prices, there would be a 6.6 point larger increase in Nominate Scores and a 10.9 percentage-point larger reduction in the likelihood a moderate legislator is elected (against a mean decline of 21.6 percentage points), which implies an equal percentage-point increase in the likelihood that a conservative Republican gains power. Similar to trade shocks, housing-price shocks on net shift electoral outcomes against moderates in favor of conservatives.[56] But distinct from trade shocks, this outcome obtains only where Republican politicians were already in power.

[56]To facilitate comparison of the relative magnitudes of the different shocks, Table 8 reports the relative impact of each shock on a district at the 75th versus 25th percentile of exposure to the shock. The impacts of the Housing shocks on electoral outcomes in initially Republican districts are similar both in sign and magnitude to those of the trade shock.

Table 8: Import Exposure and Change in Ideological 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 of Party and Representative				Δ Prob that Winner has Given Political Orientation			
	Party Change	Same Party, Diff Rep	Party, Same Rep	Nominate Score	Liberal Dem	Mod Dem	Mod Repub	Conservative Repub
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
				A. Initially Democratic District				
Δ CZ Import Penetration	31.02 ~	-8.17	-22.86	18.59	14.72	-45.74 *	0.79	30.23 ~
	(17.20)	(16.68)	(18.15)	(14.54)	(17.84)	(18.65)	(5.76)	(17.75)
Δ CZ House Price Index	0.15	-0.16	0.00	0.22	-0.04	-0.11	0.17	-0.02
	(0.16)	(0.23)	(0.24)	(0.15)	(0.28)	(0.28)	(0.12)	(0.18)
Δ CZ Bartik Shift	0.16	-2.31 ~	2.15	-0.20	-0.98	0.82	-0.88	1.04
	(1.17)	(1.24)	(1.39)	(1.07)	(1.38)	(1.82)	(0.74)	(1.29)
Scaled Effects, P75 (More Adverse) vs P25 (Less Adverse) Shock								
Import Shock	15.3	-4.0	-11.3	9.2	7.3	-22.6	0.4	14.9
Housing Shock	-3.6	3.7	-0.1	-5.3	1.0	2.6	-4.0	0.4
Bartik Shock	-0.6	8.5	-7.9	0.8	3.6	-3.0	3.3	-3.8
				B. Initially Republican District				
Δ CZ Import Penetration	-12.85 ~	39.26 **	-26.42 *	11.63 ~	0.00	-12.85 ~	-19.58	32.43 ~
	(7.69)	(15.10)	(12.99)	(7.08)	.	(7.69)	(13.63)	(16.70)
Δ CZ House Price Shock	0.28 **	-0.55 *	0.27	-0.24 *	0.00	0.28 **	0.18	-0.46 **
	(0.10)	(0.23)	(0.23)	(0.11)	.	(0.10)	(0.16)	(0.16)
Δ CZ Bartik Shift	1.44 ~	0.58	-2.03	-1.78 *	0.00	1.44 ~	-0.51	-0.93
	(0.87)	(1.84)	(1.65)	(0.83)	.	(0.87)	(1.42)	(1.49)
Scaled Effects, P75 (More Adverse) vs P25 (Less Adverse) Shock								
Import Shock	-6.4	19.4	-13.1	5.8	0.0	-6.4	-9.7	16.0
Housing Shock	-6.6	13.0	-6.4	5.8	0.0	-6.6	-4.3	10.9
Bartik Shock	-5.3	-2.2	7.5	6.6	0.0	-5.3	1.9	3.4

Notes: N=1,233 County*District cells in Panel A, 2,271 County*District cells in Panel B. The table reports regression coefficients, and scaled effects that compare the implied effects for CZs at the 75th percentile of adverse shocks (greater import competition, greater house price decline, more negative Bartik shift) vs the 25th percentile. All regressions include the full set of control variables from Table 1. Observations are weighted by a cell's share of total district population in 2000, and standard errors are two-way clustered on CZs and Congressional Districts. ~ $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$.

A third set of results in Table 8 addresses the consequences of generalized labor-demand growth, as predicted by the Bartik measure in (3), on changes in the ideological positioning of winning congressional candidates. Districts subject to relatively smaller predicted growth in labor demand see a larger increase in Nominate scores, but again only in initially Republican strongholds: adverse shifts in labor demand have a small and statistically insignificant effect on Nominate scores in initially

Democratic areas (column 4, panel A) but a positive and precisely estimated in initially Republican locations (column 4, panel B). If we consider two initially Republican districts, one at the 25th percentile of predicted labor-demand growth and the other at the 75th percentile of predicted labor-demand growth, the first district would see a 6.1 point larger increase in Nominate scores, which is similar to the impact of the import shock.[57] Thus, adverse labor demand shifts—resulting either from import penetration or from broader changes in national-industry economic conditions—result in the election of more conservative legislators. Looking across columns (5) to (8), the evidence that generalized changes in labor demand favor or disfavor any particular type of legislator is not strong. Although in initially Republican districts larger declines in predicted employment tend to help conservative Republicans (as indicated by the negative coefficient in column (8), panel B) and hurt moderate Democrats (as indicated by the positive coefficient in column (6), panel B), only the latter effect is marginally statistically significant (t-value 1.66). We conclude that although the qualitative effects of trade shocks and predicted employment changes on electoral outcomes are similar in districts that initially elected Republicans, only changes in import penetration generate political polarization and well-defined changes in the fortunes of politicians across the political spectrum.

The implementation of the Bartik shock in (3) assumes that national-industry employment changes have a common effect across workers regardless of skill type. There is abundant evidence that since the 1980s, changes in industry labor demand have been considerably more favorable for more-skilled than for less-skilled labor (Katz and Autor, 1999). To allow for differential employment shifts by skill level within industries, we construct Bartik shocks that are specific to college-educated and non-college-educated individuals. Table A8 implements this approach. We replace the single Bartik shock in Table 8 with two shocks, one the predicated change in CZ log employment for more-educated workers and the other the predicted change in CZ log employment for less-educated workers. Although the results for changes in import penetration and changes in housing prices are largely the same as those in Table 8, estimated coefficients for the education-group-specific Bartik shocks in Table A8 show little evidence of consistent impacts on electoral outcomes. The only precisely estimated effect is that larger predicted decreases in employment of non-college-educated labor results in larger increases in the likelihood that a conservative Republican wins office in initially Democratic districts (column (8) of panel A).

[57] Aggregate Bartik shocks are positive for almost all CZs in this period, reflecting modest employment growth nationwide. Were we to focus instead on manufacturing-only Bartik shocks, these would be negative at every quantile. Not surprisingly, the manufacturing-specific Bartik shock is highly collinear with (and entirely swamped by) the import exposure measure, reflecting the centrality of trade shocks to the rapid decline in U.S. manufacturing employment in the 2000s (Autor, Dorn and Hanson, 2013).

6 Concluding Remarks

The polarization of national politics has been one of the defining events in American discourse 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. Our contribution in this paper is to show that this vitriolic campaign rhetoric 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 political actors with divergent ideologies. The paradox of converging popular beliefs about the source of economic challenges accompanied by diverging beliefs about appropriate policy 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.

Current expressions of voter anxiety substantiate our finding that the electoral consequences of trade are distinct from those of generalized changes in labor demand. What may distinguish trade in terms of its impact on political outcomes is that its disruption effects are so concentrated geographically. Whereas exposure to technological change in the labor market has affected both wealthy cities populated by white-collar professionals and factory towns home to blue-collar workers, rising import penetration from China and other low-wage countries disproportionately bears on local labor markets that historically specialized in labor-intensive manufacturing (Autor, Dorn, and Hanson, 2013b). This makes the employment consequences of trade acutely recognizable and therefore politically actionable. Exposure to the post-2006 U.S. housing market collapse has a similarly uneven geographic distribution. Correspondingly, areas with larger housing-price declines also embrace ideologically more-extreme legislators, though only those on the right and only in congressional districts that leaned right initially. The connection between economic and political polarization may thus arise not from sweeping changes in the U.S. economy that have left no corner untouched but rather from shocks whose disruptive force falls heavily on an identifiable set of voters who respond with equal vehemence at the polls.

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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	14.8	3%
TX	9.3	
GA	5.5	
Total Districts in Sample	416.2	96%

The sample excludes Alaska and Hawaii where the definition of Commuting Zones is difficult, and the at-large district of Vermont, which was the only district 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 district 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: Correlations between DW-Nominate Score and issue-specific W-Nominate Scores

	Issue-Specific W-Nominate Score						
	Budget	Regulation	Domestic Social Policy	Foreign Policy	Globali-zation	Tariffs and Trade Regulation	Immigr. and Naturali-zation
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
				A. Nominate Scores in 2002			
Corr. w/ DW-Nominate	0.990	0.966	0.983	0.982	0.963	0.926	0.947
				B. Nominate Scores in 2010			
Corr. w/ DW-Nominate	0.985	0.969	0.974	0.978	0.962	0.926	0.951
				C. Δ Nominate Scores 2002-2010			
Corr. w/ DW-Nominate	0.947	0.893	0.918	0.932	0.889	0.858	0.849

Notes: This table reports pairwise correlations between DW-Nominate scores and issue-specific W-Nominate scores across legislators in 2002 and 2010 and correlations across congressional districts for the 2002-2010 change in Nominate scores.. W-Nominate scores are computed for each legislator using rollcall votes specific to issue areas between 2002 and 2010. These issue areas are based on aggregate Peltzman codes in the Nominate data: budget general and special interest (Peltzman codes 1, 2), regulation general and special interest (Peltzman codes 3, 4), domestic social policy (Peltzman code 5), defense and foreign policy budget and resolutions (Petzman codes 61, 62, 71, 72), tariffs and trade regulation (issue code 50), and immigration/naturalization (issue code 59). Legislators with fewer than 20 rollcall votes within the issue area between 2002 and 2010 are excluded. In the first row, N=428, 427, 427, 427, 421, 383, and 381 legislators, respectively; in the second row, N=423, 423, 422, 423, 417, 283, and 273 legislators; in the third row, N=423, 422, 421, 422, 410, 250, and 239 districts. Observations are weighted by a cell's share of total district population in 2000.

Table A3: Summary Statistics: Commuting Zone Import, House Price and Bartik Shocks.

	All Districts	District won by D in 2002	District won by R in 2002
	(1)	(2)	(3)
	A. Δ CZ Chinese Import Penetration 2002-2010		
Mean	0.71	0.71	0.72
25th Percentile	0.40	0.40	0.42
75th Percentile	0.90	0.89	0.96
P75 - P25	0.49	0.49	0.54
	B. Δ CZ House Price Index 2007-2010		
Mean	-19.71	-20.55	-18.97
25th Percentile	-4.57	-7.23	-3.40
75th Percentile	-28.49	-28.49	-28.49
P75 - P25	-23.92	-21.27	-25.09
	C. Δ Bartik Labor Demand Shift 2000-2010		
Mean	6.83	7.35	6.38
25th Percentile	8.92	9.04	8.53
75th Percentile	5.23	5.55	4.66
P75 - P25	-3.69	-3.49	-3.87

N=3,504 district*county cells in column 1, N=1,233 cells in districts that elected Democrats in the 2002 election in column 2, N=2,271 cells in districts that elected Republicans in the 2002 election in column 3. The interquartile ranges compare CZs with more adverse labor demand shocks (75 percentile of greater import competition, greater house price decline, more negative Bartik shift) to CZs with less adverse shocks (25th percentile of these outcomes). Observations are weighted by a cell's share of total district population in 2000.

Table A4: 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		All Districts
(1)		(2)		(3)		(4)		(5)		(6)		(7)
A. Number of Districts												
30	+	22	+	104	+	42	+	95	+	123	=	416
B. Average Change in 100*Nominate Score by Type of District												
94.75		-72.52		14.89		-2.94		6.02		-1.49		7.39
C. Contribution to Overall Change in Average Nominate Score												
6.83	+	-3.80	+	3.73	+	-0.30	+	1.37	+	-0.44	=	7.39
D. Change in Republican Percentage of Two-Party Vote by Type of District												
29.55		-18.29		-10.34		10.87		-1.20		6.19		1.24
E. Contribution to Overall Change in Pct Republican Two-Party Vote												
2.13	+	-0.96	+	-2.59	+	1.10	+	-0.27	+	1.83	=	1.24

Notes: The table reports average changes in Nominate scores (Panel B) and Republican Two-Party Vote Shares (Panel D) separately for districts that changed parties from the 2002 to the 2010 election (columns 1-2), districts that changed representatives within parties (columns 3-4), and districts that elected the same represenative in both elections (columns 5-6). Panels C and E indicate the contribution of the districts with each type of election outcome to the overall changes in Nominate scores and vote shares as indicated in column 7 (in both panels, the sum of columns 1-6 sums to the value of column 7). These contributions correspond to the products of the number of districts with a given election outcome (Panel A), and the average change in Nominate scores and vote shares conditional on that outcome (Panels B and D).

Table A5: 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	0.59		-23.61	*	-13.48	36.49	**
	(6.62)		(11.54)		(9.92)	(14.13)	
B. Linear Nominate	0.53		-23.43	*	-13.74	36.65	*
	(6.65)		(9.46)		(9.62)	(15.37)	
C. Quadratic Nominate	0.61		-23.54	**	-13.89	36.83	**
	(6.91)		(8.69)		(9.04)	(13.89)	
D. Linear Nominate x Party (Primary Spec)	0.17		-22.91	**	-13.04	35.79	**
	(7.01)		(8.56)		(9.02)	(13.54)	
E. Quadratic Nominate x Party	1.45		-22.78	**	-13.67	34.99	**
	(6.61)		(8.42)		(9.17)	(12.69)	
F. 4 Nominate Categories	4.47		-28.69	**	-8.22	32.44	*
	(6.23)		(8.73)		(7.49)	(13.55)	
G. Linear x Party + 4 Categories	7.45		-28.79	**	-7.84	29.18	*
	(5.05)		(8.73)		(7.42)	(11.74)	
			B. Outcomes in 2010 Levels				
A. No Nominate 2002 Controls	8.63		-31.64	**	-5.98	29.00	*
	(9.56)		(9.15)		(7.61)	(14.31)	
B. Linear Nominate	8.79		-31.70	**	-5.84	28.75	*
	(7.75)		(9.35)		(7.46)	(11.66)	
C. Quadratic Nominate	8.57		-31.51	**	-5.76	28.70	*
	(5.70)		(9.78)		(7.29)	(11.95)	
D. Linear Nominate x Party (Primary Spec)	9.81	~	-32.56	**	-6.24	28.98	*
	(5.56)		(10.06)		(7.33)	(12.03)	
E. Quadratic Nominate x Party	9.36	~	-30.69	**	-6.84	28.17	**
	(5.58)		(8.87)		(7.46)	(10.92)	
F. 4 Nominate Categories	4.47		-28.69	**	-8.22	32.44	*
	(6.23)		(8.73)		(7.49)	(13.55)	
G. Linear x Party + 4 Categories	7.45		-28.79	**	-7.84	29.18	*
	(5.05)		(8.73)		(7.42)	(11.74)	

Notes: N=3,504 County*District cells. The ideological categories of office-holders are defined as in Table 4. 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 share of total district population in 2000, and standard errors are two-way clustered on CZs and Congressional Districts. ~ p ≤ 0.10, * p ≤ 0.05, ** p ≤ 0.01.

Table A6: 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	Conservative Republican	
	(1)		(2)	(3)		(4)	(5)	
Δ CZ Import Penetration	-30.50	**	-0.15	-22.60	**	-7.90	30.65	**
	(10.99)		(3.39)	(8.72)		(8.14)	(11.84)	
Mean Outcome	-12.4		1.4	-3.4		-9.0	11.0	
Level in 2002	57.9		6.5	40.5		17.4	35.6	

Notes: N=3,504 County*District cells. "Liberal Democrats" and "Conservative Republicans" are here defined as politicians whose Nominate scores (multiplied by 100) are respectively below -50 or above +50. All regressions include the full set of control variables from Table 1. Observations are weighted by a cell's share of total district population in 2000, and standard errors are two-way clustered on CZs and Congressional Districts. ~ p ≤ 0.10, * p ≤ 0.05, ** p ≤ 0.01.

Table A7: 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		Conservative Republican		Tea Party Member	
	(1)		(2)	(3)		(4)		(5)		(6)	
			A. County with Democratic Majority in 2000 Presidential Election								
Δ CZ Import Penetration	7.35		4.29	-14.80		5.15		5.36		10.17	~
	(11.67)		(15.60)	(17.67)		(11.32)		(14.45)		(5.68)	
Mean	0.0		6.3	-3.9		-10.8		8.4		3.9	
Level in 2002	-11.8		34.4	35.5		21.8		8.3		1.5	
			B. County with Republican Majority in 2000 Presidential Election								
Δ CZ Import Penetration	23.65	**	-0.97	-27.84	**	-20.54	~	49.36	**	28.07	
	(8.42)		(5.06)	(8.71)		(11.05)		(15.63)		(17.19)	
Mean	14.3		-0.8	-5.2		-19.0		25.0		19.1	
Level in 2002	37.9		6.4	19.1		37.2		37.3		10.4	

Notes: N=942 County*District cells in Panel A, N=2,562 County*District cells in Panel B. All regressions include the full set of control variables from Table 1. Observations are weighted by a cell's share of total district population in 2000, and standard errors are two-way clustered on CZs and Congressional Districts. ~ p ≤ 0.10, * p ≤ 0.05, ** p ≤ 0.01.

Table A8: Import Exposure, House Price Shock, Skill-Specific Bartik Shocks and Change in Political Position of Election Winner 2002-2010. (Dependent Variables: 100 × Indicator for Change of Party, Change of Representative within Party, and No Change; 100 × Change in Nominate Score of Winner; 100 × Change in Indicators for Election of Politician by Party and Political Position)

	Change of Party and Representative				Chg Prob that Winner has Given Political Orientation			
	Party Change	Same Party, Diff. Rep.	Party, Same Rep.	Nominate Score	Liberal Dem	Moderate Dem	Moderate Repub	Conserv-ative Repub
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
				A. Initially Democratic District				
Δ CZ Import Penetration	29.69 ~	-5.91	-23.78	17.17	15.41	-45.10 *	1.33	28.36 ~
	(17.02)	(16.42)	(18.04)	(14.43)	(17.93)	(18.79)	(6.29)	(17.29)
Δ CZ House Price Index	0.18	-0.22	0.04	0.26 ~	-0.06	-0.12	0.15	0.04
	(0.17)	(0.23)	(0.24)	(0.16)	(0.27)	(0.29)	(0.13)	(0.18)
Δ CZ Bartik Shift College Workers	-0.58	-0.17	0.75	-0.90	-0.19	0.77	-0.41	-0.17
	(1.22)	(1.23)	(1.57)	(1.20)	(1.67)	(1.98)	(0.67)	(1.22)
Δ CZ Bartik Shift Non-College Wkrs	1.41	-2.60 *	1.20	1.41	-0.82	-0.58	-0.76	2.17 *
	(1.19)	(1.16)	(1.46)	(1.07)	(1.03)	(1.42)	(0.74)	(1.08)
Scaled Effects, P75 (More Adverse) vs P25 (Less Adverse) Shock								
Import Shock	14.7	-2.9	-11.8	8.5	7.6	-22.3	0.7	14.0
Housing Shock	-4.4	5.4	-0.9	-6.2	1.5	2.9	-3.6	-0.8
Bartik College	1.8	0.5	-2.4	2.8	0.6	-2.4	1.3	0.5
Bartik Non-College	-6.2	11.6	-5.3	-6.3	3.6	2.6	3.4	-9.6
				B. Initially Republican District				
Δ CZ Import Penetration	-12.52	40.79 **	-28.26 *	11.53	0.00	-12.52	-21.77	34.29 *
	(7.68)	(14.88)	(12.79)	(7.14)	.	(7.68)	(13.69)	(16.85)
Δ CZ House Price Shock	0.29 **	-0.52 *	0.23	-0.26 *	0.00	0.29 **	0.14	-0.43 *
	(0.11)	(0.23)	(0.23)	(0.11)	.	(0.11)	(0.16)	(0.17)
Δ CZ Bartik Shift College Workers	0.66	-0.27	-0.39	-1.13	0.00	0.66	1.48	-2.14
	(0.91)	(1.88)	(1.77)	(0.90)	.	(0.91)	(1.36)	(1.36)
Δ CZ Bartik Shift Non-College Wkrs	0.90	1.49	-2.39	-0.90	0.00	0.90	-1.83	0.93
	(0.86)	(1.82)	(1.81)	(0.85)	.	(0.86)	(1.48)	(1.59)
Scaled Effects, P75 (More Adverse) vs P25 (Less Adverse) Shock								
Import Shock	-6.2	20.2	-14.0	5.7	0.0	-6.2	-10.8	17.0
Housing Shock	-7.0	12.4	-5.4	6.1	0.0	-7.0	-3.4	10.4
Bartik College	-2.1	0.9	1.2	3.6	0.0	-2.1	-4.7	6.7
Bartik Non-College	-4.0	-6.6	10.6	4.0	0.0	-4.0	8.1	-4.1

Notes: N=1,233 County*District cells in Panel A, 2,271 County*District cells in Panel B. The table reports regression coefficients, and scaled effects that compare the implied effects for CZs at the 75th percentile of adverse shocks (greater import competition, greater house price decline, more negative skill-specific Bartik shifts) vs the 25th percentile. All regressions include the full set of control variables from Table 1. Observations are weighted by a cell's share of total district population in 2000, and standard errors are two-way clustered on CZs and Congressional Districts. ~ $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$.