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Did US Politicians Expect the China Shock?†

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Information sets, expectations, and preferences of politicians are fundamental, but unobserved determinants of their policy choices. Employing repeated votes in the US House of Representatives on China's normal trade relations (NTR) status during the two decades straddling China's World Trade Organization (WTO) accession, we apply a moment inequality approach designed to deliver consistent estimates under weak informational assumptions on the information sets of members of Congress. This methodology offers a robust way to test hypotheses about what information politicians have at the time of their decision and to estimate the weight that constituents, ideology, and other factors have in policy making and voting. (*JEL* D72, D78, D83, D84, F14, P33)

The China shock, the large surge in imports from China that started in the 1990s and that has turned China into one of the United States' main trading partners, has received broad attention in academia and policy making. Although early research by Autor, Dorn, and Hanson (2013, 2016) and Pierce and Schott (2016) focused mainly of its labor market effects, a large literature has expanded the analysis to health, social, and political consequences of the shock.[1] While in hindsight China's entry in the US market may seem like a preordained outcome, in a series of roll call votes during the 1990s members of Congress were faced with the choice of allowing China to maintain its normal trade relations (NTR), and ultimately obtain permanent status (PNTR).

In this paper, we ask to what degree members of the US House of Representatives were informed about the consequences of the China shock for their voters and how much the expected impact on their constituents affected their support in favor or

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[1] Among the others, see Greenland and Lopresti (2016); Feler and Senses (2017); Autor, Dorn, and Hanson (2019); Greenland, Lopresti, and McHenry (2019); Autor et al. (2020) and Pierce and Schott (2020).

against China's NTR status.[2] These two questions are intrinsically related and point to the difficulty of modeling forward-looking expectations and decisions by policy makers in a context where these choices depend on consequences that are not known at the time of the vote. To this goal, we present a model and an estimation strategy of the decisions and expectations of law makers. In this, we depart from the empirical political economy literature on legislative voting (Poole and Rosenthal 1997; Heckman and Snyder 1997; Clinton, Jackman, and Rivers 2004; Canen, Kendall, and Trebbi 2020) and extend it in a distinct direction, as a formal analysis of law makers' information sets and expectations does not typically figure in standard empirical models of voting.

To provide intuition of why appropriate modeling of policy makers expectations is relevant, consider the following. A naïve approach to estimating the importance of constituent interests may be the replacement of the politician's expectations of the China shock with their realized values. However, assuming that politicians are perfectly informed about future shocks when they are not, necessarily implies a downward bias in the coefficient that measures the preference weight placed on constituent interests within a discrete choice voting model. This is due to an intuitive error-in-variables argument (i.e., the mechanical negative covariance between the expectational error and the realized future value of the shock). Without correction, a small coefficient may be interpreted as low responsiveness of politicians to subconstituents' fortunes, possibly indicating a political accountability problem (Kalt and Zupan 1984, 1990). In reality, a small coefficient may be as well the result of assuming that politicians are better informed than they truly are. Yet, this rather points to a limited expertise or insufficient information acquisition of legislators (Krehbiel 1992). Because these interpretations have distinct policy implications and they call for different remedies, it seems relevant to be able to distinguish between them.

To address the estimation challenge, we link the political economy literature to a separate strand of the international trade literature. The novel moment inequality methodology of Dickstein and Morales (2018), developed in the context of the decision of firms to export to foreign markets, allows us to consistently operate within an expectational environment where what belongs to the information set of the decision maker is only partially observed. That is, this moment inequality approach only requires the econometrician to know a subset of the information available to the politician at the time of his or her vote for the consistent estimation of the parameters—a much less demanding restriction. This approach turns out to be particularly informative in our context. It allows us to estimate a voting model under general assumptions about the politician's information set and expectations, and therefore to answer the question of how much politicians cared about the China shock in the first place.

Naturally, economic consequences on their constituents were not the only considerations affecting individual law makers' support for NTR and our model accommodates these features.[3] It is generally believed that several members of Congress

[2] The role of electoral constituencies and subconstituencies in driving the behavior of members of Congress has played a central role in the analysis of policy support in Washington, DC at least since Fenno (1978); Peltzman (1984). See also Mian, Sufi, and Trebbi (2010, 2014) for more recent applications.

[3] There is a vast literature discussing pure economic models of voting where electoral constituents (Peltzman 1984) or subconstitutents matter for roll call voting in Congress versus ideology of members of Congress (Kalt and Zupan 1984, 1990; Levitt 1996). For a recent review see (Mian, Sufi, and Trebbi 2014).

voted to withhold NTR status in order to affect China's position on human rights, as the series of yearly roll call votes on China's NTR status between 1990 and 2001 started after the Tiananmen Square events of 1989. Therefore, we also allow the voting behavior to depend on the ideological position of the legislator together the expected electoral cost of supporting China's NTR status, taking into consideration the district-specific impact of China's continued and growing exports to the United States. Further, in the utility function, ideology also captures the position of the legislator towards free trade policy, that is the value the politician places on the collective gains from maintaining low import tariffs.

We establish two main results. The first result is a moderate role of constituent interests. An interquartile difference in the value of the China shock decreases the probability of voting in favor of NTR for China by roughly 3 to 5 percentage points, while an analogous difference in ideology creates a 13to 17 percentage point increase in the probability of supporting NTR for China. In our heterogeneity analysis, we show that constituent interests are more important for Democrats than for Republicans, and for politicians that were elected with small vote margins (a margin of responsiveness supported by other studies; see discussion in Mian, Sufi, and Trebbi 2010; Ladewig 2010).

The second main result of our analysis is that politicians possessed a significant amount of knowledge about the future China shock. In some years, we cannot reject that they perfectly forecasted the shock that would hit their district in the next five years. More precisely, for all years from 1990 to 2001 we cannot reject that politicians had, at least, enough information to forecast 55 percent of the variation in the China shock. Perhaps surprisingly, our findings imply that knowledge decreases over the 1990s, a result that is plausible given that China's comparative advantage shifted substantially during the late 1990s and early 2000s. Comparing legislators across parties, we find that Democrats were systematically more informed than Republicans, with the exception of the early 1990s, a period when they are both equally informed.[4] We also present several validation exercises for our approach, including placebo exercises and a comparison of the NTR voting for China to the NTR voting for the case of Vietnam, showing how the moment inequality estimation highlights similar patterns in terms of information sets and preferences for comparable votes. Our findings on the extent of the information sets of US politicians (and their fairly accurate expectations) appear in line with the extent of information

[4] Anecdotally, the Congressional Record reports statements by members of Congress about the expected labor market consequences of the China shock which turned out to be fairly accurate. For instance, during the PNTR debate in 2000, Congress members David Bonior (D-Michigan) and Barbara Lee (D-California) shared predictions from the Economic Policy Institute, a nonprofit, for the state of California: "In my State of California we estimate 87,294 jobs lost …" over the next decade (https://www.congress.gov/congressional-record/volume-146/house-section/page/H3157). Similarly, but citing data from the International Trade Commission, Bill Pascrell (D-New Jersey) went on record stating, "In New Jersey, we will lose 23,000 jobs. In the United States as a whole, we will suffer a net job loss of 872,000 jobs over the same ten years. We are not creating jobs in America, we are creating jobs in China." (https://www.congress.gov/congressional-record/volume-146/house-section/page/H3514). Members of Congress acquired the information, among others, from labor unions (and especially so members of the Democratic Party). As Baldwin and Magee (2000, p.83) state, around the debate over NAFTA and China NTR, "...most labor unions were convinced that the adverse employment effects would be much more widespread than economists had predicted..." and therefore the expectations of some of these members of Congress differed from part of the economic consensus at the time.

inferred from stock price responses around the China permanent NTR vote (e.g., Greenland et al. 2020).

Finally, we employ the estimated model to perform counterfactual exercises in which we give politicians perfect information about the upcoming shocks and calculate the change in voting behavior that would have resulted from the additional information. We find that overall support for China's NTR status would not have changed substantially in the presence of perfect information.

In essence, to the question in the title, "Did US Politicians Expect the China Shock?" our answer is "yes, but they did not give it substantial weight." Counterfactual simulations in Section VB, where such weight is increased for all lawmakers in our sample for given baseline information, show that pro-China legislation would have been overturned.

One important premise to the question we are posing is the assumption that the electorate is generally attuned to trade policy positions of their representatives. It would otherwise be unclear why politicians would care about the reaction of voters. While it is implausible to assume that voters have a complete command of specific trade policy measures, a number of papers have documented the impact of the China shock and the recent trade war on electoral outcomes. Autor et al. (2020) find that districts more affected by the China shock saw an increase in Fox News viewership and elected more conservative Republicans and, to a lesser extent, more liberal Democrats, thus inducing more polarization.[5] Another recent contribution by Che et al. (2020) finds that the reaction to the China shock during the 2010s was due to the antitrade turn taken by the Republican party after the appearance of the Tea Party. Blanchard, Bown, and Chor (2019) also document a significant electoral impact of the trade war on the vote share of Republicans in 2018. Interestingly, the negative effect on Republican vote share coming from retaliatory tariffs imposed by US trading partners is not mirrored by a positive effect due to the protection offered by import tariffs. In sum, these recent papers offer a clear justification for making constituent interests a major component of the decision to vote on an important trade policy measure, like maintaining and expanding China's NTR status.[6]

This paper contributes to the literature on Congressional voting, in particular on trade policy. An example is Baldwin and Magee (2000), which estimates the importance of constituent interests and campaign contributions from business and labor groups in three trade bill votes in the 1990s. Relative to Baldwin and Magee (2000), we sharpen the estimation of the role of constituent interests by employing a more precise measure of how constituents were impacted by the policy, but mostly by applying a new econometric methodology to the expectations of politicians. A more recent paper by Feigenbaum and Hall (2015) studies the impact of the China shock on congressional voting on trade bills in general and finds that congressmen in districts more negatively affected are less likely to vote in favor of trade promoting bills, as classified by the Cato Institute, a think tank in Washington, DC. Differently from Feigenbaum and Hall's (2015) retrospective view, we take a prospective angle in

[5] Colantone and Stanig (2018) document a similar result for western European countries.

[6] A recent paper by Fajgelbaum et al. (2020) points to the importance of constituent interests in the structure of tariffs in the trade war started by the United States in 2018. United States import tariffs are such that marginal counties (those with a Republican vote share of around 50 percent) receive the highest level of protection.

modeling voting behavior, where law makers are deciding to vote based on the future electoral consequences of their decision. In the broader literature on the political economy of trade policy, Rodrik (1995) offers a more conceptual framework and depicts trade policy as emerging from demand (interest groups, grassroots, etc) and supply (government) factors. This paper's contribution sheds light on the individual behavior of legislators that constitutes a crucial element of the policy supply side.

Beyond the trade policy literature, this paper speaks to the established empirical literature focused on modeling voting in legislatures. The elements of this vast scholarship that are closer to our paper span political economy and political science (Poole and Rosenthal 1984; Levitt 1996; Heckman and Snyder 1997; Poole and Rosenthal 1997; Jenkins 2000; Clinton, Jackman, and Rivers 2004; McCarty, Poole, and Rosenthal 2006; Canen, Kendall, and Trebbi 2020). Modeling prospective behavior of legislators does figure in this strand of research, as it is often postulated that a representative politician acts by "determining her roll-call vote choice based on which legislative options will maximize her future utility" (Ladewig 2010, p. 501). Stimson, MacKuen, and Erikson (1995, p. 545) argue that in congressional voting "elected politicians … sense the mood of the moment, assess its trends, and anticipate its consequences for future elections." However, expectations and information sets of lawmakers are rarely explicitly modeled as part of standard empirical approaches. The more complex exercise of assessing whether a politician may not be responding to prospective constituent conditions in their vote because of limited information or because of policy preferences appears unexplored. We contribute to this by showing how this moment inequality approach provides a useful stepping-stone in estimating prospective behavior of lawmakers.

Less directly, our application offers a complementary view to the empirical literature focused on modeling the expectations of policy makers. This analysis has traditionally found important applications in Macroeconomics (Primiceri 2006; Sargent, Williams, and Zha 2006), and in this sense, the paper connects to a broader set of questions than congressional voting alone. Our estimates also speak to the modeling of government preferences, a key area of political economy.[7]

I. Empirical Model

This section presents a simple model of probabilistic voting for members of Congress. Indicate a congressional cycle with $t = 1, 2, \ldots, T$. Each period a single bill focused on a main policy issue is introduced—in this application, maintaining normal trade relations with China. As previously discussed, such bills were typically presented to the US legislative branch and voted upon once per congressional cycle, so that t may equivalently indicate time and bill number.

Let us indicate with $x_t \in \mathbb{R}$ a policy position favorable to trade normalization, so that a "yes" vote will indicate a vote for x_t. Consequently, one interprets a "no" vote as a vote against normal trade relations, $q_t \in \mathbb{R}$. This notation allows for the two positions to be affected by some nuance over time and neither position is assumed to

[7] For early applications, see Alesina (1988); Alesina and Tabellini (1990); Drazen and Masson (1994).

be exactly constant in time. In the empirical analysis, we will simply make sure that a "yes" vote will be consistently labeled to the support for the alternative x_t.

Indicate by $i = 1, \ldots, N$ individual legislators, where N is large.[8] We will assume that individual i's preferences are described by a random utility framework. We also posit a spatial voting environment for the members of Congress.[9]

For simplicity of exposition (relaxed in the empirical application later), the deterministic component of the politician's utility is assumed to depend on (i) the distance of the bill from his/her ideological position θ_i; (ii) an electoral motive, summarized by his/her expected future electoral support $V_{i,t+1}$ (for example, due to that expressive voters who are adversely impacted by the China shock may reward or punish i based on his/her voting records.)

Concerning (i), political ideology $\theta_i \in \mathbb{R}$ is a unidimensional and fixed characteristic of i. The assumption of unidimensionality is appropriate in the time period under analysis (see McCarty, Poole, and Rosenthal 2006). The assumption of constant policy preferences has been validated in the literature on congressional voting (Poole 2007 for a discussion) and our results do not appear sensitive to replacing constant ideal points with time-varying ideal points using the estimates from Nokken and Poole (2004), which are available from the authors. We use the ideological positions θ_i from DW-Nominate first dimension scores (Poole and Rosenthal 1997).[10] This follows a common approach in modeling congressional voting when the explicit estimation of such preference parameters is peripheral to the main empirical analysis like in this case (e.g., Mian, Sufi, and Trebbi 2010, 2014).[11]

Concerning (ii), let us indicate by $S_{i,t}$ a proxy for the degree of exposure of the local labor market in the district represented by i at time t to increasing imports from China (the China shock as presented in Autor, Dorn, and Hanson 2016). Assume the potential electoral impact of the China shock in the district represented by politician i at time $t+1$ is defined by

$$V_{i,t+1} = h_t(d_{i,t}, S_{i,t+1}) + e_{i,t+1},$$

where $d_{i,t}$ is the voting decision made by the politician, and

$$
\begin{aligned}
(1) \qquad E\left[e_{i,t+1} \mid d_{i,t}, S_{i,t+1}, \mathcal{I}_{i,t}\right] &= 0, \\
h_t(d_{i,t}, S_{i,t+1}) &= \gamma_t^0 + \gamma_t^2 S_{i,t+1} + \left(\gamma_t^1 - \gamma_t^2\right) S_{i,t+1} \\
&\quad \times \mathbf{1}\left\{d_{i,t} = \text{vote for } x_t\right\},
\end{aligned}
$$

[8] What follows can be applied to each chamber independently at the cost of omitting interactions between the two chambers, such as resolutions and conferences. $N = 435$ for the House and $N = 100$ for the Senate.

[9] Spatial voting is a successful and informative modeling approach to the description of congressional behavior and it has found substantial support in the literature (Poole and Rosenthal 1997; Heckman and Snyder 1997; Clinton, Jackman, and Rivers 2004; McCarty, Poole, and Rosenthal 2006; Bateman, Clinton, and Lapinski 2017).

[10] For further reference, see www.voteview.com and Carroll et al. (2015).

[11] The reader interested in the estimation of θ_i can find a detailed analysis in Canen, Kendall, and Trebbi (2020) and references therein. Due to lack of sample overlap, the Canen, Kendall, and Trebbi (2020) estimates cannot be used in our application.

where $\mathbf{1}\{\cdot\}$ is an indicator function and $\mathcal{I}_{i,t}$ is the information set of politician i at t. Note that the function $h_t(\cdot)$ introduces both a direct effect of the China shock on electoral support independently of i's vote and a component that depends on the interpretation by the voters of their representative i's decision. Online Appendix A.A1 presents a full microfoundation of equation (1).[12]

The expected utility for a politician i of taking decision d_t, given information set $\mathcal{I}_{i,t}$ is

$$U(\xi_{i,t}, d_{i,t}; \theta_i, \mathcal{I}_{i,t}) = u(\|d_{i,t} - \theta_i\|) + \tilde{\delta} E[V_{i,t+1} \mid d_{i,t}, \mathcal{I}_{i,t}] + \begin{cases} \xi_{i,t,x}, & \text{if } d_{i,t} = \text{vote for } x_t; \\ \xi_{i,t,q}, & \text{if } d_{i,t} = \text{vote for } q_t; \end{cases} \tag{2}$$

where $u(\|\cdot\|)$ indicates an ideological loss that is function of the distance of the policy from the ideal point of i. The term $V_{i,t+1}$ indicates the future electoral outcome for the district represented by i. We assume a quadratic loss function $u(\cdot)$ and i.i.d. Gaussian term $\xi_{i,t,d} \sim N(0, \sigma_\xi^2)$. This implies the useful convolution $\xi_{i,t} = \xi_{i,t,q} - \xi_{i,t,x} \sim N(0, 2\sigma_\xi^2)$. A standard identification requirement in discrete choice problems with Gaussian shocks (e.g., in probit) further requires the normalization $2\sigma_\xi^2 = 1$, which we impose.

We define the variable $Y_{i,t}$ as an indicator function that is equal to 1 when legislator i decides to vote "yes" on x_t and 0 when the legislator votes in favor of q_t:

$$\begin{aligned} Y_{i,t} &= \mathbf{1}\{U(\xi_{i,t}, x_t; \theta_i, \mathcal{I}_{i,t}) > U(\xi_{i,t}, q_t; \theta_i, \mathcal{I}_{i,t})\} \\ &= \mathbf{1}\Big\{-\frac{1}{2}[(x_t - \theta_i)^2 - (q_t - \theta_i)^2] \\ &\qquad + \tilde{\delta}(E[V_{i,t+1} \mid x_t, \mathcal{I}_{i,t}] - E[V_{i,t+1} \mid q_t, \mathcal{I}_{i,t}]) \geq \xi_{it}\Big\}. \end{aligned} \tag{3}$$

We can write the probability of $Y_{i,t} = 1$ as

$$\begin{aligned} \Pr(Y_{i,t} = 1 \mid \mathcal{I}_{i,t}) = \Phi\Big(&-\frac{1}{2}((x_t - \theta_i)^2 - (q_t - \theta_i)^2) \\ &+ \tilde{\delta}(E[V_{i,t+1} \mid x_t, \mathcal{I}_{i,t}] - E[V_{i,t+1} \mid q_t, \mathcal{I}_{i,t}])\Big), \end{aligned} \tag{4}$$

where Φ is the standard normal cumulative density function and $E[V_{i,t+1} \mid x_t, \mathcal{I}_{i,t}] - E[V_{i,t+1} \mid q_t, \mathcal{I}_{i,t}]$ is the expected net loss (or net gain) of electoral support due to the China shock in the constituency represented by politician i in the future electoral cycle, given the information available to i at t. This implies that the probability

[12] When considering the economic effects of trade with China, it may be natural to consider the role of exports, and not only imports. Recent work by Feenstra, Ma, and Xu (2019) shows that the exports-led increase in the demand for labor has almost matched the negative employment effects of the China shock. An important observation in this regard is that Feenstra, Ma, and Xu (2019) consider US exports not only to China, but to all its trading partners. When considering only China as a destination market, exports and export growth were markedly smaller and did not have as large a positive effect on employment, as initially shown in Autor, Dorn, and Hanson (2013). Since NTR votes did not have obvious implications for the United States' worldwide export prospects, we exclude exports from our main analysis of voting decisions, but include them in a robustness section in online Appendix E.E3.

of voting "yes" depends on the expectations of the electoral consequences of voting "yes" relative to voting "no", which are unobserved by the econometrician. Intuitively, the higher is the relative expected electoral gain of voting "yes", the higher the likelihood of voting "yes". It follows from (1) that

$$E\left[V_{i,t+1} \mid x_t, \mathcal{I}_{i,t}\right] - E\left[V_{i,t+1} \mid q_t, \mathcal{I}_{i,t}\right] = \left(\gamma_t^1 - \gamma_t^2\right) E\left[S_{i,t+1} \mid \mathcal{I}_{i,t}\right].$$

Setting $\delta_t = \tilde{\delta}\left(\gamma_t^1 - \gamma_t^2\right)$ and simplifying the relative loss function $-\frac{1}{2}\left[\left(x_t - \theta_i\right)^2 - \left(q_t - \theta_i\right)^2\right]$ as $a_t \theta_i + b_t$, where $a_t = x_t - q_t$ and $b_t = \frac{1}{2}\left(q_t^2 - x_t^2\right)$, we rewrite (3) and (4), respectively, as

$$Y_{i,t} = \mathbf{1}\left\{a_t \theta_i + b_t + \delta_t E\left[S_{i,t+1} \mid \mathcal{I}_{i,t}\right] \geq \xi_{i,t}\right\}, \tag{5}$$

$$\Pr\left(Y_{i,t} = 1 \mid \mathcal{I}_{i,t}\right) = \Phi\left(a_t \theta_i + b_t + \delta_t E\left[S_{i,t+1} \mid \mathcal{I}_{i,t}\right]\right). \tag{6}$$

This structure of preferences separates (in a somewhat restrictive way) the role of a_t, interpretable as the politicians' weight on ideology, defined as an aggregate across all primary issues (e.g., taxes, gun control, government deficit, etc.) of concern to them and/or to local voters, and the role of δ_t, specifically indicating the differential alignment with constituents along the future China Shock dimension once the role of average ideology is accounted for. We discuss two distinct approaches to the estimation of these and the other parameters in the next section.

II. Expectations and Information Set of Politicians: Estimation

A key contribution of this paper is the analysis of the information set available to politicians to forecast the labor market effects of the China shock at the time of a roll call vote. There are two fundamentally different approaches, which in turn hinge on the answer to the following question: Is the politician's information set $\mathcal{I}_{i,t}$ known to the econometrician? When the answer is in the affirmative, estimation can be performed by maximum likelihood or method of moments. When the econometrician knows only a subset of the information available to politicians, then one can adopt a moment inequality estimator. We discuss these two approaches in turn, but we first start with describing the three benchmark information sets that we will consider throughout the paper.

(i) Minimal Information: The politician knows his own ideological position θ_i, but the only information a politician has about the economic impact of the China shock is the current share of population employed in manufacturing in district i, $ShareMfg_{i,t}$.

(ii) Baseline Information: The politician has access to the minimal information set, plus the current period China shock $S_{i,t}$.

(iii) Perfect Foresight: The politician has perfect foresight of the labor market consequences of the China shock, so that $E\left[S_{i,t+1} \mid \mathcal{I}_{i,t}\right] = S_{i,t+1}$.

A. *Politician's Information Set Fully Known to the Econometrician: MLE*

When the econometrician knows the content of the politician's information set, then the parameter vector $\omega_t = \{a_t, b_t, \delta_t\}$ can be estimated by maximum likelihood for each cycle. Based on expression (6), the log-likelihood function takes the form:

$$\ln\mathcal{L}\Big(\omega_t \big| \{Y_{i,t}, \theta_i, \mathcal{I}_{i,t}\}_{i=1}^N\Big) = \sum_{i=1}^N Y_{i,t}\ln\Big[\Phi\Big(a_t\theta_i + b_t + \delta_t E\big[S_{i,t+1}\big|\mathcal{I}_{i,t}\big]\Big)\Big] + (1 - Y_{i,t})\ln\Big[1 - \Phi\Big(a_t\theta_i + b_t + \delta_t E\big[S_{i,t+1}\big|\mathcal{I}_{i,t}\big]\Big)\Big]. \tag{7}$$

Maximizing (7) requires specifying the information set $\mathcal{I}_{i,t}$. In the case of perfect foresight, (7) is maximized after replacing $E\big[S_{i,t+1}\big|\mathcal{I}_{i,t}\big]$ with $S_{i,t+1}$. In the case of minimal information set, the expectation of the China shock is derived as the predicted value of the following OLS regression: $S_{i,t+1} = \beta_0 + \beta_1\theta_i + \beta_2 ShareMfg_{i,t} + \epsilon_{i,t+1}$.[13] For the baseline information set, we can perform a similar two-step procedure, albeit with an OLS regression that contains a larger set of regressors, reflecting a richer knowledge by the politician. This methodology also imposes that politicians have rational expectations, i.e., a mean zero expectation error that is uncorrelated with the expectation.

The key assumption of the maximum likelihood approach is that we, as econometricians, are confident about what enters the politician's information set. When one misspecifies the politician's information set, the parameter estimates ω_t will be biased. The direction of the bias cannot be characterized in general, so for our case we resort to Monte Carlo simulations to illustrate the problem in online Appendix B.

One specific instance lends itself to an intuitive explanation. When the econometrician incorrectly assumes that the politician has perfect foresight, the bias that arises is similar to the case of error in variables in a linear regression setting. The intuition is that $E\big[S_{i,t+1}\big|\mathcal{I}_{i,t}\big]$ is measured with error when we replace it with $S_{i,t+1}$ and that error is, by assumption of rational expectations, uncorrelated with $E\big[S_{i,t+1}\big|\mathcal{I}_{i,t}\big]$.Similarly to a linear regression setting, this will lead to an attenuation bias in the estimated coefficient δ_t. Assume that the true δ_t is negative and consider two representatives in districts A and B, who form their expectations based only on a minimal information set, which includes the manufacturing share in the region. Assume that district A and B have similar manufacturing shares, but different industrial composition. Hence, the two representatives predict a similar import shock, but in reality, district A is much more severely affected than district B. We, the econometricians, assume that these representatives are instead very well informed about the imminent import increases. Because A and B expect a similar impact, they vote similarly on the bill. The econometrician, however, observing a similar voting behavior between politicians A and B, concludes that δ is smaller (in absolute value) and that the politicians place little weight on the import shock. The bias can

[13] See Manski (1991) and Ahn and Manski (1993).

be large (around 40 percent) under realistic data configurations, as shown in online Appendix B.

B. *Politician's Information Set Partially Known to the Econometrician: Moment Inequality Approach*

In the previous subsection, we have shown that the maximum likelihood approach relies on an accurate knowledge by the econometrician of the information set possessed by the politician. The alternative estimation method proposed by Dickstein and Morales (2018), based on moment inequalities, does not require full knowledge of $\mathcal{I}_{i,t}$, but rather of a subset of variables $Z_{i,t} \subseteq \mathcal{I}_{i,t}$. That is, the politician may know more than $Z_{i,t}$ in forming his/her forecast, but she knows at least the covariates in $Z_{i,t}$. Assuming only partial knowledge of $\mathcal{I}_{i,t}$ comes at the cost of less precise identification. We will not be able to point identify the elements of the parameter vector ω_t, but only to set identify them. Whether these sets are sufficiently tight to be informative will be carefully discussed in the results section.

The second important goal of our analysis is to ascertain the extent of the information set of legislators. This, in turn, involves a formal analysis of which subset of variables a politician considers at the time of his/her vote through an application of specification selection tests proposed by Bugni, Canay, and Shi (2015). Being able to reject that certain variables are used in the politician's forecast allows us to learn about the process of decision making of legislators: what they knew and considered relevant at the time of their vote. In this exercise, the voting model and data are kept constant, but the subset of variables assumed part of the information set is varied.

In what follows, we allow politicians to have time varying information sets and we formally test whether certain groups of legislators have identical information sets or not. For instance, we assess whether members of higher levels of chamber seniority have broader information sets than lower seniority members, or whether members of opposing parties share the same information set. Questions of asymmetry of information sets across party lines are increasingly common in the political economy literature focused on polarization[14] and our application offers a formal approach to this problem for members of Congress.

A final question that the approach allows us to answer is whether, had politicians had a more complete information set, their votes for trade normalization with China would have been different. These counterfactuals are simulated within the same structure of expectations and information we just described.

Throughout, we maintain the assumption of rational expectations on the part of politicians, that is the expectational error $\epsilon_{i,t+1} = S_{i,t+1} - E[S_{i,t+1} | \mathcal{I}_{i,t}]$ has mean zero, $E[\epsilon_{i,t+1} | \mathcal{I}_{i,t}] = 0$ and is uncorrelated with $E[S_{i,t+1} | \mathcal{I}_{i,t}]$. This means that politicians do not systematically skew their prediction or ignore elements of their information set which would systematically help in forecasting $S_{i,t+1}$. In online Appendix A.A2, we discuss different plausible data generating processes that support the rational expectation assumptions, given the formulation of the baseline China shock

[14] See Alesina, Miano, and Stantcheva (2020).

measure.[15] We follow Dickstein and Morales (2018) in generating two sets of moment inequalities that identify the possible values that the parameters of interest can take (i) odd-based moment inequalities and (ii) revealed preference moment inequalities.[16] In the following subsection, we go through the main steps of the derivation of the inequalities to illustrate the basic intuition.

Odds-Based Moment Inequalities.—We use the definition in (5) to obtain

$$\mathbf{1}\{a_t\theta_i + b_t + \delta_t E[S_{i,t+1} | \mathcal{I}_{i,t}] - \xi_{i,t} \geq 0\} - Y_{i,t} = 0. \tag{8}$$

This expression depends on the unobserved shock realization $\xi_{i,t}$ and $\mathcal{I}_{i,t}$. Therefore, we take the expectation of (8) conditional on $\mathcal{I}_{i,t}$ and manipulate the expression to obtain the following equality:

$$E\left[(1 - Y_{i,t})\frac{\Phi(a_t\theta_i + b_t + \delta_t E[S_{i,t+1} | \mathcal{I}_{i,t}])}{1 - \Phi(a_t\theta_i + b_t + \delta_t E[S_{i,t+1} | \mathcal{I}_{i,t}])} - Y_{i,t} | \mathcal{I}_{i,t}\right] = 0.$$

This equality still depends on the expectation $E[S_{i,t+1} | \mathcal{I}_{i,t}]$, which in turn depends on the true information set $\mathcal{I}_{i,t}$, an object that we do not observe. However, under the assumption that the expectational error $S_{i,t+1} - E[S_{i,t+1} | \mathcal{I}_{i,t}]$ has mean zero and from the property that $\Phi/(1 - \Phi)$ is convex, one can replace $E[S_{i,t+1} | \mathcal{I}_{i,t}]$ with $S_{i,t+1} - \epsilon_{i,t+1}$ and apply Jensen's inequality to derive the following inequality:

$$E\left[(1 - Y_{i,t})\frac{\Phi(a_t\theta_i + b_t + \delta_t S_{i,t+1})}{1 - \Phi(a_t\theta_i + b_t + \delta_t S_{i,t+1})} - Y_{i,t} | \mathcal{I}_{i,t}\right] \geq 0. \tag{9}$$

Consider now a subset of the information set $Z_{i,t} \subseteq \mathcal{I}_{i,t}$. By invoking the Law of Iterated Expectations, we may replace the unobserved information set $\mathcal{I}_{i,t}$ by $Z_{i,t}$, and obtain the following inequality from (9):

$$\begin{aligned} &E\left[m_l^{ob} \middle| Z_{i,t}\right] \geq 0, \\ &m_l^{ob} = (1 - Y_{i,t})\frac{\Phi(a_t\theta_i + b_t + \delta_t S_{i,t+1})}{1 - \Phi(a_t\theta_i + b_t + \delta_t S_{i,t+1})} - Y_{i,t}. \end{aligned} \tag{10}$$

Notice that (10) is increasing in δ_t, so this condition identifies a lower bound for this parameter. We use the subscripts l, u to indicate the lower bound and upper bound inequalities.

Following a similar logic, one can derive a moment condition that further bounds the parameters of interest ω_t:

$$\begin{aligned} &E\left[m_u^{ob} \middle| Z_{i,t}\right] \geq 0, \\ &m_u^{ob} = Y_{i,t}\frac{1 - \Phi(a_t\theta_i + b_t + \delta_t S_{i,t+1})}{\Phi(a_t\theta_i + b_t + \delta_t S_{i,t+1})} - (1 - Y_{i,t}). \end{aligned} \tag{11}$$

[15] We also assess the implications of different types of violations of the rational expectations assumption in online Appendix F.

[16] Specifically, see their online Appendix C for additional details.

Notice that m_u^{ob} is decreasing in δ_t and therefore moment inequality (11) identifies an upper bound for this parameter.

Revealed Preference Moment Inequalities.—The second set of moment inequalities derives from the revealed preference argument that a politician will vote "yes" if and only if the benefit from doing so is positive, hence

$$Y_{i,t}\big(a_t\theta_i + b_t + \delta_t E\big[S_{i,t+1} \,|\, \mathcal{I}_{i,t}\big] - \xi_{i,t}\big) \geq 0. \tag{12}$$

Because, again, $\xi_{i,t}$ is unobserved, we take the expectation of (12), conditional on $\mathcal{I}_{i,t}$ and obtain the following inequality:

$$E\Big[Y_{i,t}\big(a_t\theta_i + b_t + \delta_t E\big[S_{i,t+1} \,|\, \mathcal{I}_{i,t}\big]\big) + \Gamma_{i,t} \,|\, \mathcal{I}_{i,t}\Big] \geq 0, \tag{13}$$

where $\Gamma_{i,t} = -E\big[Y_{i,t}\xi_{i,t} \,|\, \mathcal{I}_{i,t}\big] = (1 - Y_{i,t})\dfrac{\phi\big(a_t\theta_i + b_t + \delta_t E\big[S_{i,t+1} \,|\, \mathcal{I}_{i,t}\big]\big)}{1 - \Phi\big(a_t\theta_i + b_t + \delta_t E\big[S_{i,t+1} \,|\, \mathcal{I}_{i,t}\big]\big)}$, and ϕ is the standard normal probability density function. Once again the expression in inequality (13) contains the unobserved expectation $E\big[S_{i,t+1} \,|\, \mathcal{I}_{i,t}\big]$ and information set $\mathcal{I}_{i,t}$. Because $\phi/(1 - \Phi)$ is convex, with the rational expectation assumption and $Z_{i,t} \subseteq \mathcal{I}_{i,t}$, we can apply the same logic as for inequality (10). The resulting inequality will be weaker than (13) and is given by

$$E\big[m_l^{rp} \,|\, Z_{i,t}\big] \geq 0, \tag{14}$$

$$m_l^{rp} = Y_{i,t}\big(a_t\theta_i + b_t + \delta_t S_{i,t+1}\big) + (1 - Y_{i,t})\frac{\phi\big(a_t\theta_i + b_t + \delta_t S_{i,t+1}\big)}{1 - \Phi\big(a_t\theta_i + b_t + \delta_t S_{i,t+1}\big)}. \tag{15}$$

Starting from another revealed preference inequality,

$$(1 - Y_{i,t})\big(\xi_{i,t} - a_t\theta_i - b_t - \delta_t E\big[S_{i,t+1} \,|\, \mathcal{I}_{i,t}\big]\big) \geq 0,$$

we can obtain a second revealed-preference moment inequality in a similar manner:

$$E\big[m_u^{rp} \,|\, Z_{i,t}\big] \geq 0, \tag{16}$$

$$m_u^{rp} = -(1 - Y_{i,t})\big(a_t\theta_i + b_t + \delta_t S_{i,t+1}\big) + Y_{i,t}\frac{\phi\big(a_t\theta_i + b_t + \delta_t S_{i,t+1}\big)}{\Phi\big(a_t\theta_i + b_t + \delta_t S_{i,t+1}\big)}.$$

Partial Identification.—The moment inequalities defined by (10), (11), (14), and (16) are conditional on values of the vectors $Z_{i,t}$, which we allow to contain different variables, characterizing different possible information sets possessed by politicians. The following theorem indicates that the true parameter vector $\omega_t = \{a_t, b_t, \delta_t\}$ is contained in the set of parameters that are in compliance with the odds-based and revealed preference moment inequalities. Hence, the parameters of interest are partially identified.[17]

THEOREM 1 (Dickstein and Morales 2018): *At the true value of the parameter vector* $\omega_t = \{a_t, b_t, \delta_t\}$ *the following four moment inequalities are satisfied:*

$$\begin{cases} E\left[m_u^{ob} \middle| Z_{i,t}\right] \geq 0; \\ E\left[m_l^{ob} \middle| Z_{i,t}\right] \geq 0; \\ E\left[m_u^{rp} \middle| Z_{i,t}\right] \geq 0; \\ E\left[m_l^{rp} \middle| Z_{i,t}\right] \geq 0; \end{cases}$$

where $Z_{i,t} \subseteq \mathcal{I}_{i,t}$ *is the set of variables known by politician i at time t.*

Conditional moments (10), (11), (14), and (16) cannot be directly employed for empirical applications because conditioning on each possible value of $Z_{i,t}$ is computationally unfeasible. The standard solution in the moment inequality literature, which we adopt, is to transform conditional moment inequalities into unconditional moment inequalities, which can be directly employed in estimation. This is not innocuous in that information is lost in transitioning from conditional inequalities to a relatively smaller set of unconditional inequalities. As a result, the parameters that satisfy conditional moment inequalities may be a small subset of those that satisfy the unconditional moments. Whether these larger confidence sets remain sufficiently informative is again an issue to be reckoned with once we discuss our results.[18] The estimation implementation is detailed in online Appendix C.

C. *Further Robustness of the Methodology*

A relevant issue pertinent to our application is whether politicians are uncertain about the impact of future import shock on future electoral support, and need to form expectations about it, as well as the China shock. This is an issue related to the uncertainty specific to the component $\left(\gamma_t^1 - \gamma_t^2\right)$ in the model, to which we need to explore sensitivity in the construction of our estimator. In online Appendix A.A3, we clarify under which conditions we can allow this uncertainty and we reinterpret the coefficients estimates in light of this modification.

III. Institutional Background and Data

The background for the series of roll call votes that we employ in this paper is the extension of normal trade relations to the People's Republic of China, after their suspension in 1951. NTR status was restored in 1980 under Title IV of the Trade Act of 1974 and was dependent on the presence of a bilateral trade agreement to be renewed every three years and on compliance with the Jackson-Vanik amendment

[17] See online Appendix C in Dickstein and Morales (2018) for the proof of this result.
[18] For a complete discussion see Andrews and Shi (2013).

on freedom of emigration, required for nonmarket economies.[19] China's NTR status would be renewed automatically every year upon the President recommendation unless Congress disapproved it by enacting a joint resolution. It is widely recognized that these resolutions were spurred by humanitarian and foreign policy considerations following the Tiananmen Square events of 1989 (Pregelj 1998). Congress sought to provide incentives, through withholding of NTR status, to the Chinese government to address issues of human rights. In this effort, it clashed with the executive branch, a fact reflected in the several episodes in which the resolutions to disapprove NTR passed in the House, but died in the Senate or were overturned by a Presidential veto. In light of these considerations, it should be clear that we do not view the threat to local economic interests as the only driver of the legislators' roll call votes, and ideological considerations in the utility function of legislators account for this.

A. *Roll Call Votes*

The sample for our estimation includes individual legislator roll call votes for 12 House joint resolutions that took place every year from 1990 to 2001, as listed in online Appendix Table D.1. Three of these joint resolutions to disapprove NTR extension were passed by the US House of Representatives[20] in the years 1990, 1991 and 1992, but not voted on or struck down in the Senate.[21]

The website voteview.com provides the roll call votes, together with the ICPSR code for each legislator, the congressional district, the Party, and the first two dimensions of the DW-Nominate score, a multidimensional scaling application developed by Poole and Rosenthal (1997) and the first dimension of which is our proxy for θ_i (Lewis et al. 2022).[22] Instead of "yea" and "nay," we indicate all votes as pro and against China NTR for ease of interpretation (a "yea" vote in favor of disapproving China's NTR is a vote against China).

Figure 1 shows that support for NTR is not purely along party lines and changes over time. Democrats are relatively more supportive of NTR in the middle of the sample period, while Republicans become increasingly supportive of NTR over time. There is switching of positions within individual legislators as well. Figures 2 and 3 show, by party, how many legislators switch position or maintain their vote relative to the previous year. On average, every year 15 percent (17 percent) of Republican (Democratic) legislators change their position relative to the previous year. In sum, there is sufficient heterogeneity in positions across parties and within legislators over time to justify that the changes in the electoral effects of the vote could play a role beyond constant preferences and ideology of legislators.

We also include in the data the voting outcomes of the bill HR 4444 in 2000 which would grant China permanent normal trade relations, conditional on China's

[19] See *CRS Report for Congress* (Pregelj 2001) for further details.

[20] In what follows, we exclusively focus on votes in the House of Representatives to exploit the pertinent commuting zone level variation in the local labor market effects of the China shock.

[21] Two House resolutions, HR2212 and HR5318, in the 102nd Congress passed both in the House and the Senate and were vetoed by President George H.W. Bush. There was no action on NTR bills of China in the Senate after 1992.

[22] voteview.com represents one of the most comprehensive and popular sources of measures of ideological positions in US politics and is a standard reference in the political economy literature on Congress. See the Introduction for references.

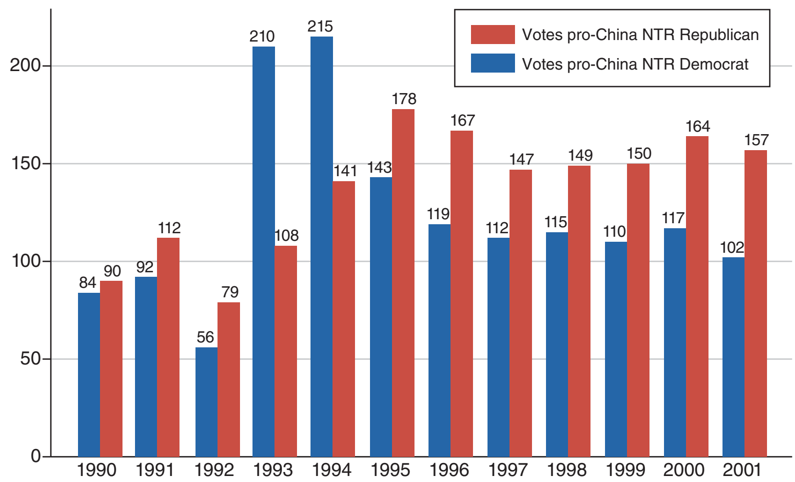

Figure 1. Roll Call Votes pro-China by Party

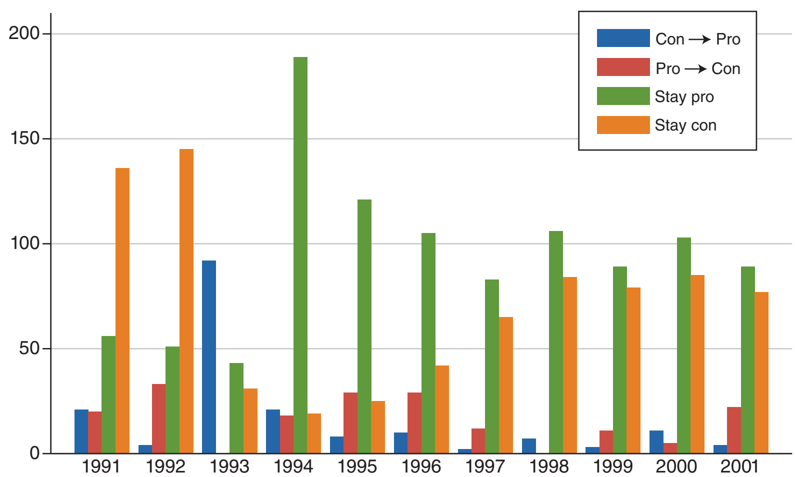

Figure 2. Roll Call Vote Switching: Democrats

accession to the World Trade Organization. The results are stable regardless of the inclusion or exclusion of this bill.[23]

Note that the period 1990–1992 covers the final years of the George H.W. Bush administration.[24] The period 1993–1996 coincides with the first Bill Clinton administration and the period 1997–2001 covers the second Clinton administration and the first year of the George W. Bush presidency. Given the important role played by the executive branch in the legislative evolution of China's NTR status, this subdivision

[23] The shares of votes in favor of China by year is reported in Figure 4.

[24] NTR votes became not perfunctory only after the historical event of Tienanmen Square. From 1979 through 1989, presidents Carter, Reagan, and Bush did not face any congressional opposition to the granting of most-favored-nation privileges to China (Johnson 2006).

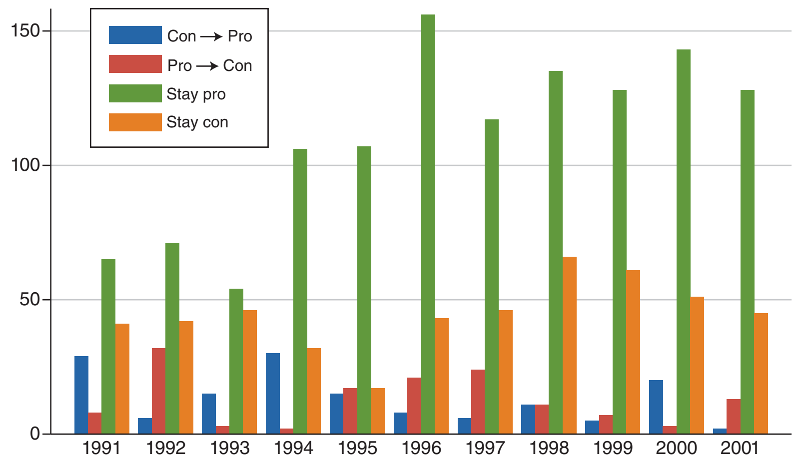

Figure 3. Roll Call Vote Switching: Republicans

of the sample period is explored in our analysis, by allowing for different parameters and expectations during each administration.

B. *The China Shock*

The exposure to the China shock at the district level is generated from the import shocks in different local labor markets nested within the district. We start with constructing the import shocks at the level of commuting zones (CZs), which are clusters of adjoining counties characterized by strong commuting ties and have been conceptualized as local labor markets in the literature (Autor, Dorn, and Hanson 2013; Acemoglu et al. 2016). The shocks from different CZs are then aggregated to congressional districts (CDs) as in Autor et al. (2020).

Exposure to Import Shock at the Commuting Zone Level.—Future supply shocks from China faced by the commuting zone j is constructed according to

$$S_{j,t+1} = \sum_k \frac{L_{j,k,t}}{L_{j,t}} \frac{\Delta M_{k,t+1}^{oth}}{Y_{k,t} + M_{k,t} - X_{k,t}}. \tag{18}$$

In this expression, $\Delta M_{k,t+1}^{oth}$ is the change in import of good k from China by eight other (non-US) high-income countries over five years in the future.[25] It reflects the rising supply capacity of China due to its economic reforms, and is arguably exogenous to the US product-demand shocks from the perspectives of the US local economies (Autor, Dorn, and Hanson 2013; Acemoglu et al. 2016; Autor et al. 2020). The future import growth is then normalized by the contemporaneous absorption (US industry output plus net imports, $Y_{k,t} + M_{k,t} - X_{k,t}$) at the industry level. $L_{j,k,t}/L_{j,t}$ denotes the share of industry k in CZ j's total employment in period t.

[25] The eight other high-income countries are Australia, Denmark, Finland, Germany, Japan, New Zealand, Spain, and Switzerland.

The Bartik-style measure (18) summarizes the exposure of CZ j to China's future supply shocks from the standpoint of t. Having a perfect foresight of (18) not only requires the information on contemporaneous employment composition of the local labor market and domestic absorption of different industries, but also knowledge on supply shocks from China five years in the future. In our analysis, the future shocks correspond to the import supply growth over the period 1990–1995, 1991–1996, …, 2001–2006, which overlaps with China's post-WTO-accession period when the United States' witnessed the most intense increase in import competition from China.[26]

While the politicians may not have full knowledge of future import shocks as in (18), they may use the information of the past shocks to form expectations. For the years 1993–2001, we construct import shock in the past five years analogously as follows:

$$S_{j,t} = \sum_k \frac{L_{i,k,t-5}}{L_{i,t-5}} \frac{\Delta M_{k,t}^{oth}}{Y_{k,t-5} + M_{k,t-5} - X_{k,t-5}}, \tag{19}$$

where $\Delta M_{k,t}^{oth}$ denotes the change in import of good k from China by eight other (non-US) high-income countries over the previous five years.[27]

The baseline measures (18) and (19) follow the specification in Acemoglu et al. (2016) and Autor et al. (2020), and can be derived from workhorse trade models with a gravity structure. However, differently from the literature focusing on the impacts of contemporaneous trade shocks on local economies, our study aims at evaluating the extent to which politicians foresaw the future import shock from China in (18), and whether they acted on the relevant information when setting China-specific trade policies. The trade, employment and output data that are employed to construct the China shock measures are detailed in online Appendix D.1.

Exposure to Import Shocks at the Congressional District Level.—Following Autor et al. (2020), we map economic outcomes in CZs to CDs as follows. We start with the geographic relationships between counties and congressional districts provided by the Missouri Census Data Center (MCDC).[28] Counties are sometimes split across different CDs, and the MCDC concordance provides information

[26] For the baseline analysis, we construct future and past China shocks based on a five-year window with the consideration that it is probably better able to reflect the underlying shift in China's import supply capacity while being consistent with the expected career horizon of the incumbents. Our results remain robust to alternative measures based on a two-year window or a ten-year window (available upon request).

[27] Due to data constraints, for years 1990–1992, we use the two-year-lagged variables to construct the past shocks. To be specific, for $t = 1990, 1991, 1992$,

$$S_{i,t} = \sum_k \frac{L_{i,k,t-2}}{L_{i,t-2}} \frac{\Delta M_{k,t}^{oth}}{Y_{k,t-2} + M_{k,t-2} - X_{k,t-2}},$$

where $\Delta M_{k,t}^{oth}$ denotes the change in import of good k from China by eight other (non-US) high-income countries over the past two years. As is discussed in online Appendix D.D1, we use the data from County Business Patterns (CBP) to construct employment shares. For 1990–1992, calculating $L_{i,k,t-2}/L_{i,t-2}$ requires the 1985–1987 CBP data with industries classified based on 1977 SIC codes. For the purpose of analysis, the data needs to be mapped to the 1987 SIC codes. However, the crosswalk from 1977 SIC to 1987 SIC involves many splits of industries. As a result, the concordance leads to a structural break in the employment measures for some localities over 1987–1988. For the concern of systematic measurement errors, we don't use the CBP data prior to 1988 for the main analysis.

[28] http://mcdc.missouri.edu/applications/geocorr1990.html.

on the distribution of the county population in each CD. We then ascribe to each county-by-congressional district cell the CZ-level import shock that corresponds to the county, and weight each cell by its share of population in the district. Lastly, we aggregate the weighted shocks across cells to the CD level. By construction, if a district spans multiple CZs, its exposure to China's rising import competition is the population-share-weighted average of the import shocks in these CZs. Since congressional districts, by construction, have similar population size, they have roughly the same weight in our analysis. In the empirical analysis, we denote the future and past import shocks at the CD level by $S_{i,t+1}$ and $S_{i,t}$, respectively.

During the sample period 1990–2001, the boundaries of county-by-congressional district cells experienced a major change in 1993, but remained stable afterwards. Therefore, for 1990–1992, we map the CZ-level import shocks to congressional districts as defined for the 102nd Congress. For 1993–2001, the mapping is based on the configuration of the 103rd Congress. This treatment does not affect the consistency of our baseline analysis because, as is discussed below, we conduct the estimation by periods based on presidential administrations, and none of the subsamples spans over 1992–1993.

Figure 4 shows the cross-district averages of past and future import shocks. The import shocks are always positive throughout the sample period, but the future shocks move less in tandem with the past shocks in the later years. For the moment inequality estimation discussed in the following section, we detrend the import shocks. The corresponding distributions reported in online Appendix Table D.2 reveal a substantial heterogeneity in exposures across districts. For the periods 1997–2001, 1993–1996, and 1990–1992, the interquartile ranges of future shock are 0.154, 0.078, and 0.143, respectively.

IV. Results from Congressional Voting on NTR with China

A. *Estimation Results*

This section reports our main results. To benchmark our approach to more standard methods using incorrect proxies for the information set of politicians, we start by reviewing estimates of a voting model using a maximum likelihood approach.

Table 1 shows that results substantially vary depending on the informational assumptions made. An important parameter of interest is the weight placed by politicians on their affected subconstituencies δ_t. A reasonable prior for this parameter would be $\delta_t < 0$, i.e., a negative utility weight placed on electoral groups adversely affected by the China shock in the politician's congressional district (ceteris paribus, the politician wishes to minimize these adverse effects). While the parameter estimates for δ are occasionally negative and sometimes of magnitude similar to those obtained using the moment inequality approach, often the coefficients are economically insignificant and not statistically different from zero. For example, for the period 1997–2001 the estimate for δ is 0.018, which is small in absolute value relative to consistent estimates obtained with the moment inequality method. The risk of attenuation from misspecification of the information set of politicians is therefore evident in the MLE case. To further consolidate the intuition we also offer Monte Carlo evidence of the problem in online Appendix B.

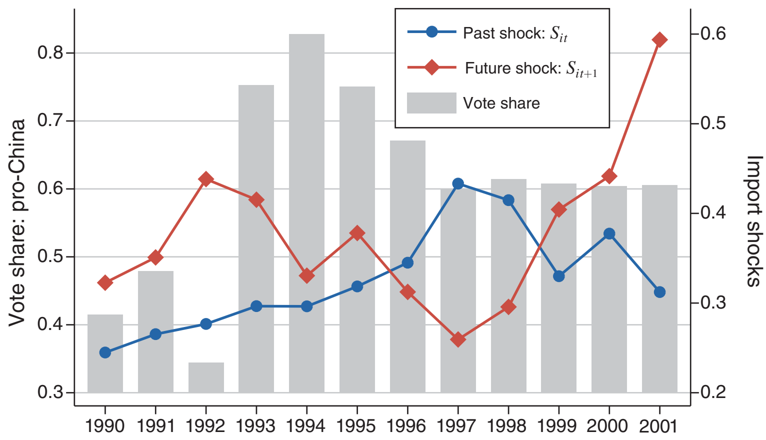

Figure 4. Import Shocks from China and pro-China NTR Vote Share

Notes: The bar chart shows the share of votes in favor of renewing China's NTR status for the bills introduced in the House over 1990 to 2001. The data in 2000 also includes the bill HR 4444 which granted China permanent normal trade relations. The dot blue line represents the average of past import shock across congressional districts, and the diamond red line shows the average of the future import shock. In order to put the data on a comparable five-year scale, past import shocks over 1990–1992 are multiplied with the factor 5/2.

Table 1—MLE Estimates Based on Different Information Sets

	$\hat{a}$ (SE)	$\hat{b}$ (SE)	$\hat{\delta}$ (SE)
Panel A. 1997–2001			
Minimal information $Z_{i,t} = \{ShareMfg_{i,t}, \theta_i\}$	0.650 (0.063)	0.262 (0.025)	−0.810 (0.314)
Baseline information $Z_{i,t} = \{S_{i,t}, ShareMfg_{i,t}, \theta_i\}$	0.643 (0.062)	0.263 (0.025)	−0.786 (0.279)
Perfect foresight	0.619 (0.062)	0.263 (0.025)	−0.018 (0.211)
Panel B. 1993–1996			
Minimal information $Z_{i,t} = \{ShareMfg_{i,t}, \theta_i\}$	−0.046 (0.084)	0.678 (0.033)	−0.440 (0.611)
Baseline information $Z_{i,t} = \{S_{i,t}, ShareMfg_{i,t}, \theta_i\}$	−0.043 (0.084)	0.679 (0.033)	−1.116 (0.549)
Perfect foresight	−0.051 (0.083)	0.678 (0.033)	−0.517 (0.452)
Panel C. 1990–1992			
Minimal information $Z_{i,t} = \{ShareMfg_{i,t}, \theta_i\}$	1.019 (0.104)	−0.179 (0.037)	0.065 (0.351)
Baseline information $Z_{i,t} = \{S_{i,t}, ShareMfg_{i,t}, \theta_i\}$	1.000 (0.105)	−0.181 (0.037)	−0.981 (0.288)
Perfect foresight	1.013 (0.104)	−0.180 (0.037)	−0.669 (0.244)
Panel D. 1990–2001			
Minimal information $Z_{i,t} = \{ShareMfg_{i,t}, \theta_i\}$	0.532 (0.044)	0.280 (0.017)	−0.453 (0.222)
Baseline information $Z_{i,t} = \{S_{i,t}, ShareMfg_{i,t}, \theta_i\}$	0.534 (0.044)	0.280 (0.017)	−0.856 (0.198)
Perfect foresight	0.523 (0.044)	0.280 (0.017)	−0.333 (0.147)

Notes: This table reports the MLE estimates based on equation (7). For the case of minimal information, we replace the term $E[S_{i,t+1} | \mathcal{I}_{i,t}]$ by the predicted value of the OLS regression: $S_{i,t+1} = \beta_0 + \beta_1 \theta_i + \beta_2 ShareMfg_{i,t} + \epsilon_{i,t+1}$. For the case of baseline information, the term is replaced by the predicted value of the OLS regression: $S_{i,t+1} = \beta_0 + \beta_1 \theta_i + \beta_2 ShareMfg_{i,t} + \beta_3 S_{i,t} + \epsilon_{i,t+1}$. For the case of perfect foresight, the term is replaced by $S_{i,t+1}$. Robust standard errors are reported in parentheses.

We now proceed to the results of estimating of equation (6) using the three information sets $Z_{i,t}$: minimal; baseline; and perfect foresight. The 95 percent confidence sets that we report are built through a grid search implementing the generalized moment selection (GMS) method in Andrews and Soares (2010) as detailed in

Table 2—Relevance of Manufacturing Employment Share and Past Import Shock in Explaining Future Import Shocks

Sample period: Dependent variable: $S_{i,t+1}$	1990–2001 (1)	1997–2001 (2)	1993–1996 (3)	1990–1992 (4)
$S_{i,t}$	0.404	0.731	0.214	0.654
	(0.017)	(0.041)	(0.013)	(0.033)
$ShareMfg_{i,t}$	0.705	0.680	0.457	0.700
	(0.021)	(0.036)	(0.020)	(0.045)
Observations	5,494	2,564	1,698	1,232
R^2	0.556	0.584	0.674	0.695

Notes: In order to put the data on a comparable five-year scale, past import shocks over 1990–1992 are multiplied with the factor $5/2$. Robust standard errors in parentheses.

Dickstein and Morales (2018). For each value of ω_t we build a modified method of moments (MMM) statistic, which tends to be large when the moment inequalities are *not* satisfied at that value of the parameters. This is formally tested by constructing the asymptotic distribution of the MMM statistic and rejecting ω_t when the MMM statistic is above the critical value corresponding to the ninetieth percentile of that distribution. Incidentally, empty confidence sets instead have to be interpreted as highly significant rejection (p-value < 0.05) of the corresponding information set. The steps to construct the 95 percent confidence sets are detailed in online Appendix C.C2.

Table 3 reports estimation results splitting the NTR votes by presidential administration and for the full sample 1990–2001. The first period 1990–1992 covers the George H.W. Bush administration, the second period 1993–1996 coincides with the first Clinton administration, and the third period 1997–2001 covers the second Clinton administration.[29] Focusing on subperiods allows for heterogeneity in information and flexibility of the parameters with respect to the behavior/agenda setting of executive branch—a more robust approach that we prefer. We will mostly discuss results for the importance of constituency interests δ, but we will gauge its magnitude in relation to the role of ideology parameter a.

Consider the pooled 1990–2001 results first. Panel A reports estimation results under the assumption that the politicians' information set includes at least the manufacturing share in their district and their own ideology. Under this assumption we report a confidence set $[-1.163, -0.150]$, so that all values in the confidence set are negative, as we would expect if politicians are more likely to vote against China's NTR status when they expect their constituency to be exposed to a larger shock. The confidence set is quite similar under the assumption of baseline information, which includes also the past shock $S_{i,t}$, $[-1.438, -0.500]$.

The magnitude of the parameter δ can be illustrated as follows. Considering two districts whose value of the China shock are at the twenty-fifth and seventy-fifth percentile. The probability of voting in favor of China's NTR status decreases by

[29] As a robustness check, we drop the year 2001, which overlaps with George W. Bush's first year in office during which the final vote took place, but was related to the previous administration's efforts (Pregelj 2001). The results remain similar.

Table 3—Parameter Confidence Sets and Specification Test p-Values

Period	CS of a	CS of b	CS of δ	p-value BP	p-value RC	p-value RS	Num obs.
1997–2001	*Panel A. Minimal information* $Z_{i,t} = \{ShareMfg_{i,t}, \theta_i\}$						
	$[0.450, 0.720]$	$[0.165, 0.260]$	$[-1.975, -0.225]$	0.405	0.405	0.405	2,564
	Panel B. Baseline information $Z_{i,t} = \{S_{i,t}, ShareMfg_{i,t}, \theta_i\}$						
	$[0.435, 0.750]$	$[0.185, 0.290]$	$[-1.560, -0.075]$	0.290	0.275	0.275	2,564
	Panel C. Perfect foresight $Z_{i,t} = \{S_{i,t}, ShareMfg_{i,t}, S_{i,t+1} - E[S_{i,t+1} \mid S_{i,t}, ShareMfg_{i,t}, \theta_i], \theta_i\}$						
	–	–	–	0.010	0.010	0.010	2,564
1993–1996	*Panel A. Minimal information* $Z_{i,t} = \{ShareMfg_{i,t}, \theta_i\}$						
	$[-0.280, 0.100]$	$[0.583, 0.703]$	$[-2.300, 0.900]$	0.330	0.330	0.330	1,698
	Panel B. Baseline information $Z_{i,t} = \{S_{i,t}, ShareMfg_{i,t}, \theta_i\}$						
	$[-0.325, 0.130]$	$[0.598, 0.740]$	$[-3.125, -0.125]$	0.395	0.395	0.395	1,698
	Panel C. Perfect foresight $Z_{i,t} = \{S_{i,t}, ShareMfg_{i,t}, S_{i,t+1} - E[S_{i,t+1} \mid S_{i,t}, ShareMfg_{i,t}, \theta_i], \theta_i\}$						
	–	–	–	0.040	0.035	0.035	1,698
1990–1992	*Panel A. Minimal information* $Z_{i,t} = \{ShareMfg_{i,t}, \theta_i\}$						
	$[0.800, 1.550]$	$[-0.325, -0.125]$	$[-1.125, 2.125]$	0.955	0.955	0.955	1,232
	Panel B. Baseline information $Z_{i,t} = \{S_{i,t}, ShareMfg_{i,t}, \theta_i\}$						
	$[1.025, 1.438]$	$[-0.275, -0.150]$	$[-1.300, 0.000]$	0.165	0.145	0.145	1,232
	Panel C. Perfect foresight $Z_{i,t} = \{S_{i,t}, ShareMfg_{i,t}, S_{i,t+1} - E[S_{i,t+1} \mid S_{i,t}, ShareMfg_{i,t}, \theta_i], \theta_i\}$						
	$[1.000, 1.550]$	$[-0.270, -0.128]$	$[-1.624, 0.096]$	0.255	0.240	0.240	1,232
1990–2001	*Panel A. Minimal information* $Z_{i,t} = \{ShareMfg_{i,t}, \theta_i\}$						
	$[0.463, 0.625]$	$[0.210, 0.270]$	$[-1.163, -0.150]$	0.185	0.185	0.185	5,494
	Panel B. Baseline information $Z_{i,t} = \{S_{i,t}, ShareMfg_{i,t}, \theta_i\}$						
	$[0.463, 0.650]$	$[0.235, 0.287]$	$[-1.438, -0.500]$	0.190	0.190	0.190	5,494
	Panel C. Perfect foresight $Z_{i,t} = \{S_{i,t}, ShareMfg_{i,t}, S_{i,t+1} - E[S_{i,t+1} \mid S_{i,t}, ShareMfg_{i,t}, \theta_i], \theta_i\}$						
	–	–	–	0.010	0.010	0.010	5,494

Note: For the case of perfect foresight, we assume that in addition to $S_{i,t}$, $ShareMfg_{i,t}$ and θ_i, politicians also possess information that is orthogonal to these covariates, i.e., $S_{i,t+1} - E[S_{i,t+1} \mid S_{i,t}, ShareMfg_{i,t}, \theta_i]$.

between 0.022 and 0.059 when the value of the expected China shock goes from its twenty-fifth percentile to its seventy-fifth percentile.[30] This is a moderate effect, especially considering the role of ideology. If we perform the analogous exercise, we find that an interquartile range shift in ideology (from relatively liberal to relatively conservative) calculated at the mean expected China shock value produces an increase in the probability of voting in favor of China of between 0.134 and 0.175 in 1990–2001. In sum, these comparisons allow us to conclude that the effect of ideology is much larger than the effect of constituent interests. These results line up with common findings in the congressional voting literature for most bills (as early as Kalt and Zupan 1984 and for a recent discussion see Poole and Rosenthal 2017).

We will discuss differences in the parameters estimates between subsets of legislators after introducing specification tests that allow us to gauge the information possessed by politicians.

[30] In particular, we calculate these percentage points as the minimum and maximum of $\Phi\left(b + \delta E\left[S_{i,t+1} \mid \mathcal{I}_{i,t}^b\right]^{75\text{th}}\right) - \Phi\left(b + \delta E\left[S_{i,t+1} \mid \mathcal{I}_{i,t}^b\right]^{25\text{th}}\right)$ evaluated at the mean $\theta_i = 0$ and $\omega_t \in \Omega_t^{95\%}$, where $\Omega_t^{95\%}$ is the 95 percent confidence set of the underlying parameters.

B. *Testing for Different Information Sets*

A fundamental first step in the analysis of politicians' decisions is to assess the exact extent of their information sets at the moment of the vote. Intuitively, when $Z_{i,t} \nsubseteq \mathcal{I}_{i,t}$, moment inequalities presented in Section IIB are misspecified, and there could be no values of ω_t that yields the data moments consistent with the misspecified inequality conditions.[31] The estimated confidence set hence could be empty. A formal statistical approach to this form of specification test is presented in Bugni, Canay, and Shi (2015), which allows us to reject the null hypothesis that there exists a value of ω_t within the parameter space that can rationalize the set of odds-based and revealed preference-based moment inequalities.

Specifically, consider two alternative information sets $Z_{i,t}^1$ and $Z_{i,t}^2$. Suppose there is no value of the parameter vector for which the set of moment inequalities hold given $Z_{i,t}^1$, yet there are values of ω_t within the parameter space that can rationalize the set of moment inequalities given $Z_{i,t}^2$. Then, one may infer that $Z_{i,t}^1 \nsubseteq \mathcal{I}_{i,t}$ and cannot reject $Z_{i,t}^2 \subseteq \mathcal{I}_{i,t}$. The rejection of the null in this framework may also indicate misspecification of the original model of decision, so simple rejection of $Z_{i,t}^1$ cannot exclude misspecification per se. What is crucial in this application is that the failure to reject $Z_{i,t}^2$ eliminates this second interpretation of the test. Misspecification of the original model would affect the analysis under both $Z_{i,t}^1$ and $Z_{i,t}^2$, as the decision model is unchanged and only the information set varies across tests, and model misspecification would imply rejection in both instances.

Online Appendix C.C3 reports the full details for the construction of the BP, RC, and RS specification test statistics following Bugni, Canay, and Shi (2015) and the corresponding p-values. Generally speaking, the test BP is less powerful than RC and RS, and rejection of the null hypothesis in any of these tests indicates a rejection of the hypothesis that specific information belongs to the information set of the members of Congress.

We report results for all three tests in Table 3. Columns 4, 5, and 6 report the p-values for BP, RC, and RS respectively. For the full sample and for two of the three presidential administration periods we reject that politicians have perfect foresight, with p-values of 0.01–0.015 depending on the test. For all periods we cannot reject that the politicians had at least the baseline information set. While it is not obvious whether the baseline information set represents the actual amount of knowledge, we believe Table 2 is informative. It shows that the manufacturing employment share and past import shock have significant explanatory power in predicting future shocks. The R^2 is generally around 55 percent and is lower in the later period, suggesting that the variance of the expectational errors have increased over time and close to the end of the sample.

Our results point to information used by politicians worsening over time and their capacity of forecasting the China shock in the following five years deteriorating. Why would legislators be better informed in the earlier part of the period? We hypothesize that this is due to the more predictable nature of the shocks in the early 1990s, when China was specialized in relatively less complex and more labor

[31] To see this, the derivation step that invokes the Law of Iterative Expectation is invalid. Hence, for some values of $Z_{i,t}$, the moment inequalities are unsatisfied for all values of ω_t.

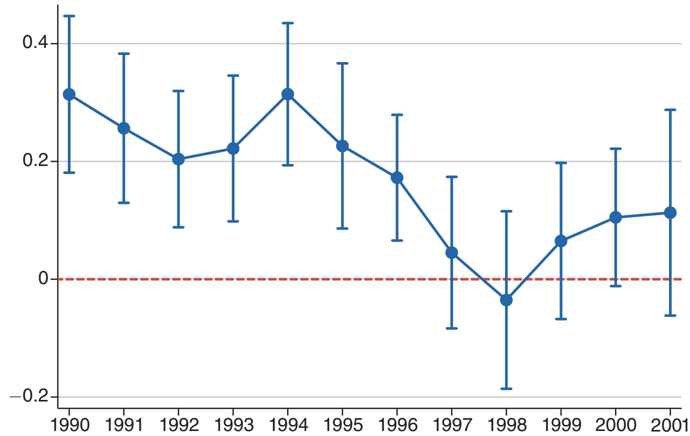

Figure 5. The China Shock over Time

Notes: We explore the autocorrelation of import supply shock from China at the four-digit SIC level by estimating the following equation:

$$\Delta \ln M^{oth}_{k,t,t+5} = \alpha_t \Delta \ln M^{oth}_{k,t-5,t} + D_t + \epsilon_{kt},$$

where $\Delta \ln M^{oth}_{k,t-5,t}$ (respectively, $\Delta \ln M^{oth}_{k,t,t+5}$) measures the change in log import from China by eight developed countries (Australia, Denmark, Finland, Germany, Japan, New Zealand, Spain, and Switzerland) over the period $t-5$ to t (respectively, t to $t+5$). D_t is the year fixed effects, and α_t captures the autocorrelation of import supply shock from China in different periods. Standard errors are clustered at the four-digit SIC level. We use the Davis-Haltiwanger-Schuh approximation of the log growth rate, i.e., $\Delta \ln M^{oth}_{k,t-5,t} \approx 2\left(\frac{M^{oth}_{k,t} - M^{oth}_{k,t-5}}{M^{oth}_{k,t} + M^{oth}_{k,t-5}}\right)$ and $\Delta \ln M^{oth}_{k,t,t+5} \approx 2\left(\frac{M^{oth}_{k,t+5} - M^{oth}_{k,t}}{M^{oth}_{k,t+5} + M^{oth}_{k,t}}\right)$, to avoid dropping observations where imports are zero. The point estimates of α_t and their 95 percent confidence intervals are reported in the figure.

intensive products. In Figure 5, we report the autocorrelation of the China shock at the industry level and it is clear that in the earlier years the shock was more predictable from year to year as the autocorrelation is above 0.3. The decline in the autocorrelation parameter begins its negative adjustment in 1994, which matches (qualitatively) the timing observed for 1993–1996 and 1997–2001 in terms of heightened difficulty of predicting the China shock. This is compatible with the evolution of China's comparative advantage from low value-added to more complex products over time.

Our evidence on the extent of the information set of US politicians and our assessment of their (fairly accurate) predictions on the industry consequences of the China shock is in line with evidence concerning expectations of other types of actors. Greenland et al. (2020) find that equity valuations around the key PNTR vote in Congress correlated with US firms' exposure to trade liberalization with China, supporting the view that stock market participants were systematically pricing firms and sectors exposure to the shock.[32] It is therefore not completely implausible that US legislators had information sets comparable with those of financial agents.

[32] For a discussion on the degree of predictability of trade policy change effects based on stock reactions of both domestic and foreign firms see also Breinlich (2014).

C. *Heterogeneity across Groups of Politicians*

Party.—The estimates and tests presented so far were performed in the universe of the members of Congress during the 1990–2001 period under an assumption of common information sets. It is plausible to hypothesize that politicians from different parties, and with different electoral prospects, might have had varying degrees of knowledge and different expectations.

In this section, we analyze three dimensions of heterogeneity: political party, tenure in office (i.e., experience), and margin of victory in the most recent election. For all tests we report the results for the baseline information set. To anticipate our findings, the picture that will emerge from this heterogeneity analysis is that Democrats at the time of NTR votes were both more informed and more sensitive to constituent interests than Republicans, and that legislators in tighter electoral races placed a heavier weight on the China shock.

In Table 4, we find that Democrats are more informed than Republicans, as we cannot reject that Democrats had at least the baseline information set in all time periods, but we can reject at standard levels of significance that Republicans know the baseline information in all periods, but 1990–1992. We believe part of the rationale for this finding comes from the reliance of Democrats and some northern Republicans on information from sources close to the labor movement.[33] When we obtain nonempty parameter estimates, we find that Democrats display higher (in absolute value) sensitivity to the China shock. Republicans' confidence set for δ straddles zero, while the entire confidence set for the Democratic members of the House is comprised of negative values. In online Appendix Table E.1, we estimate the model separately for Democrats and Republicans based on the specification with minimal information. In this case, we obtain nonempty confidence sets for both parties for each sub-sample period. For Republicans, the 95 percent CS of δ always contains zero. While the estimated confidence sets for Democrats become wider, it appears that the weight on constituent interests is smaller for Republicans compared to Democrats. These findings are in line with Democratic legislators historical alignment with workers' interests and placing more weight on the affected subconstituencies (Poole and Rosenthal 1997).[34]

To further establish the relevance of this dimension of heterogeneity, we also considered behavior of politicians beyond voting, particularly congressional speech (number of speeches related to the "China and trade" issue or the "China and labor" issue delivered). We can see that the nonvoting data aligns with preference and information patterns that we reported.

We only focus on two corroborating findings here and present a full analysis in online Appendix E.E1. First, as the China shock over 2001–2006 is realized,

[33] For example, the Economic Policy Institute predicted in 2000 "the elimination of 872,091 jobs during the next decade," due to the trade deficit with China induced by the PNTR. This is remarkably close to estimates subsequently produced by economists. Acemoglu et al. (2016) find that US trade with China eliminated 985,000 US manufacturing jobs between 1999 and 2011. Autor, Dorn, and Hanson (2013) and Pierce and Schott (2016) estimate that the "China shock" reduced on net US manufacturing employment by 1.5 million jobs between the year 1990 and 2007. Acemoglu et al. (2016) estimate close to a 1.98 million total jobs lost. See footnote 4 for some excerpts from the Congressional Record citing these sources.

[34] In online Appendix F.F2, we further discuss through Monte Carlo simulations the role of misspecifications of information sets with respect to heterogeneity across levels of ideology θ.

Table 4—Parameter Confidence Sets and Specification Test p-Values: Heterogeneity by Party (Baseline Information)

Period	Group	CS of a	CS of b	CS of δ	p-value BP	p-value RC	p-value RS	Num obs.
1997–2001	Democrats	[2.225, 4.625]	[0.850, 1.800]	[−4.425, −0.015]	0.315	0.265	0.265	1,229
	Republicans	–	–	–	0.010	0.010	0.010	1,326
1993–1996	Democrats	[0.800, 3.400]	[1.050, 2.050]	[−6.325, −0.025]	0.380	0.380	0.380	888
	Republicans	–	–	–	0.020	0.015	0.015	806
1990–1992	Democrats	[1.875, 4.375]	[0.050, 0.850]	[−3.700, −0.400]	0.415	0.410	0.410	745
	Republicans	[−1.300, 1.400]	[−0.300, 0.650]	[−1.500, 1.800]	0.315	0.310	0.310	485
1990–2001	Democrats	[1.375, 3.325]	[0.570, 1.290]	[−4.412, −0.417]	0.990	0.990	0.990	2,862
	Republicans	–	–	–	0.010	0.010	0.010	2,s617

Note: The estimation in this table is based on the baseline information $Z_{i,t} = \{S_{i,t}, ShareMfg_{i,t}, \theta_i\}$.

representatives from districts in the top tercile of the exposure to import shock from China raise the related trade and labor issues more often in their speeches, but such response is stronger for Democrats than Republicans. Second, consistently with their higher information and preference weights, Democrats start taking actions earlier. Specifically, for Democrats the congressional speech on China starts surging during the 106th Congress, 1999–2000, while for Republicans, the effect picks up in the 108th, 2001–2002.

Vote Margins and Tenure.—In Table 5, we also explore whether politicians with above-median victory margins in the previous election display differences in the sensitivity to constituent interests. The literature has found higher sensitivity for legislators in tighter races (Mian, Sufi, and Trebbi 2010). We indeed find that, across different periods, legislators in tighter races display confidence sets for δ that are entirely composed of negative values, whereas confidence sets for politicians in safe races often cover both positive and negative values, consistent with lower preference weights. It does not appear to be the case, however, that politicians in tight races were differentially informed relative to politicians elected by larger margins. While it is not an objective of this section to identify whether the heightened sensitivity to the China shock was due to state dependent preferences of politicians, changing with electoral conditions, or due to selection (although this should also reflect in different information sets), the analysis does display a potential for our approach to pick up differential elements of the behavior and knowledge of sub-groups of politicians.

Finally, we explore the role of experience. In Table 6, we divide the sample in two according to whether House members tenure is above or below the median. One may imagine that politicians that are very experienced have better access to various sources of information. We do not find support for this conjecture, as confidence sets for δ of junior legislators appear not systematically different from those of senior ones.

D. *Role of Special Interests*

Special interests' contributions[35] are often listed within the set of potential drivers of congressional voting, but not without substantial uncertainty about the economic

[35] The data on campaign donations employed in the analysis of special interests are obtained from opensecrets.org, based on official Political Action Committee disclosure forms from the US Federal Election Commission

Table 5—Parameter Confidence Sets and Specification Test p-Values: Heterogeneity by Win Margin (Baseline Information)

Period	Group	CS of a	CS of b	CS of δ	p-value BP	p-value RC	p-value RS	Num obs.
1997–2001	Win-margin $>$ median	[0.600, 1.225]	[0.060, 0.270]	[−1.375, 1.500]	0.890	0.890	0.890	1,205
	Win-margin $<$ median	[−0.125, 0.500]	[0.250, 0.438]	[−3.413, −0.525]	0.680	0.675	0.675	1,207
1993–1996	Win-margin $>$ median	[−0.375, 0.350]	[0.475, 0.700]	[−3.250, 2.750]	0.560	0.560	0.560	799
	Win-margin $<$ median	[−1.050, 0.300]	[0.625, 0.975]	[−8.550, −0.225]	0.965	0.965	0.965	800
1990–1992	Win-margin $>$ median	[0.875, 2.125]	[−0.465, −0.105]	[−2.875, 3.312]	0.765	0.755	0.755	568
	Win-margin $<$ median	[0.450, 1.700]	[−0.325, −0.025]	[−4.375, −0.750]	0.510	0.465	0.465	569
1990–2001	Win-margin $>$ median	[0.550, 0.888]	[0.120, 0.232]	[−0.900, 0.800]	0.450	0.450	0.450	2,573
	Win-margin $<$ median	[0.013, 0.475]	[0.282, 0.410]	[−2.565, −0.825]	0.615	0.615	0.615	2,575

Note: The estimation in this table is based on the baseline information $Z_{i,t} = \{S_{i,t}, ShareMfg_{i,t}, \theta_i\}$.

Table 6—Parameter Confidence Sets and Specification Test p-Values: Heterogeneity by Tenure (Baseline Information)

Period	Group	CS of a	CS of b	CS of δ	p-value BP	p-value RC	p-value RS	Num obs.
1997–2001	Tenure $>$ median	[0.750, 1.075]	[0.158, 0.255]	[−0.850, 0.600]	0.120	0.105	0.105	1,372
	Tenure $<$ median	[0.180, 0.700]	[0.238, 0.375]	[−2.750, −0.875]	0.185	0.165	0.165	1,176
1993–1996	Tenure $>$ median	[−0.400, 0.350]	[0.525, 0.700]	[−3.350, 1.363]	0.460	0.460	0.460	863
	Tenure $<$ median	[−0.700, 0.237]	[0.650, 0.925]	[−5.875, 0.312]	0.630	0.615	0.615	821
1990–1992	Tenure $>$ median	[0.825, 2.400]	[−0.175, 0.175]	[−5.513, 0.175]	0.920	0.920	0.920	572
	Tenure $<$ median	[0.938, 1.750]	[−0.495, −0.270]	[−1.600, 0.900]	0.260	0.235	0.235	634
1990–2001	Tenure $>$ median	[0.615, 0.965]	[0.200, 0.287]	[−1.758, −0.270]	0.275	0.275	0.275	2,488
	Tenure $<$ median	[0.238, 0.475]	[0.262, 0.338]	[−1.502, −0.355]	0.145	0.135	0.135	2,950

Note: The estimation in this table is based on the baseline information $Z_{i,t} = \{S_{i,t}, ShareMfg_{i,t}, \theta_i\}$.

magnitude of their effect. There is evidence of a prominent role of special interest giving in certain votes (e.g., the Emergency Economic Stabilization Act of 2008, Mian, Sufi, and Trebbi 2010), but no consensus in the political economy literature on its role for the bulk of all congressional activity (Stratmann 2005). In the case of China's NTR, we find no evidence that special interests, both in terms of campaign contributions from business organizations (corporations and business associations) or from labor unions, influence the estimated effects of ideology and constituent interests on congressional votes in our main specification. We base this assessment on three main sets of empirical evidence which we report below.

First, as an extra dimension of heterogeneity, we separate politicians into two groups, depending on whether the campaign contributions from business interests are above or below the median in the sample. The degree of heterogeneity found in Table 7 is minimal, with marginally more negative estimates of constituent weights for politicians with contributions above the median in later periods. This may suggest

(Center for Responsibe Politics 1990–2016). opensecrets.org is a website run by the Center for Responsive Politics, one of the main nonpartisan organizations in Washington DC, dedicated to electoral transparency and to the collection of information related to campaign spending and lobbying disclosures. (See https://www.opensecrets.org/pacs/.)

Table 7—Parameter Confidence Sets and Specification Test p-Values: Heterogeneity by Campaign Contributions from Business (Baseline Information)

Period	Group	CS of a	CS of b	CS of δ	p-value BP	p-value RC	p-value RS	Num obs.
1997–2001	Money $>$ median	[−0.350, 0.450]	[0.492, 0.720]	[−2.850, −0.600]	0.360	0.360	0.360	1,273
	Money $<$ median	[0.525, 1.078]	[−0.045, 0.135]	[−2.200, 0.425]	0.380	0.345	0.345	1,273
1993–1996	Money $>$ median	[−1.000, 0.250]	[0.725, 1.075]	[−7.000, −0.400]	0.715	0.695	0.695	844
	Money $<$ median	[-0.312, 0.588]	[0.450, 0.700]	[−3.650, 2.125]	0.665	0.665	0.665	844
1990–1992	Money $>$ median	[0.150, 1.100]	[−0.185, 0.025]	[−2.038, 0.325]	0.370	0.370	0.370	613
	Money $<$ median	[1.112, 3.475]	[−0.400, 0.275]	[−7.325, 0.775]	0.830	0.830	0.830	616
1990–2001	Money $>$ median	[0.052, 0.438]	[0.390, 0.490]	[−2.090, −0.900]	0.220	0.215	0.215	2,730
	Money $<$ median	[0.510, 0.912]	[0.100, 0.220]	[−1.770, 0.128]	0.440	0.440	0.440	2,733

Note: The estimation in this table is based on the baseline information set $Z_{i,t} = \{S_{i,t}, ShareMfg_{i,t}, \theta_i\}$.

that money in politics may target politicians with some type of characteristics, but ultimately the confidence sets do not point to substantial differences.

Second, we augment our specification with campaign contributions. It has to be noted that adding elements to the vector of parameters within the moment inequality approach is extremely costly due to the grid search process necessary for inference and hypothesis testing. Further, contributions could be endogenous to NTR votes, and hence it may be difficult to interpret the additional parameters. Online Appendix Table E.2 reports the results. Due to the computational burden, we consider the following specifications, each with four parameters to estimate and information sets to assess (a) baseline specification + campaign contributions from business organizations (panel A); (b) baseline specification + campaign contributions from labor unions (panel B). Our main results appear robust to these additional controls.

Third, we explore the congressional committees closer to the policy decision and likely the most important targets for special interests.[36] In particular, we exclude from the analysis politicians working in influential committees (including the Commerce and Ways and Means committees), who may be more influenced by money in politics. As is reported in online Appendix Table E.3, our results remain robust to this approach.

E. *Validation: NTR Votes for Vietnam and Votes on Other Foreign Policies*

As validation of our model and methodology, we briefly compare the results for the analysis of the NTR votes for China to a set of similar, but distinct NTR votes for the case of Vietnam. The goal here is to establish comparability in the responses of politicians across sets of votes. We view this as a way of supporting external validity of our findings. The analysis covers the votes on Vietnam's Jackson-Vanik Waiver (necessary to extend Vietnam's NTR status) that took place over the period 1998 to 2002. Based on a report by the Congressional Research Services, disapproval resolutions were not introduced in 2003, 2004, or 2005 (i.e., there is no voting data over

[36] We obtain the list of members of the US Congress and their committee assignments from Nelson (2018) and Stewart III and Woon (2016).

2003–2005). In 2006, the House passed legislation to grant Vietnam PNTR status as part of a more comprehensive trade bill and Vietnam accessed the WTO in 2007.[37]

In Table 8, the coefficient δ appears larger in magnitude than for the China shock case. Specifically, the baseline confidence set is $[-74.075, -6.700]$. However, this is because, as expected, the magnitude of the import shock from Vietnam is several orders smaller than that from China. The standard deviation of future import shock from Vietnam is 0.006, while that from China in the same period is 0.129. Adjusted for this scaling, the economic significance of constituency interests appears small for the Vietnam case as well, consistent with our findings of a low constituent weight estimated with the China NTR votes.

Concerning information sets, Table 8 confirms that members of Congress were informed about the impact of Vietnamese imports to some extent, as they were for China.[38] We cannot reject at standard significance levels the baseline information set, however, we reject that politicians have perfect foresight. Overall, across the sets of NTR votes for Vietnam and China, we do not find salient differences both in terms of economic magnitude of constituent weights and information sets.

As a final placebo check, we investigate the impact of the China shock on congressional voting on other foreign policies. In particular, we consider the annual Foreign Operations appropriations bill, which is the primary legislative vehicle through which Congress reviews the US foreign aid budget. The appropriations bills introduced over 1990–2001 cover a wide range of foreign aid programs across different geographic regions. Hence, we conjecture, the voting decisions are unlikely to be driven by constituents' considerations related to the China shock and this should be picked up by our methodology. This is the case. In online Appendix Table E.4, we find that the confidence sets of δ always contain zero across different periods, consistent with the conjecture that the China shock should have little influence on policies unrelated to issues on trade or local markets.

F. *Further Robustness*

Our baseline empirical model of voting falls in the class of "expressive voting" models, where politicians do not incorporate the likelihood of the pivotality of their vote choice on the passage of the entire bill nor voters punish or reward politicians for the passage or failure of the bill at the polling station. There are two main reasons for this modeling choice. First, it is empirically relevant, as politicians routinely campaign on their individual vote choices, and attack each other based on their respective individual voting records, rather than on the outcome of specific roll calls. Second, it is also a realistic assumption for decision making in the US House where the set of agents has large cardinality. In our context, none of the votes on NTR bills were decided by a single vote. That being said, it is interesting to engage with an analysis that incorporates pivotality and evaluate the robustness of the baseline findings to the alternative modeling choice. Online Appendix E.E2 proposes a sim-

[37] See online Appendix D.D2 for additional details on the data. Due to the congressional redistricting in 2002, for the analysis in this section, we only include the bills over 1998–2001. The past and future import supply shocks from Vietnam are constructed analogously to the China shock in Section IIIB.

[38] The manufacturing employment share and past import shock also have significant explanatory power in predicting future shocks for the case of Vietnam. The R^2 is generally around 53 percent.

Table 8—Parameter Confidence Sets and Specification Test p-Values: Votes on NTR with Vietnam, 1998–2001

CS of a	CS of b	CS of δ	p-value BP	p-value RC	p-value RS	Num obs.
Panel A. Minimal information $Z_{i,t} = \{ShareMfg_{i,t}, \theta_i\}$						
[−1.300, −0.300]	[0.600, 0.938]	[−85.000, −21.625]	0.990	0.990	0.990	1,595
Panel B. Baseline information $Z_{i,t} = \{S^V_{i,t}, ShareMfg_{i,t}, \theta_i\}$						
[−1.250, −0.613]	[0.630, 0.792]	[−74.075, −6.700]	0.435	0.435	0.435	1,595
Panel C. Perfect foresight $Z_{i,t} = \{S^V_{i,t}, ShareMfg_{i,t}, S^V_{i,t+1} - E[S^V_{i,t+1} \mid S^V_{i,t}, ShareMfg_{i,t}, \theta_i], \theta_i\}$						
–	–	–	0.025	0.020	0.020	1,595

Note: For the case of perfect foresight, we assume that in addition to $S^V_{i,t}$, $ShareMfg_{i,t}$ and θ_i, politicians also possess information that is orthogonal to these covariates, i.e., $S^V_{i,t+1} - E[S^V_{i,t+1} \mid S^V_{i,t}, ShareMfg_{i,t}, \theta_i]$.

ple pivotal voting model and reformulate the measure of trade shock accordingly. In particular, the alternative measure incorporates the "NTR gap" as defined by Pierce and Schott (2016) and takes into account that votes to revoke China's NTR status would be more effective at reducing imports from China for goods where the difference between the most-favored-nation tariff and the "column 2" tariff is larger. Our baseline findings remain robust in this environment.

Online Appendix E.E3 introduces export shocks and finds a positive effect of export opportunities on the probability of a vote in favor of China's NTR status. Yet, the information tests and the other estimates remain largely consistent with our main findings.

Online Appendix F.F1 discusses through Monte Carlo simulations the potential biases in our analysis due to violations of the rational expectation assumption. In particular, we consider the following scenarios: (i) constant under- or over-prediction of the China shock; (ii) expectational errors correlated with $S_{i,t}$; and (iii) expectational errors correlated with θ_i. In all these particular settings, the estimation based on moment inequalities and the specification tests appear to be robust to moderate violations of rational expectations. Exceptions typically emerge in cases where the variation in irrational expectational errors is larger than the variation in unpredictable components of the China shock (i.e., the component of $S_{i,t+1}$ that is unexplained by elements in $\mathcal{I}_{i,t}$).

V. Counterfactuals

In this section, we employ the estimated parameters of the model to perform two counterfactuals that answer the following questions: (i) would giving full information to legislators have changed the results of the NTR roll call votes? (ii) would the results of the NTR votes have changed if legislators had placed a larger weight on the labor market consequences of the China shock?

A. *Legislators Receiving Information about the China Shock*

After establishing that, for most of the time period in our sample, politicians had less than perfect knowledge of the labor market consequences of China's export expansion, a natural question we can ask is whether providing more accurate information to politicians would have changed their vote, and the overall passage of certain bills. The answer, for the case of China NTR votes, is not substantially.

Denote by $\mathcal{N}_t$ the set of politicians in period t, and $\Pi^{+}(\omega_t, \mathcal{I}_{i,t}, \mathcal{N}_t)$ the share of votes in favor of China. We simulate $\Pi^{+}(\omega_t, \mathcal{I}_{i,t}, \mathcal{N}_t)$ under the case of baseline information $\mathcal{I}_{i,t}^{b}$ and the case of perfect foresight $\mathcal{I}_{i,t}^{p}$. The change in share of votes (in percentage point) in favor of China when we provide politicians with full information is then given by[39]

$$\begin{aligned}(20)\qquad &\Big[\min_{\omega_t \in \Omega_t^{95\%}} \Big\{\Pi^{+}\big(\omega_t, \mathcal{I}_{i,t}^{p}, \mathcal{N}_t\big) - \Pi^{+}\big(\omega_t, \mathcal{I}_{i,t}^{b}, \mathcal{N}_t\big)\Big\} \times 100, \\ &\quad \max_{\omega_t \in \Omega_t^{95\%}} \Big\{\Pi^{+}\big(\omega_t, \mathcal{I}_{i,t}^{p}, \mathcal{N}_t\big) - \Pi^{+}\big(\omega_t, \mathcal{I}_{i,t}^{b}, \mathcal{N}_t\big)\Big\} \times 100\Big].\end{aligned}$$

Before we delve into the specific results, it is worth pointing out that the effect of information provision on NTR votes is ambiguous, and depends on (i) the underlying distribution of the expectational errors $G(\epsilon_{i,t+1})$, (ii) the weight on constituent interests that is governed by δ_t, and (iii) the policy position relative to individual ideology $a_t\theta_i + b_t$. Regarding (i), our assumption of rational expectations dictates that $\epsilon_{i,t+1}$ has a mean zero, conditional on $\mathcal{I}_{i,t}^{b}$. Moreover, as is shown in online Appendix Figure E.2, the expectation errors are more or less symmetrically distributed in different sample periods. In Figure 6, we illustrate the roles of (ii) and (iii) given a symmetric distribution of $\epsilon_{i,t+1}$. Point A represents the probability of casting a pro-China vote for legislators who are endowed with θ_i and have an expectation $E\big[S_{i,t+1} \,|\, \mathcal{I}_{i,t}^{b}\big]$. Note that in this scenario, the general policy position is in favor of China, and point A is located in the concave segment of $\Phi(\cdot)$. (This scenario is likely the case during 1993–1996 when the vote share in favor of China was above 70 percent on average.) When these politicians are supplied with complete information, they face import shocks $S_{i,t+1}$ that are dispersed around $E\big[S_{i,t+1} \,|\, \mathcal{I}_{i,t}^{b}\big]$ in a mean-preserving way. As is represented by point B, the share of pro-China vote will be lower in the counterfactual due to the concavity of $\Phi(\cdot)$ at this segment. The magnitude of the counterfactual change hinges on the dispersion of $\delta_t \epsilon_{i,t+1}$, which in turn depends on δ_t and the variation of $\epsilon_{i,t+1}$. In Section IV, we show that the effect of constituent interests reflected by the estimate of δ is moderate, and that the covariates in the baseline information set have significant explanatory power in predicting future shock especially in the earlier period (i.e., the variation of expectational errors is also moderate). These findings indicate that the effect of improving information could be small. Points C and D in Figure 6 depict the case in which general policy position is against China, and point C is located in the convex segment of $\Phi(\cdot)$. (This is related to the scenario in 1990–1992 when the vote share in favor of China was around 40 percent on average.) Under this case, had the politicians been provided with complete information on $S_{i,t+1}$, the pro-China vote share

[39] As the construction of the counterfactual is technical and not commonplace in the literature, we provide its details in online Appendix C.4.

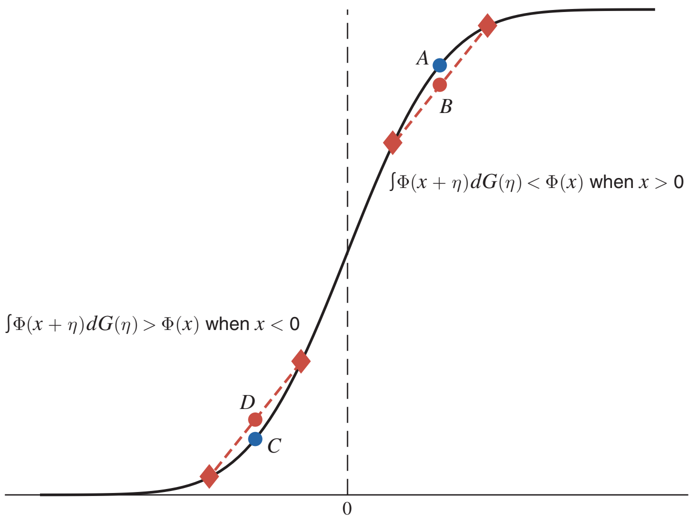

Figure 6. An Illustrative Example: Effects of Improving Information

Notes: The figure presents an illustrative example on how the provision of information on the China shock changes the pro-China vote share. The black solid curve represents the standard normal cumulative density function. In this figure, x corresponds to the component $a_t\theta_i + b_t + \delta_t E\left[S_{i,t+1} \,|\, \mathcal{I}^b_{i,t}\right]$ and η corresponds to the component $\delta_t \epsilon_{i,t+1}$. In this example, the distribution of η has a mean zero and is symmetric. The spread represented by the red dashed lines is determined by the variation of η.

would have increased. The magnitude of the counterfactual change again depends on δ_t and the variance of $\epsilon_{i,t+1}$.

We also simulate the share of politicians in the following categories: (i) vote for China in the baseline, but switch vote in the counterfactual, (ii) vote in favor of China in both the baseline and the counterfactual, (iii) vote against China in the baseline and switch in the counterfactual, and (iv) vote against China in both scenarios.

Table 9 shows the results of this counterfactual exercise. There are only a few small changes in the voting patterns across politicians who now possess full information, a result of the fact that they already possess a substantial amount of information to begin with and that the weight on constituent interests is moderate. Moreover, the vote share in favor of China's NTR status appears unaffected.[40] In summary, the evidence shows that lack of information on the part of members of Congress was hardly a driver of the particular outcome of the NTR votes.

B. *Counterfactual: Heightened Constituent Interests*

Another counterfactual that we explore here involves increasing the legislator's weight placed on his or her local constituents. Specifically, we apply to all politicians the lower bounds of δ_t corresponding to the confidence sets of two groups of politicians (i) Democrats and (ii) politicians with below median victory margins in the previous election, and then simulate the NTR votes. As is demonstrated

[40] Also note that in aggregate the counterfactual change in pro-China vote share in 1993–1996 is negative, while that in 1990–1992 is positive, which is consistent with the scenarios depicted in Figure 6.

Table 9—Effects of Improving Information

	1997–2001	1993–1996	1990–1992
1. Change in share of votes pro-CHN (%)	[−0.030, 0.012]	[−0.161, −0.000]	[−0.008, 0.064]
2. Share of always pro-CHN (%)	[55.945, 61.671]	[70.389, 76.967]	[36.626, 42.357]
3. Share of pro-CHN to against-CHN (%)	[0.077, 1.701]	[0.054, 1.618]	[0.000, 1.262]
4. Share of against-CHN to pro-CHN (%)	[0.078, 1.694]	[0.054, 1.486]	[0.000, 1.308]
5. Share of always against-CHN (%)	[36.736, 42.196]	[21.985, 27.637]	[56.268, 62.314]

Note: The simulation in this table is based on the assumption that the true information set is the baseline information set $\mathcal{I}_{i,t}^{b} = \{S_{i,t}, ShareMfg_{i,t}, \theta_i\}$.

Table 10—Effects of Heightening Constituent Interests

	1997–2001	1993–1996	1990–1992
Panel A. Baseline			
1. Value of δ	[−1.560, −0.075]	[−3.125, −0.125]	[−1.300, 0.000]
2. Share of votes pro-CHN (%)	[57.663, 61.760]	[71.999, 77.034]	[37.682, 42.427]
Panel B. Lower bound of CS for Democrats			
3. Value of δ	−4.425	−6.325	−3.700
4. Share of votes pro-CHN (%)	[17.485, 33.071]	[35.132, 58.356]	[11.983, 23.131]
Panel C. Lower bound of CS for win margin $<$ median			
5. Value of δ	−3.413	−8.550	−4.375
6. Share of votes pro-CHN (%)	[24.372, 42.458]	[24.606, 45.485]	[9.682, 19.424]

Note: The simulation in this table is based on the assumption that the true information set is the baseline information set $\mathcal{I}_{i,t}^{b} = \{S_{i,t}, ShareMfg_{i,t}, \theta_i\}$.

in section IVC, both these groups place a higher weight on the subconstituencies affected by the China Shock. By setting δ_t at the highest possible (absolute) values, we assess whether this margin of preferences may have played an important role in driving legislative outcomes for given expectations of the legislators.

Table 10 reports the confidence sets of the simulated vote shares. For all the simulations, we assume that politicians have access to the baseline information set and make voting decisions based on $E\left[S_{i,t+1} \mid \mathcal{I}_{i,t}^{b}\right]$. In panel A, we simulate the shares of pro-China votes based on the baseline estimates reported in Table 3. The predicted vote shares in favor of China align with the actual shares.

In panel B, we apply the largest possible weight that Democrats place on the expected China shock to all politicians. The counterfactual weight in 1997–2001 (respectively, 1993–1996 and 1990–1992) is nearly three times (respectively, two times and three times) larger than the lower bound of the baseline estimates of δ_t. The simulated vote share is in the range of $[17.485, 33.071]$ in 1997–2001, indicating that bills in favor of NTR with China would have not passed in this counterfactual scenario during this period. The only caveat is that the confidence set of the counterfactual vote share still straddles 50 percent for the period 1993–1996.

Panel C applies to all legislators the largest possible weight placed by the politicians facing high electoral pressure (i.e., winning margins below median). In this counterfactual, the passage of the bills in favor of NTR with China would have been overturned in periods 1993–1996 and 1997–2001.

VI. Conclusion

China's permanent NTR status in the United States, and its accession to the WTO, represent one of the most salient critical junctures in international trade (and certainly in trade policy decisions) of the last fifty years. This paper investigates whether US politicians had imperfect information about the extent of the China shock's repercussions in their home district at the time when they repeatedly voted on China's NTR status between 1990 and 2001.

To isolate the role of preferences versus information of members of Congress, we present a voting model and an application of a method of moment inequality approach designed to estimate expectations in decision making and to formally test for the content of information sets of legislators. We find that US legislators had imperfect, but fairly accurate expectations, yet placed a relatively low weight on the constituencies that ended up being adversely affected by the China shock.

The approach discussed in this paper may be a general and informative avenue for the study of policy decisions made by politicians and to understand the role of their expectations. Being able to resolve between preference and information margins is a valuable step forward for political economy and political science scholars interested in questions of both political accountability and learning in policy making. Future research should implement and extend our application outside the consequential questions of trade policy addressed in this paper. These could include quantifying the role of congress members' expectations for housing market reform ahead of the 2008–2009 financial crisis, and labor, healthcare, and fiscal policy in the 2010s.

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