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# Self-Enforcing Trade Agreements: Evidence from Time-Varying Trade Policy

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# Self-Enforcing Trade Agreements: Evidence from Time-Varying Trade Policy

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**Abstract**

The Bagwell and Staiger (1990) theory of cooperative trade agreements predicts new tariffs (i) increase with imports, (ii) increase with the inverse of the sum of the import demand and export supply elasticities, and (iii) decrease with the variance of imports. We find US import policy during 1997-2006 to be consistent with this theory. A one standard deviation increase in import growth, the inverse of the sum of the import demand and export supply elasticity, and the standard deviation of import growth changes the probability that the US imposes an antidumping tariff by 35%, by 88%, and by -76%, respectively.

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In an influential paper, Bagwell and Staiger (1990) develop a model of a cooperative trade agreement between two large countries.[1] They show that, in a dynamic, repeated trade policy-setting game, a cooperative trade policy equilibrium characterized by relatively low trade taxes can be sustained by the threat of infinite reversion to a Nash equilibrium of high trade taxes. Governments optimally choose low cooperative tariffs so that they can reap the benefits of greater trade. This cooperative equilibrium is characterized by a positive correlation between unexpected increases in import volumes and import tariffs. That is, when the import volume rises in response to an output shock, the lowest import tariff that governments can sustain as the cooperative equilibrium in the infinitely repeated dynamic trade policy game must rise. Our paper provides the first empirical investigation of the intertemporal and cross-sectional predictions of the Bagwell and Staiger model.

Import tariffs in the Bagwell and Staiger model generate terms-of-trade gains and thus vary intertemporally and cross-sectionally according to observable characteristics. The model first predicts that increases in an import tariff are more likely when import volumes increase. Second, conditioning on a positive import surge, the gains from (and thus the likelihood of) a tariff increase are rising in the inverse of the sum of the export supply and import demand elasticities. Thus, in the cross-section, a tariff increase is more likely for an import surge of a given size if import demand and export supply are more inelastic. Third, the gains from (and thus the likelihood of) maintaining a cooperative equilibrium with low trade taxes are increasing in the mean and variance of the underlying free trade volume. Therefore, conditional on an import surge of a given magnitude, an increase in a tariff for a product is more likely, the smaller is the variance of imports of that product in the cross-section.

To analyze the model, we use data on increases in US import tariffs against 49 countries under the US's antidumping and safeguard laws over the 1997-2006 period.[2] Although trade agreements like that embodied by

[1]Bagwell and Staiger (1990) helped influence a rich body of theory to understand the trade policy choices of countries that voluntarily submit to the rules of international trade agreements and their associated institutions (Bagwell and Staiger, 1999; Maggi, 1999; Ossa, 2011).

[2]We therefore examine whether antidumping and safeguard tariffs are consistent with the conditions under which trade volume shocks increase the incentive for a government to raise cooperative tariffs in order to continue participating in a self-enforcing trade agreement. Such an interpretation is consistent with Bagwell and Staiger (1990, p. 780, emphasis added), which states "[c]ountries can cooperatively utilize protection during periods of exceptionally high trade volume to mitigate the incentive of any country to

the World Trade Organization (WTO) require member countries to establish an upper limit on the tariff for their imported products, there are exceptions to WTO rules which allow governments to exceed those upper tariff limits under certain conditions. Antidumping and safeguards are two of the most important policies that major WTO economies use when they seek to implement *higher* import tariffs. Furthermore, these policies are economically important; e.g. the United States subjected 4-6% of its imported products at the 6-digit Harmonized System level to these policies during our sample period (Bown 2011; Prusa, 2011).

Our empirical results confirm a number of theoretical predictions from the Bagwell and Staiger model. In our baseline specification, we find that a one standard deviation increase in the recent growth of bilateral imports increases the probability of an antidumping tariff by 35%. We also find that the probability of an antidumping tariff increases as import demand and export supply become less elastic; a one standard deviation increase in the inverse of the *sum* of the import demand and export supply elasticities — the variable formally derived from the theory — increases the probability of an antidumping tariff by 88%. Finally, a one standard deviation increase in the standard deviation of import growth reduces the likelihood of an antidumping measure by 76%. Expanding our analysis to include safeguard tariffs as well as antidumping tariffs, we find that one standard deviation increases in these variables changes the predicted likelihood of a new time-varying tariff by 22%, 106% and -75%, respectively.

We investigate the robustness of our results to alternative explanations for time-variation in tariffs. In particular, we extend the empirical model to include political economy measures that have been widely utilized in the large literature on the use of antidumping and safeguard tariffs (e.g., Finger, Hall and Nelson, 1982; Feinberg, 1989; Knetter and Prusa, 2003; Crowley, 2011). We confirm that the quantitative importance of the key theoretical determinants generated from the Bagwell and Staiger (1990) model — import growth, import variance, and the trade elasticities — is similar to or greater than that of the traditional political-economy measures — industry concentration, employment, and inventory levels — that previous research has shown to be important determinants of these forms of time-varying tariff protection. Most significantly, inclusion of these political economy measures in our augmented empirical model does not affect our key findings.

unilaterally defect, and in so doing can avoid reversion to the Nash equilibrium. Thus, surges in the underlying trade volume lead to periods of "special" protection as countries attempt to *maintain some level of international cooperation.*"

The terms-of-trade motive for trade policy plays a critical role in the Bagwell and Staiger (1990) theory.[3] Our empirical investigation of these economic forces complements two other recent empirical contributions documenting how trade policy formation is determined by economic incentives in addition to political economy and income redistribution motives. Broda, Limão and Weinstein (2008) provide two pieces of evidence broadly consistent with the idea that countries exploit their market power in trade. First, they find that countries that are not members of the WTO systematically set higher tariffs on goods that are supplied inelastically. Second, they find that trade barriers on products not covered by the WTO agreement are significantly higher when the importing WTO member has greater market power. In a separate setting, Bagwell and Staiger (2011) focus on a set of countries newly acceding to the WTO between 1995 and 2005 in order to examine the role of market power in negotiating tariffs for new WTO members. They find evidence consistent with the theory of the terms-of-trade effect; the tariff to which a country negotiates is further below its non-cooperative level, the larger was its import volume before accession negotiations began.

The current paper contributes to this empirical literature by exploiting intertemporal and cross-sectional variation to explain government use of time-varying trade policies. In particular, we study how countries adjust their trade policies over time in response to shocks to trade flows and how these adjustments vary cross-sectionally according to industry structure. The earlier empirical literature has examined the cross-sectional variation in a country's *tariff level* (Broda, Limão and Weinstein, 2008) or the magnitude of a country's *tariff reduction* when moving from a non-cooperative policy to a trade agreement (Bagwell and Staiger, 2011). Our paper departs from this literature by focusing on one important WTO member's time-varying *tariff increases* in the face of trade volume shocks whose influence may vary due to heterogeneity across import demand and export supply elasticities.

The rest of this paper proceeds as follows. Section 1 briefly reviews the Bagwell and Staiger (1990) theory before introducing our empirical model of US antidumping and safeguard tariff determination. Section 2 presents a discussion of the data used in the estimation. Section 3 presents the estimates of the model of US tariff formation over the 1997-2006 period. Finally, section 4 concludes.

[3] Irwin (1996) provides a full account of the intellectual history of the terms-of-trade (or "optimal tariff") theory, which he finds dates back at least to Robert Torrens in the early nineteenth century. More recent treatments include the seminal work of Johnson (1953-1954).

# 1 Tariffs under a Cooperative Trade Agreement

## 1.1 The Bagwell and Staiger (1990) theory

Bagwell and Staiger characterize the most cooperative trade policy equilibrium in a two country partial equilibrium model of trade. In this model, stochastic output leads to fluctuations in the volume of trade over time that provide an incentive for countries to adjust the level of trade policy restrictiveness. We focus on the empirical predictions of Bagwell and Staiger's extension of their model which examines trade policy under more general import demand and export supply functions, $M(k^*, P^*)$ and $X(k, P)$. Specifically, $P$ is the (domestic) exporter's price, $P^*$ is the (foreign) importer's price, and $k$ and $k^*$ are general shift parameters such that $\partial M(k^*, P^*)/\partial k^* > 0$ and $\partial X(k, P)/\partial k > 0$. Letting $V^f$ designate the free trade volume of imports and exports, assume an increase in either shift parameter causes an increase in the volume of trade, i.e., $dV^f/dk^* > 0$ and $dV^f/dk > 0$. Bagwell and Staiger analyze the choice of a specific import tariff, $\tau^*$, and a specific export tax, $\tau$, where $P^* - P = \tau^* + \tau$ in equilibrium.

The national welfare for each country is defined as the sum of consumer's surplus, producer's surplus and tariff or tax revenue and can be denoted $W(k, k^*, \tau, \tau^*)$ for the domestic (exporting) country and $W^*(k, k^*, \tau, \tau^*)$ for the foreign (importing) country. The Nash equilibrium in the one-shot trade policy setting game is characterized by an import tariff, $\tau_N^*(k, k^*)$, and an export tax, $\tau_N(k, k^*)$, that are each inefficiently high.[4] Bagwell and Staiger use their stochastic output model to prove that, provided the discount factor is not too high, a cooperative equilibrium characterized by an import tariff, $\tau_c^*$, that is lower than the Nash equilibrium tariff and an export tax, $\tau_c$, that is lower than the Nash equilibrium export tax can be supported by the threat of infinite reversion to the Nash equilibrium in a dynamic infinitely repeated game.[5]

For the most cooperative equilibrium to exist, both countries must benefit from cooperation. The "no defection" condition requires that, for every possible volume of trade, the discounted present value of gains from cooperation to the foreign importing country, defined as $\omega^*(\cdot)$, exceeds the within-period gain of defecting from the cooperative agreement, defined as

[4] Our analysis focuses on the interior solution to the one-shot game with positive trade taxes. We rule out prohibitive trade taxes or taxes that reverse the natural direction of trade.

[5] A maintained assumption is that output follows an i.i.d. process. This, in turn, implies that trade volume shocks are i.i.d. Bagwell and Staiger (2003) describe a richer environment with serially correlated shocks.

$\Omega^*(\cdot)$.[6] If the incentive to defect, $\Omega^*(\cdot)$, increases, equation (1) implies that the cooperative trade policies, $\tau_c^*$ and $\tau_c$, must rise in order to maintain the inequality.[7]

$$\Omega^*(k, k^*, \tau_c(k, k^*), \tau_D^*(k, k^*, \tau_c(\cdot)) \leq \omega^*(\tau_c(k, k^*), \tau_c^*(k, k^*)) \tag{1}$$

Consider the special case of two countries that start from a most cooperative trade policy equilibrium of free trade, $\tau_c^* = 0, \tau_c = 0, P^*(\cdot) = P(\cdot) = P^f$. What incentive is there for the importing country to deviate from this cooperative policy? The gains to the importing country of defecting to a policy $\tau_D^*$ from a cooperative equilibrium of free trade can be written:

$$\begin{aligned}\Omega^*(k, k^*, 0, \tau_D^*) =& [P^f - P(k, k^*, 0, \tau_D^*)]M(k^*, P^*(k, k^*, 0, \tau_D^*)) \\ & - \int_{P^f}^{P^*(k,k^*,0,\tau_D^*)} [M(k^*, P^*) - M(k^*, P^*(k, k^*, 0, \tau_D^*))]dP^*\end{aligned} \tag{2}$$

Equation (2) states that if the importing country defects to its best response tariff, $\tau_D^*$, and the exporting country maintains a cooperative policy of free trade, $\tau_c = 0$, then the change in the importing country's welfare in the period in which it defects is equal to its terms-of-trade gain (the first term) less the efficiency loss associated with distorting the consumption price in its economy away from the free trade price and reducing the import volume to an inefficiently low level (the second term).

Further, Bagwell and Staiger have shown, by direct calculation, that the incentive to defect from a cooperative free trade equilibrium is increasing in positive shocks to trade volume if and only if the efficiency loss of the tariff policy is sufficiently small:

$$\frac{d\Omega^*(\cdot)}{dk^*} > 0 \text{ iff } \frac{\partial M(k^*, P^f)}{\partial k^*}\Big[\frac{P^f}{\eta_x^f + \eta_m^f}\Big] > \int_{P^f}^{P^*(k,k^*,0,\tau_D)} \frac{\partial M(k^*, P^*)}{\partial k^*} dP^*, \tag{3}$$

where $\eta_x^f$ is the export supply elasticity evaluated at free trade and $\eta_m^f$ is the import demand elasticity evaluated (positively) at free trade.

[6] Defection from the cooperative agreement by the foreign importing country consists of the importing country choosing its unilateral best response, $\tau_D^*(k, k^*, \tau_c(\cdot))$, to the domestic exporting country's most cooperative trade policy, $\tau_c(k, k^*)$.

[7] Symmetry implies a similar "no defection" expression for the exporting country.

Equation (3) provides the basis for the Bagwell and Staiger result that the most cooperative tariff increases in response to a positive import volume shock under fairly general conditions. Intuitively, if the most cooperative tariff fails to rise, the importing country will defect because the within-period gain from defecting exceeds the discounted present value of infinite reversion to the Nash equilibrium. This expression provides our first set of testable empirical predictions. An increase in import volume raises the incentive to defect provided that import demand and export supply are sufficiently inelastic, i.e., $1/(\eta_x^f + \eta_m^f)$ is large. Thus, the likelihood of a tariff increase rises with an increase in import volume. Moreover, equation (3) indicates that, for a given increase in import volume, $\partial M/\partial k^*$, the likelihood of a tariff increase is increasing cross-sectionally in the inverse of the sum of the import demand and export supply elasticities. For highly competitive sectors with highly elastic import demand and export supply, the inverse of the sum of the export supply and import demand elasticity will approach zero, providing no incentive to defect, even for large increases in import volume.

Next, we turn to the incentives to maintain cooperation. In any period, the gains to the importing country of maintaining cooperation can be written as:

$$\begin{aligned}\omega^*(\tau_c(k,k^*),\tau_c^*(k,k^*)) \equiv & \frac{\delta}{1-\delta}[EW^*(k,k^*,\tau_c(k,k^*),\tau_c^*(k,k^*)) \\ & - EW^*(k,k^*,\tau_N(k,k^*),\tau_N^*(k,k^*))]\end{aligned} \tag{4}$$

where $\tau_c^*(k,k^*)$ is the cooperative import tariff, $\tau_c(k,k^*)$ is the cooperative export tax, $\tau_N^*(k,k^*)$ is the Nash equilibrium import tariff, and $\tau_N(k,k^*)$ is the Nash equilibrium export tax. Equation (4) indicates the gains to cooperation are equal to the discounted present value of the difference between expected welfare under cooperative trade policies and expected welfare under Nash equilibrium trade policies. While the gains to a country of defecting from a cooperative agreement vary period-by-period with the realization of the within-period free trade volume, the discounted present value of the expected gains to maintaining a cooperative equilibrium $(\tau_c, \tau_c^*)$ is time-invariant.[8]

To develop empirical predictions, we consider the special case of the Bagwell and Staiger model with linear import demand and export supply,

[8] Because trade volume shocks are assumed to be i.i.d., expected welfare is time-invariant.

$M(k^*, P^*) = k^* - aP^*$ and $X(k, P) = k + aP$.[9] Further, we restrict our attention to symmetric trade policy functions in both the static and dynamic games, $\tau_N(\cdot) = \tau_N^*(\cdot)$ and $\tau_c(\cdot) = \tau_c^*(\cdot)$ . We present the cooperative trade policies as functions of the underlying free trade volume, $\tau_c(V^f) = \tau_c^*(V^f)$.

Starting with equation (4), direct calculation of the gains to cooperation, where punishment involves infinite reversion to the interior Nash equilibrium of the static game, yields:

$$\begin{aligned} \omega^*(\tau_c(V^f)) &= \omega(\tau_c(V^f)) \\ &= \frac{\delta}{1-\delta}\Big\{\frac{5}{12a}\big(\sigma^2_{V^f} + [EV^f]^2\big) - \frac{a}{4}\big(\sigma^2_{\tau^c} + [E\tau^c(V^f)]^2\big)\Big\} \end{aligned} \tag{5}$$

where $EV^f$ and $\sigma^2_{V^f}$ are the mean and variance of the underlying free trade volume and $E\tau^c(V^f)$ and $\sigma^2_{\tau^c}$ are the mean and variance of the cooperative tariff function. From equation (5), it is clear that the implications from Bagwell and Staiger regarding the stochastic output model are preserved in the special case of linear import demand and export supply. In particular, the expected future gains to cooperation are increasing in the mean, $EV^f$, and variance, $\sigma^2_{V^f}$, of the underlying free trade volume, holding the cooperative trade policy, $\tau_c(V^f) = \tau_c^*(V^f)$, fixed. Further calculation reveals the following rule for the most cooperative trade policy:

$$\tau_c^*(V^f, \omega^*) = \tau_c(V^f, \omega^*) = \begin{cases} 0 & \text{if } V^f \in [0, \overline{V}^f] \\ \frac{1}{2a}(V^f - \overline{V}^f) & \text{if } V^f \geq \overline{V}^f \end{cases} \tag{6}$$

where $\overline{V}^f = \sqrt{6a\omega^*}$, the cutoff value of trade volume below which the most cooperative policy is free trade.[10]

As in Bagwell and Staiger, equations (5) and (6) imply that, in the cross-section, a given increase in imports above the expected value will result in a higher cooperative tariff for the sector that has the smaller variance of imports.[11] In other words, an increase in the tariff is more likely when an

[9] For this special case, $V^f = (k + k^*)/2$ and $\tau_N = \tau_N^* = (k + k^*)/4a$.

[10] Note that while we treated $\omega^*$ as a constant for the purpose of calculating (6), $\omega^*$ is the function given in equation (5). Using the fixed point argument in Bagwell-Staiger, $\omega^*(\cdot)$ and, thus, $\tau_c^*(\cdot)$, can be expressed as functions of the model's exogenous parameters.

[11] The cross-sectional implications from the single sector model of Bagwell and Staiger (1990) come from equations (19) and (20) which together imply that the magnitude of the tariff increase is greater for sectors in which import surges are uncommon. For import surges of the same size in two different sectors, the magnitude of the tariff increase will be larger in the sector with the lower variance of imports.

import surge in a sector appears to be unusual. The final empirical prediction that we take to the data is therefore that a tariff increase is more likely in sectors in which the standard deviation of that sector's imports is lower. Nevertheless, it is worth highlighting that the interpretation that we adopt for our empirical specification below does rely on the single sector set-up of the Bagwell and Staiger model. Our approach implicitly assumes that, in a game played between countries with multiple sectors, the retaliation threat to deviation in a single sector is localized to that sector. Empirically, this assumption seems reasonable because governments incur non-trivial administrative costs in order to change tariffs and most retaliation threats made under the WTO system have been limited to small sets of goods.[12]

## 1.2 An empirical model of time-varying US tariffs

Our empirical strategy is to aggregate the comparative static predictions of equations (3), (5) and (6) into a single estimating equation. Equation (3) indicates that the incentive to defect will vary intertemporally with changes in import flows and cross-sectionally with the elasticities of import demand and export supply. In particular, the terms-of-trade theory implies that a change in imports will only affect the incentive to defect, and hence raise cooperative tariffs, if export supply and import demand are relatively inelastic. Thus, the empirical specification must allow for an *interaction* between imports and elasticities. Equations (5) and (6) together indicate that, in the cross-section, cooperative tariff increases will be more likely and/or larger in sectors with less volatile imports. Combining these predictions, we estimate the following equation:

$$y_{ikt} = \beta_0 + \beta_1 M_{ikt} + \beta_2 \Big( \frac{1}{\eta_{xk} + \eta_{mk}} \Big) + \beta_3 \Big( M_{ikt} * \frac{1}{\eta_{xk} + \eta_{mk}} \Big) + \beta_4 \sigma^m_{ik} + \varepsilon_{ikt}, \tag{7}$$

where $y_{ikt}$ is a measure of a trade policy change imposed against country $i$ for products of sector $k$ in year $t$, $M_{ikt}$ is a measure of the change in imports of $k$ originating from country $i$ in year $t$, $1/(\eta_{xk} + \eta_{mk})$ is the inverse of the

[12] More generally, in a multi-sector model in which the incentive constraints are pooled across sectors, the associated welfare loss due to the breakdown in cooperation could reflect the variance of trade volume aggregated across all sectors. We thank a referee for pointing out this possibility; we leave the empirical investigation for future research. Maggi (1999) is one theoretical approach that examines the pooling of incentive constraints in a multi-country, multi-sector model. However, his model focuses on multiple trading partners and emphasizes the role of multilateral cooperation, rather than extending a two-country, one-sector model to multiple sectors.

sum of the export supply and import demand elasticities for product $k$, and $\sigma_{ik}^m$ is a measure of the variance of imports of product $k$ from country $i$. We augment (7) to include the change in the bilateral real exchange rate between the importing country and country $i$ to control for aggregate relative price changes.

Empirically, changes in the incentive to defect can be interpreted as affecting the probability of a tariff increase or as determining the magnitude of a cooperative tariff increase. Our primary approach is to examine how variation in the data affects the probability of an antidumping or safeguard tariff across time, countries and industrial sectors. We report estimates from both a probit model and a logit model of tariff imposition. As a robustness check, we also use a censored Tobit model to determine the size of antidumping tariffs that are imposed, interpreting $y_{ikt}$ as an antidumping tariff.

## 2 Data used to estimate US tariff formation

We estimate the empirical model of US antidumping and safeguard tariff formation on a panel dataset constructed from several primary data sources: (1) trade policy data for the US come from the World Bank's Temporary Trade Barriers Database (Bown, 2010b), (2) US bilateral imports at the industry level come from the US International Trade Commission's DataWeb, (3) industry-level foreign export supply elasticities facing the US come from Broda, Limão, and Weinstein (2008), (4) industry-level US import demand elasticities come from Broda, Greenfield, and Weinstein (2006), (5) variables describing the characteristics of US domestic industries come from the US Census Bureau, and finally, (6) annual bilateral real exchange rates in foreign currency per US dollar come from the USDA Economic Research Service. Summary statistics for all variables in the dataset are reported in Table 1.[13]

The *ikt* panel includes 49 countries denoted $i$, 283 North American Industry Classification System (NAICS) 2007 manufacturing industries $k$ at the 5- or 6- digit level of aggregation, depending upon availability, for the years ($t$) 1997 through 2006.[14]

[13] Table 1 includes footnotes which describe how some variables are scaled by factors ranging from 1/100 to 1/10,000 prior to estimation. Our discussion of all quantitative empirical results fully accounts for this scaling.

[14] These 49 countries are: Argentina, Australia, Austria, Bangladesh, Belgium, Brazil, Canada, Chile, China, Colombia, Costa Rica, Denmark, Ecuador, Egypt, El Salvador, Finland, France, Germany, Greece, Hong Kong, Hungary, India, Indonesia, Ireland, Is-

The Temporary Trade Barriers Database provides detailed information on US antidumping and safeguard tariffs including the date a petition to restrict imports was filed, the identity of the country accused of dumping, the identity of countries included in the safeguard tariff, tariff-line information on the products involved, the outcome of the investigation and the magnitude of any final antidumping tariff imposed by the US against country $i$.

All tariff-line level (8- or 10- digit Harmonized System (HS)) trade policy data were concorded to 283 NAICS (2007 version) 5- and 6- digit US industries to merge into the $ikt$, foreign country-industry-year panel.

Industry-level foreign export supply elasticities facing the US at the 4-digit HS level were concorded to NAICS 5- and 6- digit industries. Because multiple 4-digit HS sectors can sometimes map into each NAICS industry, we record the median 4-digit HS export supply elasticity that maps into each NAICS industry as the elasticity for an industrial sector. Similarly, import demand elasticities at the 3- digit HS level were concorded to NAICS industries with the median elasticity in each industry used as the elasticity in the sector. To address a concern that some observations with extremely high or low import demand or export supply elasticities could be affecting our results, as a robustness check, we experiment with estimating our model on a smaller sample of data for which we drop observations in which either the inverse import demand or inverse export supply elasticity is in the top 5% or bottom 5% of the distribution of our primary estimation sample.

The leading alternative explanation for changes in tariffs over time is that political economy concerns lead governments to protect certain sectors of the economy. Other industry characteristics are frequently used in the literature to control for political economic determinants of an industry's propensity to obtain import protection. We follow Staiger and Wolak (1994) and Crowley (2011) in the choice of domestic industry characteristics to include in our analysis. Because a free-rider problem must be overcome in filing a request for import protection on behalf of the industry, more concen-

rael, Italy, Japan, Kenya, Malaysia, Mexico, Netherlands, New Zealand, Norway, Peru, Philippines, Poland, Portugal, Singapore, South Africa, South Korea, Spain, Sweden, Switzerland, Taiwan, Thailand, Trinidad, Turkey, United Kingdom, and Venezuela. Data on US manufacturing industries are available at the 5- digit level over the entire sample period. For some larger 5- digit industries, data is also available at the 6- digit level over the entire sample period. When the more disaggregated 6- digit industry data were available for all 6- digit industries within a 5- digit industry, we replaced the more aggregated 5- digit industry data with the less aggregated 6- digit industry data. Because we require two years of lagged data for our explanatory variable, we estimate the model on policy data from 1999-2006.

trated industries are thought to have a higher propensity to seek and to be awarded antidumping or safeguard tariff protection. Thus, we include the 4-firm concentration ratio (the shipments of the 4 largest shippers relative to total industry shipments). Further, we include a measure of industry size, total employment, because large industries may be better able to assume the large legal fixed cost of filing an antidumping or safeguard petition. Total employment also serves as a measure of an industry's political importance. The vertical structure of an industry may matter; upstream industries producing simpler commodities may file more petitions because they are more sensitive to industry price changes. We proxy for the vertical structure of an industry with the value-added to output ratio. Finally, because the current values of industry-specific variables may be endogenous to the antidumping or safeguard tariff, we use lagged values of these variables in estimating the model.

Furthermore, the WTO's Agreement on Antidumping and the WTO's Agreement on Safeguards specify empirical "injury" criteria that must be satisfied in order for a country to impose a special antidumping or safeguard tariff (Finger, Hall and Nelson, 1982; Feinberg, 1989; Knetter and Prusa, 2003; Crowley, 2011). In some specifications, we include the ratio of inventories to shipments to capture the WTO's injury criteria.

## 3 Empirical Results: US Tariff Formation

The empirical results reported in tables 2-4 provide evidence that the United States uses time-varying tariffs as predicted by the theoretical model of Bagwell and Staiger (1990). We examine 49 of the US's trading partners and find that the likelihood of an antidumping (antidumping or safeguard) tariff rises by 35% (22%) in response to a one standard deviation increase in bilateral import growth, rises by 88% (106%) in response to a one standard deviation increase in the inverse sum of the elasticities of export supply and import demand, and falls by 76% (75%) in response to a one standard deviation increase in a measure of the variance of import growth. Because the terms-of-trade theory describes the trade policy choices of large countries, we also report results for a sample limited to the US's top ten trading partners by import volume and find results that are quantitatively larger for some variables.[15] Analysis of the magnitude of antidumping tariffs also aligns with the theoretical predictions of the Bagwell and Staiger model.

[15]These 10 countries are Canada, China, France, Germany, Italy, Japan, Mexico, South Korea, Taiwan, and the United Kingdom.

Finally, we show that our results are robust to augmenting our empirical specification to include variables that have been widely used in the political economy literature on antidumping and safeguard policy.

We begin by describing the reported estimates of the binary model of the US government's decision to impose a final antidumping (or safeguard) tariff against country $i$ in industry $k$ after an investigation begun in year $t$. Estimates from a probit model are presented as marginal effects in which a one-unit increase in a variable is associated with an incremental increase in the probability that the US will impose an antidumping or safeguard tariff. From our estimating equation (7), the marginal effect of a change in bilateral import growth, $M_{ikt}$, on the probability of a tariff works through the *direct* effect of a change in this variable as well as *indirectly* through the interaction term, $\left(M_{ikt} * \frac{1}{\eta_{xk}+\eta_{mk}}\right)$. Thus, for each specification, we report only the *total* marginal effect of bilateral import growth as:

$$\frac{\partial Pr(y_{ikt} = AD|\boldsymbol{x})}{\partial M_{ikt}} = \phi(\boldsymbol{\beta'x})\Big(\beta_1 + \beta_3\Big(\frac{1}{\eta_{xk} + \eta_{mk}}\Big)\Big) \tag{8}$$

where we use the sample averages of $\boldsymbol{\beta'x}$ and $(1/(\eta_{xk} + \eta_{mk})$ in all calculations. $\phi(\cdot)$ is the standard normal density and is used in all probit specifications. Similarly, the marginal effect of a change in the inverse sum of the elasticities of export supply and import demand works through a direct effect and the interaction term. An analogous formula is used to calculate the marginal effect of a change in the elasticity measure.[16]

## 3.1 Baseline Results

Turn next to the results in Table 2, which analyzes the imposition of antidumping tariffs. Consistent with the theory, new US antidumping tariffs are more likely to be imposed when there has been a surge in past import growth, import demand and export supply are relatively inelastic, and import growth is less volatile.

Column (1) of Table 2 presents results for the basic specification of the model. First, the marginal effect of the growth of bilateral imports from

[16]For Table 2 specification (5) we report the marginal effects of the logit model and use $[1/(1+\exp(-\boldsymbol{\beta'x}))]*[1-(1/(1+\exp(-\boldsymbol{\beta'x})))]$ for the density $\phi(\cdot)$. For Table 3 specification (5) we report the coefficients from a Tobit model. Thus, $\phi(\cdot)$ is replaced with a 1 in calculating the interactions terms. The standard error of the marginal effect of a change in bilateral import growth on the probability of a tariff for the probit specifications is given by $\phi(\boldsymbol{\beta'x}) * (Var[\hat{\beta}_1] + Var[\hat{\beta}_3]\Big(\frac{1}{\eta_{xk}+\eta_{mk}}\Big)^2 + 2Cov[\hat{\beta}_1, \hat{\beta}_3])^{1/2}$. The logistic density is used in lieu of the normal density for calculating the standard error in Table 2 specification (5).

country $i$ in industry $k$ in the year before an antidumping petition is filed is estimated at 4.44 and is statistically different from zero. In our discussion of results for this model, we focus our interpretation on the increase in the probability above the mean value, calculated by multiplying the estimated marginal effect (e.g., 4.44) by a one standard deviation change in the explanatory variable (e.g., the lagged value of import growth of $0.947 \times 10^{-4}$, from Table 1). In this case, the growth of bilateral imports is associated with an increase in the probability of an antidumping tariff of 0.04 percentage points. In the bottom panel we use our estimated probit model to predict the probability of an antidumping tariff for a one standard deviation increase in bilateral import growth when all other variables are evaluated at their means. The predicted probability of 0.23% represents a 35% increase in the likelihood of an antidumping tariff relative to its mean value.

Our second result from specification (1) is that antidumping tariffs are more likely in sectors in which the export supply and import demand are relatively inelastic. Intuitively, when export supply is more inelastic, the terms-of-trade gain from a tariff is larger. When import demand is less elastic, the domestic efficiency costs of the tariff are smaller. Empirically, a one standard deviation increase in the log of the inverse sum of the export supply and import demand elasticities increases the probability of an antidumping tariff by 0.09 percentage points. In the lower panel, the predicted probability from the probit model for this change in the elasticity measure is 0.32%, an 88% increase in the likelihood of a tariff.

The other two explanatory variables in the baseline specification in Table 2 are the standard deviation of import growth and the percent change in the bilateral real exchange rate. The marginal effect on the standard deviation of import growth of -0.16 indicates that the likelihood of a tariff is decreasing cross-sectionally as import growth becomes more volatile. In other words, increases in trade protection are more likely for sectors in which an import surge is relatively unusual. A one standard deviation increase in the standard deviation of import growth reduces the probability of an antidumping tariff to 0.04% from a sample mean of 0.17%, a decline of 76%. Finally, a real appreciation of the US dollar increases the likelihood of an antidumping tariff. Quantitatively, a one standard deviation increase in the bilateral real exchange rate yields a modest increase in the probability of an antidumping tariff to 0.20%, an 18% increase relative to the mean in the sample. This finding is in line with previous work by Knetter and Prusa (2003) and Crowley (2011), all of which find evidence from other time periods that the probability of an antidumping tariff is higher when the real dollar appreciates.

Column (2) presents our first robustness check by using the inverse of the sum of the export supply and import demand elasticities instead of the natural log of its value. This is exactly the measure of market power used in the Bagwell and Staiger model without transforming the data for this variable to create a more normal-shaped distribution. All marginal effects have the same signs as those reported in column (1). Quantitatively, the predicted probabilities associated with a one standard deviation increase in each of the variables of interest are virtually identical to those reported in column (1).

Specification (3) provides a second robustness check to examine the sensitivity of the results to outliers in the distribution of import demand and export supply elasticities, a concern noted in Broda, Limão and Weinstein (2008). For this specification, we start with the estimation sample in column (1) and drop those observations for which the inverse import demand elasticity is in the top 5% or the bottom 5% of the distribution of inverse import demand elasticities and the observations that are in the top 5% or bottom 5% of the distribution of the inverse export supply elasticities. Restricting the sample in this way produces small increases in the magnitudes of the estimated marginal effects for all variables. This generates modest increases in the quantitative impact of each variable of interest on the predicted probabilities. A one standard deviation increase in import growth increases the likelihood of a antidumping tariff by 42% and a one standard deviation increase in the elasticity measure raises the probability of a tariff by 84%. Increasing the standard deviation of import growth by one standard deviation reduces the chance of a tariff by 79%. Lastly, the predicted probability of an antidumping tariff increases by 11% with a one standard deviation appreciation in the real exchange rate.

Table 2 column (4) focuses on the US's top ten trading partners by import volume. This is an important sample for examining the Bagwell and Staiger theory as their model describes the policy choices of large countries that are assumed be capable of influencing the terms-of-trade. In this sample, the likelihood of an antidumping tariff is more than two and a half times larger than in the full sample. For this sample, a one standard deviation increase in lagged bilateral import growth increases the probability of an antidumping tariff by 50%. This is modestly larger than the increase observed in the full sample of 49 countries. A one standard deviation increase in the elasticity measure increases the likelihood of a tariff by 59%. Increasing the standard deviation of import growth by one standard deviation reduces the likelihood of protection by 52%. Finally, the effect of an increase in the bilateral real exchange rate by one standard deviation is slightly larger among the top 10 trading partners; it increases the likelihood of an antidumping tariff by 33%.

The final specification of Table 2 examines the standard errors of our estimates by implementing the variance estimator of Cameron, Gelbach and Miller (CGM) (2011) in a logit model.[17] The CGM procedure constructs a variance estimator that allows two-way nonnested clustering. In our application, one might be concerned that errors are correlated with industry groups, $k$, and within country groups, $i$. The marginal effects reported in column (5) from the logit model are similar to the marginal effects from the probit model reported in column (4) and have no discernable quantitatively different effect on the predicted probabilities reported in the bottom panel of Table 2. However, the CGM variance estimator yields standard errors that are larger than the Huber-White robust standard errors reported for the probit specifications in columns (1) - (4). In terms of hypothesis testing, using the CGM standard errors, the marginal effect of the growth of imports is statistically significant at the 1% level, but the statistical significance of estimates on the natural log of the inverse sum of the export supply and import demand elasticities and of the bilateral real exchange rate declines to the 10% level. Using the CGM procedure, the estimate of the marginal effect of the standard deviation of import growth is no longer statistically different from zero.

### 3.2 Robustness Checks: Market Share, Safeguards, and China

Table 3 introduces a new explanatory variable to proxy for the unexpected import surge in the Bagwell and Staiger model. Some of the papers in the literature on the terms-of-trade theory of trade agreements (Bagwell and Staiger, 1999; Ossa, 2011) emphasize the importance of the market access implied by a negotiated tariff rate over tariff rates and import volumes. In mapping the repeated static environment of Bagwell and Staiger (1990) to an empirical environment characterized by domestic economic and trade growth, in Table 3 we use country $i$'s *share* of the importing country's market as our measure of expected import volume. From this, we define an import surge at $t-1$ as an increase in country $i$'s share of the US's market for $k$ between $t-2$ and $t-1$.

The first column of Table 3 reports our basic specification using the market share variable in lieu of the import growth measure. The results are consistent with those of the baseline specification (1) of Table 2. A one

[17] Judson Caskey provided the STATA code for the CGM variance estimator in a logit model.

standard deviation increase in a country's change in US market share at time $t - 1$ increases the probability of a US antidumping tariff by 18%. A comparison of the estimates for the impacts of the other variables included in the column (1) specifications of both Table 2 and Table 3 reveals that they are virtually identical.

The remaining specifications in Table 3 explore the robustness of our results through additional sensitivity analyses. Specification (2) reports estimates on a subsample of data made up of the top 10 foreign sources of US imports during this period. It provides additional evidence that the estimated impact of these explanatory variables is economically important.

In specification (3), we redefine the dependent variable to allow our time-varying trade policy to reflect safeguard tariffs *in addition to* antidumping tariffs. While there were many fewer instances compared to antidumping in which the United States used its safeguard policy during this time period, a focus on antidumping alone does miss out on one particularly important trade policy change that took place. In 2002, the United States used its safeguard tariff to restrict imports of steel in product lines that covered roughly $5 billion in annual US imports. Inclusion of these steel safeguard tariffs and a few other US safeguard policy actions during 1997-2006 does not change the qualitative nature of our results. Compared to specification (2), the results reported in column (3) suggest a slightly larger impact (relative to the predicted probability at the means) of the elasticities, standard deviation of import growth, and real exchange rate.

Specification (4) presents an analysis of China, the most frequent target of US antidumping tariffs during this period (Bown, 2010a) and an increasingly important source of US imports. While China accounts for only 2.5% of the observations in our baseline sample of data, it is the target of 44% of US antidumping tariffs in our sample. With the exception of the variable capturing the change in US market share (for which the marginal effect is positive, though not statistically different from zero), the estimated marginal effects are of the theoretically-predicted sign and are statistically significant. Furthermore, as the lower half of Table 3 indicates, the economic magnitudes of their estimated impact on the probability of US tariff formation during this period are also sizeable.[18]

Finally, specification (5) redefines the dependent variable as the *size* of the imposed US antidumping tariff and re-estimates the model on the top 10

[18] Because the sample of data in specification (4) consists of only one trading partner, Huber-White robust standard errors should correct the variance estimator for correlated errors within industries. This is an alternative way to address a concern that correlated errors might be non-nested in both country groups and industry groups.

trading partner sample of data using a Tobit model that is censored at zero. To interpret the quantitative significance of the estimates of the Tobit model, we start with the observation that the mean value of the antidumping tariff in this sample, defined as ln(1 + antidumping tariff), is reported in Table 1 as 0.0030. The antidumping tariff is reported in percentage points, thus the value 0.0030 can be expressed as a mean tariff of 0.3%.[19] Using the estimated coefficient of 2800.09 in the top row of column (5), we find that a one standard deviation increase in country $i$'s market share leads to an increase in the dependent variable of 0.168. Adding this to the sample mean tariff and transforming yields an increase in the antidumping tariff rate of 18.39 percentage points associated with a one standard deviation increase in the change of country $i$'s US market share. A similar calculation finds that a one standard deviation increase in the natural log of the inverse of the sum of the export supply and import demand elasticities is associated with a 45.27 percentage point increase in the antidumping tariff. A one standard deviation increase in the variability of import growth reduces the antidumping tariff rate by 26.76 percentage points. Finally, a one standard deviation increase in the growth of the bilateral real exchange rate increases the tariff rate by 27.41 percentage points. In summary, the results from the Tobit model confirm the Bagwell and Staiger predictions regarding changes in the cooperative tariff.

### 3.3 Model Extensions: Domestic Industry Characteristics and Political Economy

Table 4 presents a final set of robustness checks in which we extend the baseline model to include additional industry level covariates that the previous literature has suggested are significant determinants of time-varying antidumping and safeguard tariffs. We first establish the benchmark by re-estimating the baseline model for the full sample of trading partners with the dependent variable now defined as an indicator for whether the United States implemented an antidumping or safeguard tariff. Specification (1) indicates that the size of the marginal effects for the variables motivated by the Bagwell and Staiger theory, as well as their estimated impact on the predicted probability of a new tariff, are consistent with the results found thus far.

Specification (2) extends the model by adding four new industry level

[19]Recall that most observations in our sample face an antidumping tariff of 0% while a small number of observations face large positive values. The mean tariff in the sample of the US's top 10 trading partners, conditional on a positive duty, is 116.7%

covariates, three of which also have intertemporal variation. The estimated impact of each variable for this sample of data and these trade policies is statistically significant and consistent with our expectations based on evidence from previous research - the probability of new tariffs is increasing in industry concentration, the number of employees in the industry, and the ratio of inventories to shipments, whereas the probability is decreasing in the ratio of value-added to shipments. Most relevant for our purposes is that inclusion of these industry-level covariates does not change the sign and the statistical significance, and it does not significantly affect the size of the estimated marginal effects for the main variables of interest. Furthermore, it also worth noting that one standard deviation changes to the variables motivated by the Bagwell and Staiger theory generate changes to the predicted probability of new import tariffs that are frequently of similar or greater magnitude than these political-economic covariates that have been the emphasis of the traditional literature. Specifically, a one standard deviation change to the elasticities increases the predicted probability of a new tariff by 97% to 0.63. Specification (2) indicates that the most economically important domestic industry covariate is employment; a one standard deviation change to the number of workers in the industry increases the predicted probability of a new tariff by 94% to 0.62.

Our final robustness check of Table 4 re-estimates specification (2) with the inclusion of sector-level indicator variables for industries which produce steel or chemicals products. While only 1.3% of the observations in our dataset are for the steel industry, 26.7% of the antidumping tariffs recorded in the dataset are in steel. Similarly, while only 2.3% of the observations in the dataset are of chemicals, 11.8% of the antidumping policies in the dataset are against chemical exporters. Nevertheless, the results presented in specification (3) indicate the determinants of new US antidumping and safeguard tariffs are robust to the introduction of special controls for these sectors. First, the positive coefficient on the steel (chemical) indicator is strong evidence in favor of new tariffs against exporters from these sectors that goes beyond the basic economic variables of the Bagwell and Staiger (1990) model; a discrete change from a non-steel (non-chemicals) to a steel (chemicals) industry increases the probability of an antidumping tariff by 4 percentage points (1 percentage point), a large effect given that the probability of a new tariff for a non-steel, non-chemical sector is less than 1 percent. However, even after controlling for these sectors, the estimates of the other marginal effects are mostly unchanged, suggesting that the basic results are not driven by observations from the steel and chemical industries. The sole exception is the reduced impact of the elasticities variable; after controlling directly for steel and chemicals in specification (3), a one standard deviation change to the elasticities increases the predicted probability of a new tariff by only 22% to 0.39. Nevertheless, even in this specification the result is economically important and statistically different from zero.

To conclude this section, a large literature has explored the political-economic determinants of US antidumping and safeguard tariff policy. Our paper is the first to develop an empirical model of US tariff formation in which antidumping and safeguard policies are treated as time-varying cooperative tariffs in a self-enforcing trade agreement. Our evidence is consistent with the Bagwell and Staiger (1990) theory, as we find that US antidumping and safeguard tariffs are more likely the larger is lagged import growth, the greater the increase in the exporter's share of the US market, the lower the variance of imports, and the less elastic are US import demand and foreign export supply.

# 4 Conclusion

Our paper generates supportive evidence for the Bagwell and Staiger (1990) model of self-enforcing trade agreements. More generally, we show that the theory of cooperative trade agreements provides an empirically useful framework for understanding important trade policies like antidumping and safeguard tariffs. Using data from 1997-2006, we find that these new US tariffs are consistent with an increase in the incentive to raise "cooperative" tariffs as in the Bagwell and Staiger (1990) model of self-enforcing trade agreements. This paper presents three pieces of evidence supportive of this theory: the likelihood of these new import tariffs is increasing in the size of import surges, decreasing in the elasticities of import demand and export supply, and decreasing in the standard deviation of import growth. A one standard deviation increase in each of these variables is economically important, changing the probability that these tariffs will be imposed by 35%, by 88%, and by -76%, respectively. Our results are robust to restricting our analysis to the US's top ten trading partners and to analyzing the imposition of antidumping and safeguard tariffs. The results provide empirical support for models of trade agreements that emphasize the importance of the terms-of-trade motive in tariff setting, and they complement other empirical research (Broda, Limão, and Weinstein 2008; Bagwell and Staiger, 2011) on trade policy formation.

This empirical investigation of US trade policy raises additional questions for future research. The use of antidumping and safeguard policies has proliferated since the early 1990s; currently these policies are frequently used by a number of major emerging economies in the WTO such as India, China and Brazil. This use has been especially endemic to the global economic crisis of 2008-10 (Bown, 2011). To what extent does the theoretical model of Bagwell and Staiger (1990) apply to these economies' use of time-varying tariffs, and what other roles might such policies play in supporting cooperative trade agreements between these economies in the WTO system? Finally, and perhaps most importantly, a more thorough understanding of the use of such policies would also better inform us as to the potential *limits* to cooperation between sovereign nations through trade agreements, an ongoing sticking point in trade negotiations.

Table 1: Summary Statistics: US Antidumping and Safeguard Tariff Imposition

| | Full sample | | Top 10 trading partners only | | China only | |
|---|---|---|---|---|---|---|
| | Mean | St. dev. | Mean | St. dev. | Mean | St. dev. |
| **Dependent Variables** | | | | | | |
| Antidumping (AD) tariff imposed | 0.0017 | 0.0418 | 0.0046 | 0.0675 | -- | -- |
| AD or safeguard tariff imposed | 0.0032 | 0.0562 | 0.0060 | 0.0770 | 0.0323 | 0.1768 |
| ln(1+AD tariff) | -- | -- | 0.0030 | 0.0547 | -- | -- |
| AD tariff conditional on a positive value | 89.7 | 94.4 | 116.7 | 104.0 | 161.5 | 99.4 |
| **Explanatory Variables** | | | | | | |
| Growth of imports_ikt-1† | 0.102 | 0.947 | 0.084 | 0.567 | -- | -- |
| Change in US market share_ikt-1^ | 0.000 | 0.004 | 0.001 | 0.006 | 0.005 | 0.012 |
| $\ln\left[1/\left(\eta_x^f + \eta_m^f\right)\right]$_k ‡ | -1.991 | 1.517 | -1.995 | 1.526 | -1.982 | 1.523 |
| $1/\left(\eta_x^f + \eta_m^f\right)$_k ^ | 0.241 | 0.170 | -- | -- | -- | -- |
| Standard deviation of import growth_ik^ | 0.723 | 0.660 | 0.378 | 0.435 | 0.393 | 0.425 |
| Percent change in real exchange rate_it-1‡ | 0.007 | 0.116 | 0.001 | 0.087 | 0.017 | 0.015 |
| ***Domestic industry variables*** | | | | | | |
| ln(Four firm concentration ratio)_k*‡ | 3.468 | 0.608 | -- | -- | -- | -- |
| ln(Employment)_kt-1*‡ | 10.377 | 1.029 | -- | -- | -- | -- |
| Value-added/Shipments_kt-1*‡ | 0.513 | 0.118 | -- | -- | -- | -- |
| Inventories/Shipments_kt-1*‡ | 0.129 | 0.063 | -- | -- | -- | -- |
| Indicator for industry k is steel* | 0.013 | 0.113 | -- | -- | -- | -- |
| Indicator for industry k is chemicals* | 0.021 | 0.144 | -- | -- | -- | -- |
| **Observations** | 82,341 | | 20,775 | | 2,075 | |

*These variables are based on only 81,943 observations. † Rescaled by a factor of $10^{-4}$ for estimation.
^ Rescaled by a factor of $10^{-2}$ for estimation. ‡ Rescaled by a factor of $10^{-3}$ for estimation.

Table 2: US Antidumping Tariff Imposition: Marginal Effects from a Binary Model using Import Growth

| | Baseline specification (1) | Substitute alternative elasticity measures (2) | Remove elasticity outliers (3) | Top 10 trading partners only (4) | Logit model with multiway clustering (5) |
|---|---|---|---|---|---|
| Growth of imports_ikt-1 | 4.44*** | 4.86*** | 5.66*** | 28.93*** | 27.58*** |
| | ( 1.55) | (1.75) | (1.63) | (8.59) | (9.69) |
| $\ln\left[1/\left(\eta_x^f + \eta_m^f\right)\right]$_k | 0.58*** | -- | 0.86*** | 1.36*** | 1.31* |
| | (0.14) | | (0.20) | (0.39) | (0.75) |
| $1/\left(\eta_x^f + \eta_m^f\right)$_k | -- | 0.36*** | -- | -- | -- |
| | | (0.05) | | | |
| Standard deviation of import growth_ik | -0.16*** | -0.18*** | -0.18*** | -0.54*** | -0.54 |
| | (0.02) | (0.02) | (0.03) | (0.16) | (0.45) |
| Percent change in real exchange rate_it-1 | 1.09** | 1.15** | 1.07* | 13.91*** | 12.05* |
| | (0.55) | (0.58) | (0.59) | (2.91) | (7.13) |
| **Observations** | 82,341 | 82,341 | 67,262 | 20.775 | 20,775 |
| **Log-likelihood** | -1002.19 | -998.17 | -857.30 | -582.18 | -582.23 |
| **Predicted probability of antidumping tariff, expressed in percent,† ...** | | | | | |
| ...at means | 0.17 | 0.17 | 0.19 | 0.46 | 0.46 |
| …for one standard deviation increase to growth of imports | 0.23 | 0.24 | 0.27 | 0.69 | 0.71 |
| …for one standard deviation increase to elasticities | 0.32 | 0.26 | 0.35 | 0.73 | 0.74 |
| …for one standard deviation increase to standard deviation of import growth | 0.04 | 0.04 | 0.04 | 0.22 | 0.22 |
| …for one standard deviation increase to real exchange rate | 0.20 | 0.20 | 0.21 | 0.61 | 0.60 |

Notes: Dependent variable is a binary indicator that a US antidumping tariff was imposed on exporting country i in industry k after an investigation initiated in year t. Probit model used to estimate all specifications except for the logit model used to estimate specification (5). Huber-White robust standard errors in parentheses, except for specification (5) which implements Cameron, Gelbach and Miller (2011) multiway clustering on industry and trading partner. ***, **,* indicate statistical significance of marginal effects at the 1%, 5% and 10% levels, respectively. †Predicted probabilities expressed in percent terms; e.g., 0.17 is a predicted probability of seventeen hundredths of one percent, or 0.0017.

Table 3: US Antidumping and Safeguard Tariff Imposition: Marginal Effects from a Binary Model using Change in Market Share

| | Substitute change in US market share for import growth (1) | Top 10 trading partners only (2) | AD and safeguard tariff policies‡ (3) | China only‡ (4) | Tobit model with dependent variable as ln(1+AD tariff) (5) |
|---|---|---|---|---|---|
| Change in US market share_ikt-1 | 5.48*** | 14.41*** | 15.82*** | 18.28 | 2800.09*** |
| | ( 0.87) | (2.80) | (3.13) | (22.67) | (553.22) |
| $\ln\left[1/\left(\eta_x^f + \eta_m^f\right)\right]$_k | 0.58*** | 1.35*** | 1.86*** | 6.76** | 244.10*** |
| | (0.14) | (0.42) | (0.48) | (2.95) | (77.78) |
| Standard deviation of import growth_ik | -0.15*** | -0.38*** | -0.60*** | -3.26*** | -73.25*** |
| | (0.02) | (0.12) | (0.15) | (1.06) | (24.64) |
| Percent change in real exchange rate_it-1 | 1.20** | 14.82*** | 22.50*** | 582.99*** | 2777.37*** |
| | (0.61) | (3.08) | (3.56) | (217.19) | (518.14) |
| **Observations** | 82,341 | 20,775 | 20,775 | 2,075 | 20,775 |
| **Log-likelihood** | -995.40 | -579.51 | -716.28 | -285.03 | -634.95 |
| **Predicted probability of antidumping (or safeguard)‡ tariff, expressed in percent,†** | | | | | |
| ...at means | 0.17 | 0.46 | 0.60 | 3.23 | -- |
| ...for one std. dev. increase to change in US market share | 0.20 | 0.56 | 0.72 | 3.44 | -- |
| ...for one std. dev. increase to elasticities | 0.32 | 0.72 | 0.99 | 4.38 | -- |
| ...for one std. dev. increase to std. dev. of import growth | 0.05 | 0.27 | 0.29 | 1.79 | -- |
| ...for one std. dev. increase to percent change in real exchange rate | 0.20 | 0.61 | 0.85 | 4.19 | -- |

Notes: Dependent variable for specifications (1) and (2) is a binary indicator that a US antidumping tariff was imposed on exporting country i in industry k after an investigation initiated in year t. ‡Antidumping or safeguard tariff indicator used as dependent variable in specifications (3) and (4). Probit model used to estimate all specifications except for the Tobit model (censored at zero) used to estimate specification (5). Huber-White robust standard errors in parentheses. ***, **,* indicate statistical significance of marginal effects at the 1%, 5% and 10% levels, respectively. †Predicted probabilities expressed in percent terms; e.g., 0.17 is a predicted probability of seventeen hundredths of one percent, or 0.0017.

Table 4: US Antidumping and Safeguard Tariff Imposition: Import Growth and Industry Effects

| | Antidumping and safeguard tariff policies | Add political-economy covariates | Add steel and chemical indicators | Predicted probability of antidumping or safeguard tariff for one standard deviation increase in each explanatory variable, expressed in percent† | | |
|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (1) | (2) | (3) |
| Growth of imports_ikt-1 | 6.11*** | 3.44*** | 3.34** | 0.39 | 0.38 | 0.39 |
| | (1.71) | (1.14) | (1.37) | | | |
| $\ln\left[1/\left(\eta_x^f + \eta_m^f\right)\right]$_k | 1.19*** | 0.71*** | 0.24*** | 0.66 | 0.63 | 0.39 |
| | (0.19) | (0.12) | (0.06) | | | |
| Standard deviation of import growth_ik | -0.25*** | -0.14*** | -0.16*** | 0.08 | 0.10 | 0.09 |
| | (0.03) | (0.02) | (0.02) | | | |
| Percent change in real exchange rate_it-1 | 4.91*** | 2.70*** | 2.75*** | 0.42 | 0.40 | 0.40 |
| | (0.65) | (0.48) | (0.43) | | | |
| ***Domestic industry variables*** | | | | | | |
| ln(Four firm conc. ratio)_k | -- | 0.25*** | 0.13 | -- | 0.38 | 0.35 |
| | | (0.10) | (0.11) | | | |
| ln(Employment)_kt-1 | -- | 1.04*** | 0.70*** | -- | 0.62 | 0.47 |
| | | (0.14) | (0.10) | | | |
| Value-added/Shipments_kt-1 | -- | -3.58*** | -1.08** | -- | 0.20 | 0.28 |
| | | (0.64) | (0.54) | | | |
| Inventories/Shipments_kt-1 | -- | 6.82*** | 4.98*** | -- | 0.47 | 0.41 |
| | | (0.98) | (0.75) | | | |
| Indicator for industry k is steel | -- | -- | 0.04*** | -- | -- | 0.33 |
| | | | (0.01) | | | |
| Indicator for industry k is chemicals | -- | -- | 0.01*** | -- | -- | 0.39 |
| | | | (0.00) | | | |
| **Predicted probability of antidumping or safeguard tariff, expressed in percent,† at means** | | | | 0.32 | 0.32 | 0.32 |
| **Observations** | 81,943 | 81,943 | 81,943 | | | |
| **Log-likelihood** | -1631.52 | -1512.05 | -1346.50 | | | |

Notes: Dependent variable is a binary indicator that a US antidumping tariff or safeguard was imposed on exporting country i in industry k after an investigation initiated in year t. Probit model used to estimate all specifications. Huber-White robust standard errors in parentheses. ***, **,* indicate statistical significance of marginal effects at the 1%, 5% and 10% levels, respectively. †Predicted probabilities expressed in percent terms; e.g., 0.32 is a predicted probability of thirty-two hundredths of one percent, or 0.0032.

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