# TEMPORARY TRADE BARRIER IMPLEMENTATION AND MARKET POWER: EVIDENCE FROM LATIN AMERICAN ECONOMIES

by

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# Abstract

This paper examines temporary trade barrier (TTB) implementation by 13 Latin American economies on a bilateral basis from 2000-2009 considering market power and import shocks. Additionally, we augment our analysis by including the effect of the presence or absence of tariff water on TTB implementation. We find evidence that market power and tariff water play an integral role in TTB implementation while import shocks do not. Using a probit model we estimate that a one standard deviation increase in market power and the absence of tariff water indicator increase the probability that a country imposes an antidumping tariff by 71 and 20 percent respectively, evaluated at their means. Interestingly, we do not find that import shocks have a significant impact on TTB implementation.

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# Chapter 1 - Introduction

From the 1950s to present international trade has been liberalized on many levels[1]. During the 1995 Uruguay Round of multilateral negotiations the General Agreement on Tariffs and Trade (GATT) was transformed into the World Trade Organization (WTO). At this juncture 112 new WTO members took further steps to liberalize international trade by decreasing their most-favored-nation (MFN) tariffs and by forming preferential trade agreements. Yet, even as many barriers to trade were continuing to be eliminated or reduced—especially tariff levels among developed countries—there remain two protectionist trade policy avenues available to developing economies.

The first, "temporary trade barriers" (TTBs), have emerged as new and significant obstacles to free trade, especially over the last 30 years. TTBs most notably consist of antidumping (AD) tariffs but may also appear in the form of safeguard (SG) tariffs and countervailing duties (CVs). In the 1980s TTB's were a protectionist tool primarily employed by four industrialized nations, the U.S., EU, Canada and Australia (Knetter and Prusa 2003). However since the creation of the WTO in 1995 the use of TTBs among developing countries has drastically augmented. Bown (2011) estimates that in the mid-1990s approximately 0 percent of imported products in developing countries were subject to TTBs but by the late 2000s up to 3 percent of products were subject to TTBs in major emerging countries such as Argentina and Brazil. Finger, NG and Wangchuk (2001) posit that this precipitous rise in TTB use is non-coincidentally aligned with open trade policy because TTBs serve as a "pressure valve" necessary for importing governments to offer occasional sector-specific protection as demanded by myriad political economy and

[1] The GATT and WTO have had eight successful rounds of negotiation. Beginning with 23 countries at the Geneva Convention in 1947 to the latest Doha round which began in 2001 and is ongoing. Currently there are 162 members and observers of the WTO.

macro-economic factors. Thus, these two competing phenomena, of broad trade liberalization and the subsequent rise of TTBs, creates a dichotomous institutional framework leading to many important research questions.

The second prevailing avenue of economic trade protection is principally only at the disposal of developing economies and is a direct function of WTO policies. WTO members do not directly set precise tariff levels, rather they negotiate legal upper-bound limits over which any tariff set will be in violation of the WTO policy. The difference between the applied most favored nation (MFN) tariff and the bound tariff is known as tariff water. In principle, tariff water provides flexibility for governments to protect domestic industries against negative trade shocks by simply exercising their market power in the international market, or to respond to domestic political economy forces. Contrary to TTBs, which are used by developed and developing economies alike, large developed economies like the U.S. have virtually no tariff water with which to behave non-cooperatively and have little variation over time in their applied tariffs. Conversely, developing economies have substantial tariff water in their schedules providing flexibility to potentially engage in protectionist trade policies. Note, the thirteen Latin American countries analyzed in this paper have an average level of tariff water of nearly 27 percent, illustrating a clear second avenue by which developing economies may respond to negative trade shocks.

Thus, this paper analyzes TTB implementation by Latin American countries when subject to negative trade shocks and given market power, as well as analyzing the role of the presence or absence of tariff water on TTB implementation. This approach is motivated by theoretical predictions and empirical examples. Consider the latter: In the early to mid-2000s Brazil imported approximately 20 distinct shoe[2] products per year from China at the HS-06 level. From 2000 to

[2] Shoe or gaiter products all with HS-03 stem 640.

2007, 93 percent of these products had average levels of tariff water of approximately 14 percent. Given estimates in Nicita, Olarreaga and Silva (2015) that tariff water exceeding a threshold of 1 percent is sufficient to transition from cooperative to non-cooperative behavior. Tariff water of 14 percent provides ample flexibility for Brazil to potentially engage in protectionist policies through adjustment of the applied MFN tariff. This transpired in 2008 when tariff water dropped to just 2.4 percent on average, with tariff water observed in only 3 of the 21 imported shoe products that year. Correspondingly with the new low levels of tariff water, Brazil implemented antidumping tariffs against two thirds of the shoe products imported from China in 2008. In 2009, tariff water continued to be observed at an average 2.7 percent for the shoe products while the antidumping measures remained in place. This example demonstrates the intriguing interplay, unique to developing countries, that exists between TTB implementation and tariff water.

There is substantial literature on the determinants of trade policy with respect to TTBs as well as applied MFN tariffs and tariff water, and we briefly discuss the pertinent highlights. In the case of TTBs, rigorous empirical analysis on the rise of TTB use and the determinants of TTB implementation has generally focused on developed countries and in particular the United States. Blonigen and Prusa (2001) conduct a literature review on the determinants and effects of antidumping use as a temporary trade barrier and the effects of significant rises in these protective measures. They cite numerous studies that estimate the determinants of U.S. AD filings over a large period of time (1958—1992). The literature they review indicates major observable variables that seem to be the fundamental drivers of AD petitions for the U.S.; such as import penetration, domestic industry employment and capital stock of a given industry.[3]

[3] In particular, Sabry (2000) analyzes US antidumping procedures and data to identify successful AD petitions. The paper finds that the import penetration ratio and concentration levels are major determinants of the decision to file an antidumping petitions for the United States. Additional studies include Finger (1981), Herander and Schwartz

Knetter and Prusa (2003) provide an additional investigation and empirical analysis of the rise and use of antidumping tariffs primarily focused on the U.S. They analyze how general macroeconomic factors and specifically fluctuations in real exchange rates affect the AD filing patterns of four major users of AD tariffs (U.S., Canada, EU and Australia) during the 1980s and 1990s. They find that real exchange rates and domestic real GDP growth both have statistically significant impacts on AD filings and that dollar appreciations propagate antidumping activity[4]. Notably, the vast majority of TTB analysis and antidumping analysis in particular, has focused on the United States or a few other major developed economies. This paper strives to add to the body of knowledge and empirical evidence on the determinants of temporary trade barrier use and trade liberalization analysis in Latin American countries.

With respect to tariff water and applied MFN tariffs there is a reasonable body of literature on choosing optimal tariffs and the fundamental role market power plays in tariff setting. Optimal tariff theory suggests that tariffs set non-cooperatively should be positively related to the importing countries' market power. Otherwise known as the terms-of-trade argument, this implies that if a country has market power they can pass off the cost of a non-cooperative tariff onto the weaker exporting nation by causing a decrease in the price of the exported good (Bagwell and Staiger, 2011). Thus, the greater the market power, the higher the tariffs.

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(1984), Feinberg and Hirsch (1989), Hanse (1990), Krupp (1994), Lichtenberg and Tan (1994), Furusawa and Prusa (1996), and Blonigen (2000)

[4] Leidy (1997) conducts a similar analysis to Knetter and Prusa (2003) and get analogous results but uses a smaller sample of U.S. aggregate filings. Whereas Feinberg (1989) which analyzes the impact of exchange rate movements on U.S. AD filings with respect to four major exporters (Brazil, Japan, Korea and Mexico), finds that a depreciation of the U.S. dollar with respect to a given foreign currency begets increased AD petitions. The proposed mechanism is that a foreign firm's exports to the U.S. are at a lower price in foreign firms dollars given a U.S. dollar depreciation thus leading to an increase in the probability of a finding injury and thus an AD filing.

Broda, Limao and Weinstein (2008) quantify the importance of the terms-of-trade motive in tariff setting. One of their key findings demonstrates how, prior to WTO accession, countries with market power systematically exploited their terms-of-trade motive in the form of higher import tariffs on goods in which they had a high degree of market power. This is imperative for our analysis. Although all countries in our sample set are WTO members and are restricted by bound tariffs, they maintain significant tariff water in their schedules allowing for policy flexibility and the interplay between tariff water and TTB implementation.

Furthermore Nicita, Olarreaga, and Silva (2015) provide rigorous analysis on the interaction between trade policies and tariff water as well as providing additional evidence supporting the importance of including tariff water in a trade policy setting analysis. Nicita et al. (2015) note that tariff water is observed in more than three quarters of WTO members' tariff lines and this figure is even higher for our Latin American country set. As such, we include in our analysis a country's ability, or not, to raise their applied MFN tariff rates. Moreover, it is possible that trade policy in Latin American countries may be a protective response to trade shocks as well as an exploitation of market power. Thus, in addition to more traditional analysis of the impetus of TTB use we also work to provide empirical evidence on the effect of the presence or absence of tariff water and market power on implementation of TTBs by Latin American countries. We do this by building on several key papers discussed below.

Bown and Crowley (2013) develop an empirical model which considers governments' use of antidumping and safeguard measures as time-varying trade policies given shocks to trade flows. Their empirical analysis provides evidence in support of the theoretical model constructed by Bagwell and Staiger (1990) that yields three predictions about the variation of import tariffs due to macroeconomic factors, taking into account the importing country's market power and efficiency losses related to protectionist measures. First, increases in import growth raise the probability of the implementation of a temporary trade barrier. Second, conditional on an increase in import growth the probability of an import tariff increases the more inelastic the export supply[5], i.e., the greater the market power of a given country. Third, conditional on a given increase in import growth, the lower the volatility of import growth, the higher the likelihood of the implementation of a temporary trade barrier.

Their primary empirical approach is to examine changes in the incentive to defect from cooperative trade policy by calculating the probability of an antidumping or safeguard tariff across time, countries and industrial sectors against U.S. imports. Their empirical results strongly support the three predictions that import growth, market power and import variance are significant factors in determining the implementation of TTBs. Their analysis focuses broadly on manufacturing and more specifically on steel and chemicals. In this paper, we consider their primary variables with adaptations to their empirical strategy as well as additional factors we believe contribute to the precipitation of TTB use in Latin American economies.

Bown and Crowley (2014) shift their focus from specific analysis of the U.S. economy to a broader perspective of how major developing economies conduct their trade policy from 1989-2010 and provide analysis on the impetus behind each countries' respective policies. They do not however attempt to explain the application of TTBs on a bilateral basis but rather they estimate the impact of country level fluctuations on trade policies. In our analysis we consider the use of TTBs on a bilateral basis from 2000-2009.

[5] Bagwell and Staiger (1990) also include the import demand elasticity in their discussion. However we believe that market power and trade shocks can be viewed as alternative theories of trade policy determination and adhere to the strict terms of trade argument in our analysis which focuses solely on market power.

More specifically, Bown and Crowley (2014) analyze the increase in implementation of TTBs and strive to discern whether the increased use of TTBs signals commitment to WTO trade policy or if it is a retreat from liberal trade policy. Their investigation centers on counter-cyclical trade policy changes for emerging economies. They show that given a decrease in domestic real GDP growth in a country the potential terms-of-trade gain from protectionist trade policy can exceed the cost, especially if the country is in a persistent recession, as proposed by Bagwell and Staiger (2003), and that countries act according to these incentives. Ultimately, they find that import protection via increased use of TTBs reacts counter cyclically to real GDP growth, as predicted. Declines in growth for both the home and foreign countries result in statistically significant increased uses of TTBs (Bown and Crowley, 2013).

Bown and Crowley (2014) do not account for market power in their analysis of emerging economies. This is important, as it is a well-established theory that terms-of-trade gains are related to a countries market power. Hence, this paper incorporates market power into the analysis of Latin American economies using estimates constructed by Olarreaga, Nicita and Silva (2015). Additionally, as indicated above, many developing countries have significant leeway in satisfying their WTO obligations while changing applied MFN tariffs[6]. Thus, it is plausible that Latin American countries are employing TTBs for market power reasons as well as considering levels of tariff water, in addition to reacting to trade shocks.

[6] Bown and Crowley (2014) also consider the relationship between the share of products legally bound by the WTO and TTB policies. They note a clear pattern; major emerging economies whose products are bound by WTO constraints disproportionately employ more TTBs. The authors consider a tariff within 10 percentage points of its legally binding rate to be "bound" in the sense that governments have little space in which to maneuver tariff levels to affect industry protection. They observe increased implementation of TTB protection when there is an increase in the number of imported products that face restrictions against raising MFN applied tariff rates, implying the demand for TTBs as a protectionist policy is a function of the proximity of WTO bounds.

Therefore in this paper, we consider variables used in the empirical strategy of Bown and Crowley (2013) and their underlying theoretical predictions on the influence of import shocks on the use of TTBs in Latin American economies. Additionally we consider the findings described in Nicita et al. (2015) and Bown and Crowley (2014) to incorporate the effect of market power and tariff water on the implementation of TTBs.

Our empirical results confirm multiple theoretical predictions and question others. In our primary econometric specification we find that a one standard deviation increase in market power increases the probability of an antidumping tariff by 71 percent relative to its mean. We also find that the probability of an antidumping tariff increases given a one standard deviation in our lack of tariff water indicator variable by 20 percent relative to its mean. However, we do not find evidence that an increase in import growth has a statistically significant impact on a countries decision to implement a temporary trade barrier. We also find that a one standard deviation increase in a measure of variance of imports yields a 19 percent increase in the likelihood of an antidumping measure relative to the mean. Additionally, to investigate the robustness of our results, we extend our empirical analysis to control for political economy factors as well as including countervailing duties or safeguards in addition to antidumping tariffs. In the following chapter we provide a description of our data and discussion of our data set compilation. Chapter 3 presents our empirical strategy. In chapter 4 we present our empirical results and we provide concluding remarks in chapter 5.

# Chapter 2 - Data

In this section we discuss the compilation of data sets that allow us to investigate the role of trade shocks, market power and tariff water in determining TTB use by Latin American countries. In order to empirically analyze the determinants of TTB use and their interaction with macroeconomic factors we use TTB data from thirteen countries: Argentina, Brazil, Chile, Columbia, Costa Rica, Jamaica, Mexico, Panama, Paraguay, Peru, Trinidad and Tobago, Uruguay and Venezuela[7]. The primary source for TTB data is the Temporary Trade Barriers Databases found on the World Bank webpage and constructed by Chad Bown. This is the most comprehensive and detailed database available for historical trade policy data. Antidumping tariff data was gathered from the Global Antidumping Database (GAD), countervailing duty data was gathered from the Global Countervailing Duties Database (GCVD), and safeguard tariff data was gathered from the Global Safeguards Database (GSGD) (Bown 2015).

Ultimately, the dependent variable in our econometric analysis is the count of HS-06 country $i$ imported products $k$ that are subject to a new temporary trade barrier against exporting country $j$ in time $t$ resulting in import protection. The TTB data are originally presented as cases from importing country $i$ against exporting country $j$ at time $t$ for product $k$ at the Harmonized System 08-digit (HS-08) level. The highly dis-aggregated HS-08 product level data must be retooled to the HS-06 product level because this is the most finely dis-aggregated level of data that can be compared across countries (Bown and Crowley 2014). Thus, using the product specific information provided in Bown (2015) we reformat all HS-08 digit product codes to concord to the HS-06 level. To avoid double counting, when cases that include import protection at the HS-08

[7] Not all thirteen countries have AD, SG and CV data. We use available data from Bown (2015).

digit level fall into the same HS-06 level as a previously imposed measure, we treat these observations as a single TTB implementation at the HS-06 level. With the data aggregated at the HS-06 product level we have a data set with case initiation dates and the TTB measure imposed for all products at the HS-06 level for importing country *i* against exporting country *j*. This process is repeated for each country for antidumping, safeguard and countervailing data.

Table 2.1 illustrates key observations from the aforementioned TTB data from 2000 to 2009 for the thirteen Latin American economies that constitute our sample set. Each importing country is listed in Column 1. Column 2 shows there were a total of 738 TTB cases initiated against unique products at the HS-06 level from 2000 to 2009 by the thirteen Latin American economies. Columns 3 through 5 list cases by antidumping, safeguards and countervailing duties respectively for each country. Note that the majority of cases are antidumping, followed by safeguards with minimal countervailing cases amongst the country set[8]. As illustrated in the table there is a wide range of TTB usage between these Latin American countries. Column 6 shows that Argentina on average implemented about 30 new TTBs per year from 2000-2009 whereas the next highest user on average was Chile at 12 new cases per year. Furthermore, six countries in particular initiate the vast majority of TTBs during this time period. Argentina, Brazil, Chile, Mexico, Peru and Venezuela account for over 91 percent of all TTBs implemented from 2000 to 2009.

[8] Safeguard measures apply against all importers. Antidumping measures and Countervailing duties are applied on a bilateral basis against specific products and exporters.

**Table 2.1[9]: Temporary Trade Barrier Descriptive Statistics**

| Country | Total number of new products from 2000-2009 subject to: | | | | Average TTB Implementation: 2000-2009 | Average Share of Products subject to new TTBs: 2000-2009* | Average Share of Products subject to all imposed TTBs in effect: 2000-2009* |
|---|---|---|---|---|---|---|---|
| | TTB | AD | SG | CV | | | |
| Argentina | 268 | 168 | 100 | 0 | 29.8 | 0.687 | 2.024 |
| Brazil | 90 | 89 | 0 | 7 | 10.0 | 0.207 | 0.745 |
| Chile | 109 | 3 | 106 | 0 | 12.1 | 0.274 | 1.635 |
| Colombia | 29 | 29 | 0 | 0 | 3.2 | 0.074 | 0.296 |
| Costa Rica | 3 | 2 | 0 | 1 | 0.3 | 0.009 | 0.052 |
| Jamaica | 8 | 4 | 6 | 0 | 0.9 | 0.028 | 0.191 |
| Mexico | 35 | 33 | 0 | 2 | 3.9 | 0.076 | 0.458 |
| Panama | 13 | 0 | 13 | 0 | 1.4 | 0.049 | 0.215 |
| Peru | 86 | 81 | 0 | 5 | 9.6 | 0.238 | 1.314 |
| Paraguay | 1 | 1 | 0 | 0 | 0.1 | 0.004 | 0.024 |
| Trinidad and Tobago | 6 | 6 | 0 | 0 | 0.7 | 0.021 | 0.143 |
| Uruguay | 4 | 4 | 0 | 0 | 0.4 | 0.013 | 0.116 |
| Venezuela | 86 | 1 | 84 | 1 | 9.6 | 0.218 | 1.704 |
| Total | 738 | 421 | 309 | 16 | 6.3 | 0.146 | 0.686 |

*Source: Author using data from Bown (2015)*
*Reported as percentages

Column 7 describes each specific importing economy's entire stock of HS-06 products with non-zero imports for a given year that are subject to newly imposed TTBs. From 2000-2009 Argentina subjected on average 0.7 percent of its yearly HS-06 imports to new TTBs, more than any other country in the sample set. The previous TTB analysis literature usually focuses on the U.S. in part, because it is considered one of the greatest users of TTBs by a simple count measure of TTBs implemented. However, the U.S. only imposes TTBs against 0.16 percent of its products per year which is significantly less than Argentina as well as multiple countries in our sample set.

[9] The total number of products against which there are TTBs is less than the sum of the AD, SG and CV columns. This is due to the fact that when there is a product that is subject to two or more TTBs this is counted only as a single incidence, to avoid double counting. There are a few instances when this occurs which explains the disparity.

Thus TTB implementation as a share of imported products for countries in our sample set is relatively large or at least in line with the U.S.

The final column provides information on the share of imported products subject to all imposed TTBs in effect over the time period 2000-2009 for our sample set. This was constructed with the set of HS-06 products using a non-redundant count of affected HS-06 products divided by the specific economy's entire stock of HS-06 products with non-zero imports for a given year. Notice that we only consider a TTB measure applied in year 2000 or after and if it remains in effect until the revocation date. Argentina subjected an annual average of over 2 percent of their yearly imports to TTBs from 2000-2009 followed by Venezuela, Chile and Peru, all with well over 1 percent of their total products subjected to TTBs.

The TTB data from Bown (2015) were merged with a detailed and comprehensive trade and market power dataset constructed and provided by Nicita et al. (2015) to form a panel dataset allowing for analysis between TTB incidence, applied MFN tariffs, import data, tariff water and market power. Table 2.2 provides a summary of the tariff, market power and import data for each of the countries in Column 1. Column 2 shows the average applied MFN tariff for each country over the time period 2000-2009. Costa Rica has the lowest applied MFN tariff at 6 percent and Mexico has the highest applied rate at 15 percent. The country group average nearly 11 percent. Column 3 shows the average WTO bound tariff rate from 2000-2009. Panama has the lowest WTO tariff bound at 23 percent and Trinidad and Tobago has the highest at 57 percent with the country group average at 36 percent. Column 4 shows the average level of tariff water for each country over the same time period where tariff water is defined as:

$$Tariff\ Water = WTO\ Tariff\ Bound - Applied\ MFN\ Tariff$$

Panama has the lowest average tariff water at 16 percent where Trinidad and Tobago has the highest average at 48 percent and the group average is 26 percent. It is important to note the ubiquitous presence of tariff water throughout the country set. As discussed in the introduction, Nicita et al. (2015) demonstrate the importance of tariff water in setting cooperative tariffs. The levels of tariff water present in our country set are clearly non-trivial across the sample set. In this paper we consider a tariff within 5 percentage points of its legally binding rate to be "bound" as opposed to the mark of 10 percentage points used by Bown and Crowley (2014) given results in Nicita et al. (2015) that changes to tariff levels can be observed when tariff water exceeds 1 percent.

**Table 2.2: Tariff, Market Power and Import Data Description 2000-2009**

| Country | Applied MFN Tariff | WTO Bound Tariff | Tariff Water | Share of Products Bound by WTO | Export Supply Elasticity | Market Power | Median Import growth |
|---|---|---|---|---|---|---|---|
| Argentina | 0.124 | 0.317 | 0.19 | 0.047 | 60.70 | 0.0165 | 6.644 |
| Brasil | 0.137 | 0.311 | 0.17 | 0.067 | 49.34 | 0.0203 | 13.15 |
| Chile | 0.066 | 0.251 | 0.19 | 0.000 | 68.93 | 0.0145 | 4.357 |
| Colombia | 0.123 | 0.412 | 0.29 | 0.000 | 66.77 | 0.0150 | 7.448 |
| Costa Rica | 0.058 | 0.428 | 0.37 | 0.037 | 102.0 | 0.0098 | 3.319 |
| Jamaica | 0.084 | 0.508 | 0.42 | 0.032 | 113.9 | 0.0088 | 6.586 |
| Mexico | 0.154 | 0.350 | 0.20 | 0.158 | 28.63 | 0.0349 | 0.483 |
| Panama | 0.077 | 0.233 | 0.16 | 0.348 | 104.1 | 0.0096 | 16.39 |
| Peru | 0.096 | 0.301 | 0.20 | 0.001 | 85.34 | 0.0117 | 7.695 |
| Paraguay | 0.118 | 0.329 | 0.21 | 0.079 | 126.4 | 0.0079 | 4.516 |
| Trinidad and Tobago | 0.084 | 0.566 | 0.48 | 0.015 | 126.6 | 0.0079 | 2.015 |
| Uruguay | 0.126 | 0.314 | 0.19 | 0.051 | 134.4 | 0.0074 | 0.411 |
| Venezuela | 0.130 | 0.356 | 0.23 | 0.058 | 67.31 | 0.0149 | 0.000 |

*Source: Author using data from World Bank's Wits at wits.worldbank.org and data provided by Nicita et al (2015)*

Column 5 shows the share of imported products by each country that have tariff water levels of 5 percent or less, i.e. products that are bound by the WTO. The entire sample has just 6 percent of their products bound by the WTO from 2000-2009, further indicating that the countries in our sample set retain significant flexibility to make trade policy decisions through the applied MFN tariff mechanism and the existence of pervasive tariff water.

Column 6 shows the export supply elasticity for each country. Uruguay has the highest export supply elasticity at 134 and Mexico has the lowest at 28 followed by Brazil and Argentina. The group average is 87. These results are as predicted, given that larger economies generally have smaller export supply elasticities. Thus expecting the greatest users to have lower export supply elasticities implies they will have greater market power. This is observed in column 8 where Mexico has the highest market power followed by all other major users of TTBs with the exception of Colombia which has the fourth highest market power but does not implement many TTBs. These estimates concur with the relative size of the economies. Although the average market power is not large in comparison with the U.S. or European Union, the standard deviation is substantial, and there are also myriad products for which the non-cooperative tariffs are above 10 percentage points, leaving ample room for policy setting[10].

The last column shows the median import growth for the country set from 2000-2009. Import growth for each year, *t-1*, is calculated using the following formula.[11]

$$import\ growth_{i,t-1} = \frac{(imports_{i,t-1} - imports_{i,t-2})}{imports_{i,t-2}}$$

Since antidumping measures require time to implement, it is unlikely that the causal effect of an import surge on antidumping measure implementation will play out within a single calendar year. Thus, using a lagged measure of import growth allows us to analyze the effect of a change

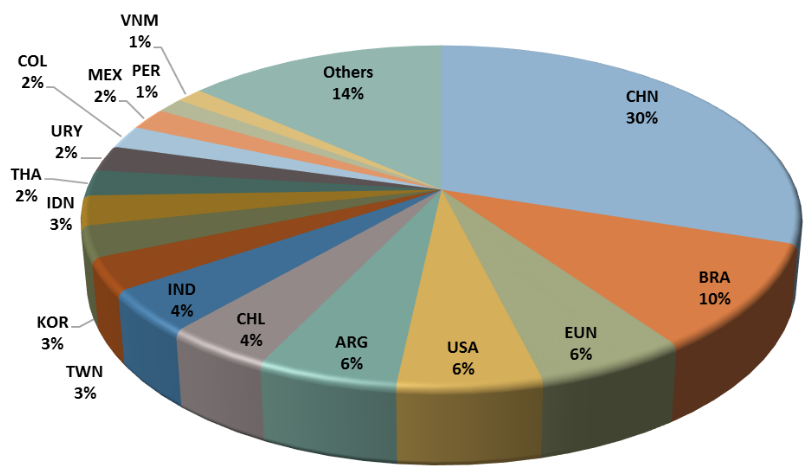
[10] As discussed in Broda et al. (2008) and Nicita et al. (2008) market power is related to economic size, the share of the importer in a particular market and product characteristics where imports of differentiated products result, on average, in greater degrees of market power.

[11] We report the median import growth as opposed the mean import growth from 2000-2009 because of the large variance in import growth from year to year. The median provides a more accurate description.

in the growth of imports on the following year's probability of an implementation of a new TTB measure against a specific HS-06 product. All countries exhibit positive median import growth from 2000-2009.

Lastly it is important to note that TTB import protection can be selectively imposed, targeting both specific countries and products. Thus TTBs are precise instruments of protection and can be concentrated against certain trading partners. We observe this in the data. Figure 2.1 provides information on the most affected exporters to the thirteen Latin American countries. Approximately 30 percent of all newly imposed TTBs from 2000 to 2009 by the thirteen Latin American economies were implemented against Chinese products. China is followed by 10 percent against Brazilian products, and 6 percent against European Union, US and Argentinian products. The rest of the newly implemented TTBs were against products from 46 countries. There is not a discernable pattern with respect to the remaining countries. They consist of developed and emerging economies across diverse geographic regions.

**Figure 2.1: Incidence of TTBs on Trade Partners**




*Source: Author using data from Bown (2015)*

Though our primary focus is on data at the HS-06 level, we also conduct the above analysis for all data at the HS-04 level. This allows us to consider our empirical analysis at multiple product levels to test the robustness of our results. To carry out these robustness checks, note that we use data on the elasticity of export supply estimated at the HS-04 digit level. The correlation between the elasticity estimates at the HS-06 and HS-04 levels is extremely high, at 0.92. Lastly, for our final robustness check we re-estimate several specifications on the HS-04 and HS-06 data sets that have been aggregated to the importer level.

# Chapter 3 - Empirical Strategy

In this chapter we describe our empirical strategy for analyzing how countries adjust their use of temporary trade barriers over time given market power, shocks to trade flows and tariff water. The inspiration of our econometric model comes from the econometric strategy employed by Bown and Crowley (2013). Our baseline specification is:

$$Y_{ijkt} = \beta_0 + \beta_1 \left(\frac{1}{e_{xik}}\right) + \beta_2 M_{ijkt-1} + \beta_3 \left[ \left(\frac{1}{e_{xik}}\right) \times M_{ijkt-1}\right] + \beta_4 \sigma_{ijk} + \varepsilon_{ijkt}$$

Where $Y_{ijkt}$ is a binary measure of a temporary trade barrier imposed by country *i* against country *j* in year *t* against product *k*. We principally consider the implementation of antidumping tariffs but also incorporate safeguard tariffs and countervailing duties into our later analysis. $\left(\frac{1}{e_{xik}}\right)$ is the inverse of the export supply elasticity for product *k* from country *i*, otherwise known as market power. As shown by the standard terms-of-trade argument, the greater the market power the greater the increase in tariff protection. Broda et al. (2008) and Nicita et al. (2015) empirically confirm these predictions. Thus, $\beta_1$ is expected to be positive, indicating that greater market power begets increased use of TTB measures.

$M_{ijkt-1}$ is a measure of the lagged growth in imports for product *k* to country *i* from country *j* in year *t*. The theory from Bagwell and Staiger (1990) suggests that surges in imports bring about strong incentives for importing countries to defect from cooperative trade policy and therefore increase the probability of a given country implementing protective trade instruments. Thus the expected sign of $\beta_2$ is positive. Given that the implementation of protective trade measures due to an import surge is also related to the elasticities of the countries involved, the variable $\left[ \left(\frac{1}{e_{xik}}\right) \times M_{ijkt-1}\right]$ allows for the interaction between import growth and elasticities. $\beta_3$ is predicted to be positive according to Bagwell and Staiger (1990). $\sigma_{ijk}$ is a measure of the standard deviation of imports of product $k$ by country $i$ from country $j$. Bagwell and Staiger (1990) posit that the more volatile the import growth in a given industry the more unlikely the imposition of a tariff against that industry, hence $\beta_4$ is predicted to be negative.

Given that $Y_{ijkt}$ is a binary variable, we report estimates using a probit model. We use values of market power provided by Nicita et al. (2015) to estimate the above baseline specification. Additionally, before estimating our baseline model, due to the fact that our sample set countries have both extremely high and low values of export supply elasticity we estimate our model on a data set where we drop the top and bottom 5 percent of the distribution of the primary sample to assuage concerns that extreme values are affecting the results. Thus, our baseline specification (1), is estimated over imported products $k$ from all industries, where $i$=sample Latin American country set, $j$=the set of exporting countries and $t$=2000—2009. Our baseline model has 1,122,820 observations.

We then augment the baseline specification to include only manufacturing products based on the NAICS 6-digit level of classification by concording the NAICS to our HS-06 digit level data set. This forms specification (2), which takes the same econometric form as specification (1) estimated over the manufacturing product only data set. Specification (2) has 1,083,006 observations.

Next, since we purpose to investigate the role of tariff water in the trade policy of Latin American Countries, in addition to import shocks we add a tariff water indicator variable to specification (2) to capture trade policy substitution over time and across instruments. Bown and Crowley (2014) analyze this effect by using a variable constructed by lagged change in the share of imported products under WTO discipline. However in this paper we take a different approach. We do not consider the share of products bound by the WTO, rather we analyze specific products and consider a product to be bound if its applied MFN tariff is within 5 percentage points of its WTO bound. Thus we introduce an *absence of water* indicator variable:

$$W_{ijkt} = \begin{cases} 1 \; if \; tariff \; water \leq 5 \; percentage \; points \\ 0 \; if \; tariff \; water \; > 5 \; percentage \; points \end{cases}$$

Where a value of "1" for any country *i* against country *j* and product *k* in time *t* acts as a proxy for the *lack of policy space* with respect to applied MFN tariffs in which a country could protect their industries by increasing applied MFN tariffs, indicating TTBs are their primary protectionist option. A value of "0" indicates there exists sufficient tariff water in which a country, in principle, could forgo the use of a TTB and simply raise their applied MFN tariff to impede a surge in imports. Thus we expect the coefficient of $W_{ijkt}$ to be positive.

Therefore our primary econometric model is specification (3), and has 1,035,535 observations:

$$Y_{ijkt} = \beta_0 + \beta_1 \left(\frac{1}{e_{xik}}\right) + \beta_2 M_{ijkt-1} + \beta_3 \left[ \left(\frac{1}{e_{xik}}\right) \times M_{ijkt-1} \right] + \beta_4 \sigma_{ijk} + \beta_5 W_{ijkt} + \varepsilon_{ijkt}$$

Lastly, given challenges to identifying the causal effect of import shocks and tariff water on the implementation of TTBs across time, industries and countries we control for multiple fixed effects. Notably, it is difficult to acquire reliable information on output levels, employment, capital levels, etc. for all of the developing countries in our sample set. Thus, we are unable to include myriad political economy factors in our analysis, yet we control for them by using industry, country and time fixed effects. Furthermore we expand upon our baseline and primary specifications by employing various additional robustness checks.

# Chapter 4 - Empirical Results

Tables 4.1, 4.2 and 4.3 present empirical results on the impetus of TTB for 13 Latin American countries. To simplify the exposition, we briefly summarize our results and then go over each Table in detail. We find that the likelihood of an antidumping tariff being imposed rises by a robust 71 percent in response to a one standard deviation increase in market power relative to its mean value. When including countervailing duties or safeguards in addition to antidumping tariffs, the response is a 70 percent and a 52 percent increase in the likelihood of TTB imposition, respectively. The magnitude of these results is significant and in line with Bown and Crowley (2013). Additionally, we find that a one standard deviation increase in the lack of tariff water indicator leads to a 20 percent increase in the likelihood of an antidumping tariff relative to its mean. When including countervailing duties or safeguards in addition to antidumping tariffs, the response is a 28 and a 20 percent increase in the likelihood of TTB imposition, respectively. These results are significant and agree with our prediction that given a lower level of tariff water, a country is more likely to implement a TTB.

We also find that a one standard deviation increase in bilateral import growth does not yield statistically significant estimates for antidumping tariffs. This is also the case for countervailing duties or safeguard tariffs in addition to antidumping tariffs. Lastly, we find that a one standard deviation increase in a measure of variance of imports yields a 19 percent increase in the likelihood of an antidumping measure relative to the mean. Including countervailing duties or safeguards in addition to antidumping tariffs, yields predictions 19 and 18 percent increases respectively. These results contrast with Bown and Crowley (2013).

Our collective results strongly suggest that the driving forces in trade policy setting for Latin American countries are market power and the absence of tariff water as opposed to import shocks. Finally, we demonstrate that our results are robust to augmenting our primary specification with additional variables as well as when controlling for country, industry and year fixed effects.

We present our empirical results in three tables. Table 4.1 presents our baseline probit model and the baseline with multiple extensions, such as accounting for manufacturing products, the lack of tariff water and controlling for country, industry and year fixed effects. In table 4.2 we present robustness checks at the HS-04 and HS-06 digit level of our baseline model and of our primary model incorporating various additional variables. In table 4.3 we estimate our model at the HS-04 and HS-06 digit levels on the original data sets aggregated to the importer level as a further robustness check.

The Tables 4.1 and 4.2 illustrate results in two ways. First in the top panel of each Table, the coefficients of probit estimates are presented where $\beta_i$ is the direct effect of an independent variable on the predicted probability of an increase in the likelihood of the dependent variable. For example, our baseline model has a positive coefficient $\beta_1$, indicating an increase in the probability of the implementation of antidumping tariffs given an increase in market power. Analogous analysis is used to interpret the coefficients of other independent variables.

Additionally, given the fact that there is interaction between market power and import shocks, $\left[\left(\frac{1}{e_{xik}}\right) \times M_{ijkt-1}\right]$, as well as direct coefficients, we need to consider the direct and the indirect effects of import shocks and market power on the probability of TTB implementation. To accomplish this we report the cumulative effects of the independent variables as the percent change in predicted probability of the implementation of an antidumping tariff (or safeguard or countervailing duty) given a one standard deviation increase in the independent variable with respect to the predicted probability evaluated at the mean.

To accomplish this we must first calculate the cumulative marginal effects accounting for both direct and indirect effects of market power and import growth. We calculate the cumulative marginal effects by:

$$\frac{\partial \Pr(Y_{ikt} \ = \ \text{AD}|\mathbf{x})}{\partial \text{M}_{idk}} = \phi(\boldsymbol{\beta}'\mathbf{x})\left(\beta_1 + \beta_3\left(\frac{1}{e_{xik}}\right)\right)$$

Where $\boldsymbol{\beta}'\mathbf{x}$ and $\left(\frac{1}{e_{xik}}\right)$ are evaluated at the sample mean and $\phi(\cdot)$ is the standard normal density used for all probit models. We then multiply the estimated cumulative marginal effect of $\beta_i$ by a one standard deviation change in the independent variable. Finally, we divide by the mean of the predicted probability of an antidumping measure or other TTB implementation, yielding an estimate of the percent change in predicted probability of TTB implementation relative to its mean.

### 4.1 Baseline Results

Turn now to the results in Table 4.1. Our first result is that the coefficient of market power is positive and significant indicating that an increase in market power leads to an increased probability of an antidumping tariff being imposed. In the lower panel of Table 4.1 column 1 note that a one standard deviation increase in market power evaluated with respect to the sample mean represents a 68 percent increase in the probability of an antidumping measure. This strongly indicates that market power plays an integral role in trade policy determination for our sample set.

The coefficient of import growth is positive and significant indicating that an increase in import growth leads to an increase in the likelihood of the implementation of an antidumping tariff. However, note that the coefficient of the interaction term is negative and insignificant. Thus, when analyzing the lower panel of Table 4.1 column 1, the net marginal effect of import growth after taking into account the indirect effect from the interaction term is not positive nor significant. This result is consistent across many specifications, signaling that Latin American economies are not responding to trade shocks as theoretically predicted or in line with previous empirical evidence from the U.S.

The coefficient of the standard deviation of import growth is positive and significant. The lower panel shows that when evaluated with respect to the sample mean a one standard deviation increase in the measure of variance of import growth leads to a 10 percent increase in the likelihood of an antidumping tariff. This result deviates from the theoretical predictions of Bagwell and Staiger (1990).

Column 2 of Table 4.1 re-estimates the baseline model over manufacturing products as defined by the 6-digit level of the North American Industry Classification System (NAICS) concorded with our data set at the HS-06 product level. The results are extremely similar to specification (1) with the only differences being slightly larger coefficients and modest increases in the percent change of predicted probability of an antidumping tariff implementation given a one standard deviation increase in market power and the standard deviation of import growth with respect to their means.

Column 3 presents our primary model; which consists of specification (2) and the addition of a lack of tariff water indicator variable. The results for specification (3) are virtually equivalent to those of specification (1) and (2) for the previously discussed independent variables. Our additional independent variable, the lack of tariff water indicator, has the predicted positive and significant coefficient, indicating an increase in the probability of an antidumping tariff given a decrease in tariff water.

**Table 4.1: Results HS-06 Antidumping Implementation: Coefficients and Predicted Probabilities from a Probit Model**

| | Baseline | Manufacturing Products Only | Add Tariff Water Indicator | Fixed Effects of Specification (3) | |
|---|---|---|---|---|---|
| | (1) | (2) | (3) | Industry X Country | Industry X Year |
| Market Power | 0.0878** | 0.0900** | 0.0909** | 0.0725* | 0.0854** |
| | (0.0322) | (0.0322) | (0.0325) | (0.0367) | (0.033) |
| Import Growth[a] | 0.454*** | 0.445*** | 0.417*** | 0.318** | 0.454** |
| | (0.112) | (0.113) | (0.117) | (0.121) | (0.174) |
| Import Growth X Marketpower | -8.120 | -8.094 | -7.693 | -2.508* | -2.806 |
| | (6.379) | (6.384) | (6.371) | (1.059) | (2.509) |
| Standard Deviation of Import Growth[a] | 0.0807*** | 0.143*** | 0.145*** | 0.147*** | 0.152*** |
| | (0.0102) | (0.0147) | (0.0147) | (0.0144) | (0.0136) |
| Lack of Tariff Water Indicator | — | — | 0.198*** | 0.321*** | 0.263*** |
| | — | — | (0.0590) | (0.0925) | (0.0574) |
| Observations | 1122820 | 1083006 | 1035535 | 1035535 | 1035535 |
| Prob > Chi-Squared | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| Percent change in predicted probability of TTB implementation given a one standard deviation increase in each independent variable with respect to the sample mean: | | | | | |
| Market Power | 68** | 70** | 71** | 61* | 73** |
| Import Growth | -111 | -112 | -107 | -28 | -27 |
| Standard Deviation of Import Growth | 10** | 19*** | 19*** | 19*** | 20*** |
| Lack of Tariff Water Indicator | — | — | 20** | 45*** | 37*** |

*Notes:* Robust standard errors are reported in parentheses. A probit model is used to estimate all specifications. Note that columns 4 and 5 estimate country, industry and time fixed effects. To control for fixed effects in columns (4) and (5) we re-estimated the model with a linear probability model. The estimates were analogous to using the probit model.

***Significant at the 1 percent level

** Significant at the 5 percent level

*Significant at the 10 percent level

[a] Rescaled by a factor of $10^{-5}$ for estimation

The lower panel of the table yields predictions in line with the previous specifications for the baseline model independent variables. For the lack of tariff water indicator we observe a 20 percent increase in the predicted probability of the implementation of an antidumping measure given a one standard deviation increase in the lack of tariff water indicator with respect to its sample mean. This estimate is significant and provides empirical evidence that tariff water levels below five percentage points indicate products are effectively bound by the WTO and countries are unable to increase their applied MFN tariff rate, thereby increasing the probability of a country implementing an antidumping measure for a bound product. This concurs with our prediction.

In column 4 of Table 4.1 we estimate our primary model, specification (3), over demeaned variables to control for both industry and country fixed effects given the dearth of available industry and political economy data for our country set. The coefficients for market power and lack of tariff water indicator are positive and statistically significant indicating that an increase in market power, or an absence of tariff water, leads to an increase in the probability of the implementation of an antidumping tariff. Similarly, in the lower panel of Table 4.1 a one standard deviation increase in market power and lack of tariff water indicator evaluated with respect to their sample means, represents a 61 and 45 percent increase in the probability of an antidumping measure respectively. These estimates confirm our earlier results while controlling for country and industry fixed effects.

The estimates for import growth and the standard deviation of import growth also closely dovetail with our primary results. The cumulative effect of import growth on the percent change in predicted probability of antidumping tariff implementation given a one standard deviation increase in import growth remains insignificant while a one standard deviation change in the measure of variance precipitates at 19 percent increase in the probability of antidumping tariff implementation.

In column 5 of Table 4.1 we present estimates of our primary model over demeaned variables to control for industry and time trend fixed effects. The results are analogous to our first robustness check as well as to our primary model for both coefficient estimates and predicted probability percent change evaluated at the mean given a one standard deviation change in the independent variables. They confirm the previous results that increases in market power and lack of tariff water indicator lead to robust increases in the predicted probability of the implementation of an antidumping tariff. While import growth and the standard deviation of import growth yield counter-theoretical predictions, further contributing to our results that market power is a major driver of Latin American trade policy.

Importantly all specifications in Table 4.1 have a miniscule p-values of less than 0.000 leading us to conclude that at least one of the coefficients of the regression is not equal to zero for all specifications. Additionally, there are over 1 million observations for each specification in Table 4.1.

## 4.2 Robustness

Table 4.2 introduces multiple robustness checks to our baseline and primary models. Column 1 of Table 4.2 presents estimates of our primary model, specification (3), for our data set retooled to the HS-04 digit level. The coefficient estimates of market power are positive and significant and are in line with those at the HS-06 level. The percent change in predicted probability of a newly imposed antidumping measure given a one standard deviation increase in market power with respect to the mean is a robust 87 percent, confirming our earlier results. The lack of tariff water coefficient is positive but not significant and the total marginal effect of a one standard deviation change evaluated at the mean is not statistically significant. Other coefficients and predicted probabilities are closely in line with previous estimates from Table 4.1, specification (3) at the HS-06 digit level.

Column 2 presents results from our primary model with the addition of a dummy variable for Chinese products. While Chinese products make up just 7 percent of all unique imported products by our sample set countries, nearly 51 percent of all antidumping measures are implemented against Chinese products. Brazil by comparison, is the next highest recipient of antidumping measures and is subject to just 12 percent of the antidumping measures employed. The coefficient estimates and predicted probabilities with respect to a one standard deviation change in the independent variables evaluated at the mean closely mirror those of the primary model, (Table 4.1, column 3) with the exception of slightly larger coefficients across the board. When controlling for Chinese products, the coefficient on market power remains positive and significant and the increase in the likelihood of an antidumping tariff evaluated at its mean is 67 percent given a one standard deviation increase, akin to previous estimates. The Chinese product indicator variable is statistically significant and positive indicating that if a product originates in China there is an increase in the likelihood of an antidumping measure. The percent change in predicted probability of an antidumping measure given a one standard deviation increase in the Chinese product indicator variable yields a 60 percent increase with respect to the sample mean.

**Table 4.2: Robustness results from probit models**

| | Specification (3): HS-04 | Specification (3) + China Indicator | Specification (3): AD and CV[b] | Specification (3): AD and SG[c] | Country Fixed Effects: Specification (3) HS-04 | | Specification (3): Brazil Only |
|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | Industry X Country | Industry X Year | (7) |
| Market Power | 0.0983*** | 0.103** | 0.0882** | 0.0587 | 0.0169 | 0.109*** | 0.166** |
| | (0.0282) | (0.0337) | (0.0322) | (0.0314) | (0.0443) | (0.0279) | (0.0610) |
| Import Growth[a] | 0.00506 | 0.483*** | 0.420*** | 0.245* | -1.198 | 0.455 | 2.109** |
| | (0.0852) | (0.126) | (0.119) | (0.115) | (1.599) | (0.681) | (0.694) |
| Import Growth X Marketpower | -0.217 | -15.83 | -8.424 | -0.486 | -6.398 | -5.585 | -10.49 |
| | (0.572) | (11.72) | (6.971) | (1.569) | (4.373) | (8.030) | (7.864) |
| Standard Deviation of Import Growth[a] | 0.0907*** | 0.127*** | 0.145*** | 0.141*** | 0.0906*** | 0.0940*** | 0.146*** |
| | (0.00836) | (0.0158) | (0.0145) | (0.0142) | (0.00847) | (0.00813) | (0.0378) |
| Lack of Tariff Water Indicator | 0.0274 | 0.238*** | 0.190** | 0.141* | 0.0438 | 0.0947 | 0.507*** |
| | (0.1000) | (0.0614) | (0.0589) | (0.0585) | (0.179) | (0.0865) | (0.0952) |
| Indicator of Chinese Products | — | 0.675*** | — | — | — | — | — |
| | — | (0.0352) | — | — | — | — | — |
| Observations | 539544 | 1035535 | 1035535 | 1035535 | 539544 | 539544 | 111781 |
| Prob > Chi-Squared | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| Percent change in predicted probability of TTB implementation given a one standard deviation increase in each independent variable with respect to the sample mean: | | | | | | | |
| Market Power | 87*** | 67** | 70** | 52* | 10 | 91*** | 105* |
| Import Growth | -4.5 | -227 | -107 | 3.1 | -167 | -104 | -138 |
| Standard Deviation of Import Growth | 12*** | 16*** | 19** | 18*** | 12*** | 12*** | 18*** |
| Lack of Tariff Water Indicator | 3.8 | 32*** | 28*** | 20** | 6 | 13 | 66*** |
| Indicator of Chinese Products | — | 60*** | — | — | — | — | — |

*Notes:* Robust standard errors are reported in parentheses. A probit model is used to estimate all specifications. Note that columns 5 and 6 estimate country, industry and time fixed effects. To control for fixed effects in columns (5) and (6) we re-estimated the model with a linear probability model. The estimates were analogous to using the probit model.

***Significant at the 1 percent level

** Significant at the 5 percent level

*Significant at the 10 percent level

[a] Rescaled by a factor of $10^{-5}$ for estimation

[b]Antidumping tariff or countervailing duty used as dependent variable in specification (3)

[c]Antidumping tariff or safeguard tariff used as dependent variable in specification (4)

In column 3 we re-estimate the primary model where the binary dependent variable is the implementation of either an antidumping measure *or* countervailing duty. The coefficients and predicted changes in probabilities given a one standard deviation increase in each independent variable are nearly identical to our primary model, specification (3) from Table 4.1, further confirming our initial results. In column 4 we re-estimate the primary model where the binary dependent variable is the implementation of either an antidumping measure *or* safeguard tariff. Column 4 also yields results generally in line with our primary model coefficients. The percent change in the predicted probability of an antidumping or safeguard tariff being implemented is 52 percent for a one standard deviation change in market power, modestly less than with specification (3) Table 4.1, but still coinciding with the general results.

Columns 5 is analogous to column 4 from Table 4.1 except it controls for industry and country fixed effects over the data set retooled to the HS-04 level from the HS-06 level. Coefficient estimates in column 5 generally agree with the consistent trends seen throughout our multiple specifications; however market power, import growth and lack of tariff water indicator variables do not have statistically significant coefficients. Only the standard deviation of import growth yields a statistically significant percent change in predicted probability of an antidumping measure.

Columns 6 is analogous to columns 5 from Table 4.1 except it controls for industry and time fixed effects over the data set retooled to the HS-04 level from the HS-06 level. The market power coefficient is positive and statistically significant. The percent change in predicted probability of an antidumping tariff given and one standard deviation increase in market power yields a robust result of a 91 percent increase with respect to the mean. These results concur with previous estimates on the effects of market power. Other estimates are in line with previous results with the exception of the lack of tariff water indicator variable, which does not yield a statistically significant estimate. Thus, when controlling for industry and time fixed effects our primary result of positive and significant market power prevails.

Column 7 considers our primary model, specification (3) from Table 4.1, with specific analysis of Brazil. Brazil is a significant user of TTBs; they subjected 2 percent of their products per year to new TTBs from 2000-2009 on average. Additionally they had an average of 17 percentage points of tariff water from 2000-2009 and had the second greatest market power in our sample set. Thus, Brazil's characteristics make it a unique candidate for specialized analysis. Estimates of the primary model generally concord with previous results. However the predicted increases in likelihood of an antidumping tariff due to a one standard deviation increase in market power and lack of tariff water yield more robust results of 105 and 66 percent increases evaluated at the mean. Given the specific country characteristics this accords with our predictions.

Finally note that columns 1, 5 and 6 with estimates on the HS-04 data set have over 500,000 observations. Columns 2, 3 and 4 that estimate on the HS-06 data set have over 1.3 million observations. Column 7, which considers only Brazil, still has over 100,000 observations. Lastly, all specifications in Table 4.2 have a miniscule p-values of less than 0.000 leading us to conclude that at least one of the coefficients of the regression is not equal to zero for all specifications.

In Table 4.3 we offer a further robustness check of our primary results. Previous models were estimated over bilateral data sets at the HS-06 and HS-04 digit levels although import shocks vary drastically amongst our country set. Thus, it is plausible that a given industry will not seek protection for any specific import shock by one exporter to an already volatile market. Rather it is possible that the industry may be more responsive given aggregate import shocks. To examine this potentiality we estimate our models on the data sets aggregated to the importer level as opposed to previous bilateral analysis. Columns 1 and 2 estimate our baseline and primary models while column 3 estimates our baseline model after controlling for industry and country fixed effects on the HS-06 data set aggregated to the importer level. Columns 4 through 6 estimate the same models on the HS-04 data set aggregated to the importer level.

**Table 4.3: Probit estimates over aggregated data sets**

| | HS-06 Baseline | HS-06 Specification (3) | HS-06 Baseline: Country x Industry Fixed Effects | HS-04 Baseline | HS-04 Specification (3) | HS-04 Baseline: Country x Industry Fixed Effects |
|---|---|---|---|---|---|---|
| Market Power[a] | 0.0848* | 0.0889* | 0.0597 | 0.186*** | 0.181*** | 0.151** |
| | (0.0370) | (0.0373) | (0.0409) | (0.0377) | (0.0381) | (0.0513) |
| Import Growth[b] | -0.927 | -0.661 | -474.2 | -8.837 | -9.353 | -7830.5 |
| | (2.090) | (1.838) | (907.1) | (6.478) | (6.642) | (5403.4) |
| Import Growth X | -2.476 | -2.458 | 112.2 | 19.23 | 21.23 | -59476.6 |
| Marketpower | (5.742) | (5.535) | (794.4) | (15.92) | (15.64) | (31253.4) |
| Standard Deviation of | 0.0577*** | 0.0578*** | 52.44*** | 0.0463*** | 0.0455*** | 45.36*** |
| Import Growth[b] | (0.00968) | (0.0102) | (11.09) | (0.00681) | (0.00684) | (7.926) |
| Lack of Tariff Water | — | 0.340*** | — | — | 0.188 | — |
| Indicator | — | (0.0691) | — | — | (0.122) | — |
| Observations | 264206 | 254344 | 264206 | 84565 | 84362 | 84565 |
| Prob > Chi-Squared | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 |

*Notes:* Robust standard errors are reported in parentheses. A probit model is used to estimate all specifications. Note that columns 3 and 6 estimate country, industry and time fixed effects. To control for fixed effects in columns (3) and (6) we re-estimated the model with a linear probability model. The estimates were analogous to using the probit model.
***Significant at the 1 percent level
** Significant at the 5 percent level
*Significant at the 10 percent level
[a] Rescaled by a factor of $10^{-3}$ for estimation
[b] Rescaled by a factor of $10^{-5}$ for estimation

Table 4.3 confirms the results in tables 4.1 and 4.2. Market power, with the exception of column three, yields results that are positive and significant indicating that even with respect to total imports, market power plays an important role in the implementation of antidumping tariffs. Additionally, at the HS-06 digit level the lack of tariff water indicator is both positive and significant, analogous to previous results. Notably import shocks, even when analyzed with respect to aggregate imports, mirror our earlier estimates presented in Tables 4.1 and 4.2. The coefficients are not significant, illustrating that across multiple levels of analysis import shocks are not significant determinants of TTB implementation amongst our country set. Furthermore, estimates of a one standard deviation change in the independent variables, evaluated with respect to their means are analogous to previous results in Tables 4.1 and 4.2.

# Chapter 5 - Conclusions

In this paper we take a unique approach to analyzing trade policy decisions by Latin American countries when subject to import shocks and given market power. Adapting Bown and Crowley's (2013) empirical strategy to include additional explanatory variables and variables unique to developing economies we provide new insights into trade policy decisions by Latin American economies. Using detailed temporary trade barrier data from Bown (2015) as well as comprehensive trade and market power data from Nicita et al. (2015) we present results from a probit model on the unique interplay between market power, tariff water and import shocks for 13 Latin American economies from 2000-2009.

First, we find strong supportive evidence for the prediction that an increase in market power leads to an increase in use of temporary trade barriers. Empirically our results show that a one standard deviation increase in market power increases the predicted probability of an antidumping tariff by 71 percent compared to the mean. Second, we find that when products do not have large amounts of tariff water, i.e. they are bound by the WTO, countries are more likely to implement temporary trade barriers against those products. Our results show that a one standard deviation increase in the lack of tariff water indicator increases the predicted probability of an antidumping tariff by 20 percent compared to the mean. Importantly, these two results are robust to controlling for industry, country and time trend fixed effects as well as the addition of countervailing duties or safeguard tariffs with antidumping tariffs. Furthermore we show that these results are robust to augmenting our empirical specifications to include additional variables.

Our third and fourth results are that a one standard deviation increase in bilateral import growth does not yield statistically significant estimates for a change in the likelihood of a country implementing antidumping tariffs and that a one standard deviation increase in a measure of variance of import growth yields a positive and significant increase in the probability of an antidumping tariff. These results hold when considering countervailing duties or safeguards in addition to antidumping tariffs, and for all fixed effects at the HS-04 and HS-06 data level. These two results are in contrast to evidence found by Bown and Crowley (2013). A potential explanation of the dissimilar results is that we are analyzing a fundamentally different set of countries. Bown and Crowley (2013) only consider the U.S.; a large and stable economy, while our country set consists of developing economies of varying sizes with volatile trade situations. Thus, there may exist such extreme volatility that a country is forced to react with TTB implementation. It is unclear that similar results from large and stable economies should be expected for country set with fundamentally different import flows.

Collectively our results yield an intriguing empirical contribution to the literature. The implication is that the overwhelming driving force in trade policy setting for Latin American countries during this time period was market power. Additionally, TTB implementation is augmented when countries have products bound by the WTO, suggesting tariff water plays an important role in trade policy setting among Latin American countries. Lastly, our results robustly suggest that import growth is not a significant driver of trade policy for our sample set.

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