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THE 'EMULATOR EFFECT' OF THE URUGUAY ROUND ON U.S. REGIONALISM*

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ABSTRACT

In this paper we investigate whether multilateral trade liberalisation undermines or encourages preferential trade liberalisation. We use highly disaggregated data at the tariff-line level and find an 'emulator effect' of multilateral trade liberalisation on preferential trade agreements. Indeed, our regression results show that products for which the U.S. agreed to cut its MFN rate substantially between the end of the Tokyo and Uruguay Rounds of GATT negotiations (1979-1994) are also the products for which subsequent tariff cuts on a preferential basis are boldest. In this sense, regionalism complements multilateralism. Our empirical investigations also reveal that sectoral differences are important determinants of the preference margin, as well as the identity of trading partners. When crossed with the results of Limão (2006) and others on the causal link between regionalism and multilateral liberalisations in the realm of the GATT/WTO institutional framework, our results establish that the multilateral trading system and the incentives to expand regionalism interact in complex ways.

JEL F13, F14, F15, N70 **Keywords**: Regionalism, Multilateralism, Stumbling bloc, Uruguay Round.

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1. INTRODUCTION

One striking feature of the current world trade system is the explosion of *regionalism*, that is, the growth in the number of preferential trade agreements (PTAs). Only 37 such agreements were in place at the launch of the World Trade Organisation (WTO) in 1994. In July 2007, 380 PTAs had been notified to the GATT/WTO and 205 of them were in force.[1] Though such numbers should be treated with caution for a variety of reasons (Baldwin 2007), it is beyond doubt that these numbers are growing. What is driving this growing proliferation of PTAs? Are regionalism and multilateralism substitutes or complements? Answering such questions is important, not least because several scholars fear that regionalism is a dynamic substitute, or *stumbling block*, to multilateral free trade and a menace to the multilateral trading system incarnated by the GATT/WTO (Bhagwati 1991, Levy 1997, Bagwell and Staiger 1998, Krishna 1998, McLaren 2002).

In this paper, we study the 'substitute vs. complement' question from the opposite angle: we assert, as Ethier (1998), that 'regionalism is an endogenous response to the multilateral trading system (Ethier 1998: 1216). Our research question is thus: is multilateralism driving the proliferation of PTAs in any way?

This question has received surprisingly little academic interest so far. To the best of our knowledge, Ethier (1998) and Freund's (2000a) theoretical papers are unique exceptions. Our paper studies this question from an empirical perspective, focusing on the United States.

Figure 1 illustrates the fact that U.S. tariffs, both preferential and multilateral, are falling over time. Is there any (causal) link between the two series?

Figure 1 about here

According to received wisdom, PTAs are a way for countries that are frustrated with the painfully slow pace of the Doha round (launched in December 2001) or pundits who expect the Doha round to collapse altogether to overcome these problems and

[1] PTAs are also known as regional trade agreements (RTAs) in WTO terminology. They include customs union (about 10% of such PTAs) and free-trade areas (FTAs) and partial scope agreements (90%). Source for figures: 'WTO regional trade agreements' website, updated July 2008: http://www.wto.org/english/tratop_e/region_e/region_e.htm.

liberalise their bilateral trade faster than in the multilateral forum (The Economist 2008). It is also conceivable that any multilateral trade deal leaves bargaining currency on the table; some subsets of countries have then some incentives to cut bilateral tariffs beyond the multilateral level. Consequently, this line of thinking would interpret the wave of regionalism that followed the conclusion of the Uruguay Round in December 1994 as a consequence of the limitations of the Uruguay Round. A logical corollary is that preferential margins should be systematically larger for tariff lines on which the multilateral cut has been shallowest.

If this view is right, then regionalism is a *substitute*, albeit imperfect, to multilateralism. Being part of a PTA represents for a WTO member a departure from the guiding principle of non-discrimination defined in Article I of GATT, Article II of GATS, and elsewhere. The consequence for the avowed aim of the GATT/WTO would then be quite grim: the world renounces to global free trade. Regionalism is one defining symptom of this scourge.

In this paper, we question this view. Specifically, we use highly disaggregated data for the preference margins granted by the United States to their PTA partners as of 1996, as well as data for the multilateral tariff cut agreed during the Uruguay Round (1994). Our data set comprises 8,661 tariff lines and ten PTAs that run from 1996 to 2007, totalling 28,454 observations for our reference sample.[2] We chose this time period so that we can be confident that Uruguay Round negotiators were negotiating multilateral cuts without having these PTA ‘concessions’ at the back of their mind at the time. We then run a regression at the good-partner level of the form:

$$\text{(preference margin)} = \beta\text{(multilateral cut)} + \text{controls} + \varepsilon \qquad (1)$$

(‘controls’ include a battery of fixed effects that minimise the omitted variable bias). If the estimated β is negative, then preference margins are highest for tariff lines (what we refer to as ‘goods’) that had the shallowest multilateral cuts and this would support the ‘substitution’ hypothesis. However, our data strongly reject this hypothesis: for the U.S. trade policy at least, the current wave of regionalism is, if

[2] An observation is a combination tariff line $\times$ partner for which a preference margin has been negotiated. Most tariff lines are included in at least one PTA. See section 3 for details.

anything, a *complement* to past cuts granted at the multilateral level (all our estimates of β are positive). Few goods are included in all PTAs and most goods are included in only a few PTAs. We exploit this variation to run a regression at the good level of the form:

$$(\#\text{preference margins}) = \gamma(\text{multilateral cut}) + \text{controls} + \xi \qquad (2)$$

Here we find that the number of times a good is included in a PTA (i.e. the frequency at which the U.S. grants a preference margin on this good to its PTA partners) is increasing in the MFN tariff cut. Therefore, both pieces of evidence suggest that past multilateral cuts acted as *dynamic complements* for current PTA preference margins. We refer to (1) as the *intensive margin* of *the emulator effect* of multilateralism on regionalism and we interpret (2) as its *extensive margin.* We expect it to vary across industries, PTAs and other dimensions. We run a series of extensions to account for such heterogeneity in various ways; our regressions reject the *emulator effect* in none of them.[3]

In the first of this series of extensions, we allow β to vary across the ten PTAs of our sample; we find that the 'emulator effect' is strongest for the generalised system of preferences (GSP) for least developed countries (LDCs) entered into force in 1997 and weakest for the bilateral trade agreement with Australia that entered into force in 2006. Second, we allow the intercept and the β for agricultural goods to be different from the constant and the coefficient for non-agricultural ones: we find that the average preferential margin granted for the former set of goods is lower than for the latter, but we find that the emulator effect is much stronger for the agricultural goods than for the non-agricultural ones; the estimated β is significant for both types of goods at the one percent confidence level. Third, we expect the determinants of preference margins for intermediates goods and final goods might be quite different; we thus break the sample into seven categories, loosely regrouped in three 'meta-categories' (final goods, intermediate goods and the rest). In all seven of them, the estimated β's are all positive in a statistical sense. A clear pattern also emerges: the emulator effect is strongest for intermediate goods and weakest for final goods. One

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[3] To keep the paper within reasonable length, we conduct our robustness check exercises for the intensive margin only.

possible explanation for this is that the emulator effect is non-linear. To check for this, we run a quantile regression: interestingly, we find that the emulator effect is lowest for the smallest preference margins and it is highest for the largest ones. It is statistically negative for none of the quantiles.[4] As a further robustness check, we also run our regressions in logs. The overall fit is lower, but the emulator effect also finds strong support in this final specification.[5]

Finally, our dataset reveals that only a subset of tariff lines are affected by a PTA and that very few tariff lines are affected by all PTAs in our sample. The selection of goods into a given PTA is unlikely to be exogenous. We thus re-run our baseline regression using a two-step Heckman procedure. Our regressions confirm the existence of a bias but the coefficient of interest remains positive (its magnitude is reduced by only 20%).

To summarise our findings, in virtually any specification that we estimate in this paper, we find that tariff cuts negotiated at the multilateral level triggered a positive, subsequent response at the preferential level.

From our empirical exercise, we thus conclude that the current wave of regionalism reflects the past success of the GATT/WTO more than the limitations of the Uruguay round or the current stalemate of the Doha round. This does not call for complacency in the current context of WTO director Pascal Lamy's attempt to save the Doha round, but instead suggests some caution before interpreting the current wave of regionalism as the consequence of a failure of the multilateral trading system. We view our paper as a first step towards understanding the causes of regionalism.

The section in the immediate sequel introduces the theoretical backdrop for our work. Section 3 introduces our dataset and some summary statistics. Section 4 presents our estimation strategy. Section 5 contains our benchmark regression results, followed by a series of extensions and robustness checks in Section 6. Section 7 discusses our

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[4] We also run distinct quintile regressions for each of the seven categories. Two patterns emerge, which are in line with our previous findings. First, the estimated β is positive for all of the 7x5 = 35 quintiles, but two (the first two quintiles of 'consumption goods'). Second, the estimated β is rising monotonically with the quintile for all of the seven categories.

[5] We also run the log specification using quantile regressions: we find that the relationship does not have a constant elasticity, so we stick with our regression in levels in the core of the paper.

findings and broadens the perspective to include the recent evidence on the *stumbling bloc versus building bloc* controversy.

2. THEORETICAL BACKDROP

Are multilateral and regional trade agreements complements or substitutes? Does multilateralism stimulate or impede regionalism?[6] This question receives a large coverage in the media but it is largely neglected in the academic literature. Ethier (1998) are Freund's (2000a) theoretical papers are rare exceptions.[7] Ethier (1998) combines a trade theory with a theory of vertical foreign direct investments to show that when the market for final goods expands as a result of multilateral trade liberalisation among rich countries, then the demand for intermediate goods rises and this may encourage the formation of PTAs in which a rich country sources the production of intermediate inputs to the developing PTA partner. By contrast, Freund (2000), develops an oligopolistic model of intra-industry trade à-la Brander and Krugman (1983), which is well-suited to study rich-rich PTAs by construction (see Krishna 1998). In her model, the preferential and the external tariff are complements: an exogenous reduction of either makes a reduction of the other more desirable.[8] She then extends her static framework to a repeated game and goes on to show that an exogenous reduction of the multilateral tariff makes a PTA more likely to be sustainable (in game theoretical terms). Thus, albeit for different reasons, all three models vindicate the idea that a multilateral move towards freer trade increases the incentive to form preferential agreements. Our data (described in Section 3) comprises

[6] A large literature has addressed the symmetric question: is regionalism a building or a stumbling block towards multilateral free trade? We cover some of the findings of this body of research in Section 7.

[7] Maggi and Rodriguez-Clare (1998) and Ornelas (2005b) study the impact of regionalism on the incentives to adopt multilateral free/freer trade; interestingly, the decision to adopt an FTA is endogenous in their models.

[8] To understand this result, consider her three-country symmetric model. Two countries may form a PTA and impose a bilateral tariff τ on each other's exports; the multilateral tariff governing any other bilateral trade relation is t, with $t > \tau$. Governments maximise social welfare (consumer surplus + producer surplus + tariff revenue). If, say, t falls then imports from the third country increase in each PTA member country, which reduces profits, consumer surplus and tariff revenue. A marginal symmetric reduction of τ may reduce profits further but it unambiguously increases consumer surplus and tariff revenues. As a result, the optimal τ is increasing in t. The same qualitative result obtains in a customs union; see Freund (2000, p364). A companion paper to this one provides evidence for this effect for the U.S. (Fugazza and Robert-Nicoud 2008).

PTAs formed between the U.S. at one end and both developing and rich countries at the other end. Our empirical evidence is consistent with this positive equilibrium relationship.

These models are static in essence: Ethier's (1998) model is akin to a one-shot sequential game and Freund (2000) focuses on stationary equilibria of her dynamic game. However, one predominant pattern of worldwide tariff data is that virtually any bilateral or multilateral tariff goes down over time (recall Figure 1). In addition, multilateral tariff are negotiated under the strict and – in order to satisfy the unanimity rule – time consuming auspices of the GATT/WTO, whereas countries can start negotiating preferential tariffs more easily and more frequently. Thus, we assess that a reasonable assumption is that multilateral and preferential tariffs are chosen sequentially. In particular, multilateral free trade is typically *not* an available option while countries negotiate a PTA.[9] So far, existing theories that explicitly seek to understand why tariffs are falling over time do not address the interaction between preferential and multilateral tariffs (Staiger 1995, Maggi and Rodriguez-Clare 2007, Baldwin and Robert-Nicoud 2007) and some static models that do address such interactions do make the assumption that multilateral free trade is on the table.

To overcome these limitations, Robert-Nicoud (2008) extends the Krishna-Freund-Ornelas oligopolistic framework to a dynamic setting, where the distinction between the extensive margin of trade (i.e. the number of trading firms) and the intensive margin of trade (i.e. the volume of trade by firm) is explicitly introduced.[10] Such extensions allow one to study the dynamics of trade liberalisation and the two-way interaction between PTA and multilateral tariffs sparked by the political-economy model. To see this, assume that the extensive margin of trade adjusts more slowly than the intensive margin, an assumption that can easily be rationalised.[11] Since one margin of adjustment is not available in the short run, the oligopoly's attitude to any

[9] Of course, unilateral free trade *is* an option. Actually, many countries (mostly developing) apply tariffs rates ('applied rates') that are below the tariff ceiling negotiated under the GATT/WTO ('bound rates').

[10] See Maggi and Rodriguez-Clare (1998) for a similar approach in a competitive model.

[11] Specifically, it takes time and real resources to establish trading relations and enter new markets; see Baldwin (1989). Rauch (1999) asserts that networks are important for establishing trading relations in differentiated products and provides evidence consistent with this view.

kind of trade liberalisation is more hostile than in the long run. The model predicts that a lower external (or MFN) tariff makes it more likely that the PTA members bilaterally decrease their internal tariff. This is consistent with the relationship that our paper uncovers.

3. DATA & SUMMARY STATISTICS

There are three types of negotiated tariffs: the bound MFN tariffs (BMFN), the applied MFN tariff (AMFN) and the preferential tariff. As already mentioned, the tariff of interest to WTO negotiators is the BMFN tariff. Tariffs cuts observed during GATT/WTO trade negotiations refer to cuts in the BMFN tariffs. The AMFN tariff could be no higher than the BMFN and its level is (unilaterally) chosen by the imposing country. In the case of the United States (and others), the BMFN and AMFN tariffs coincide exactly and are usually decreasing in the post-Uruguay round period, so we refer to them as the MFN tariff for short. Our key explanatory variable, denoted by ΔMFN , is defined as the (non-negative) difference between the Uruguay and Tokyo MFN rates.

Each preferential tariff, if it exists at all, is PTA-specific. We define the *preferential margin* as the (non-negative) difference between the MFN tariff and the preferential tariff prevailing at the end of the implementation period (when available) and we denote it by ΔPT . Specifically, we observe in our sample that many preferential tariffs are falling over time before stabilising at a constant level. When this is the case, we take this constant level to construct our 'preference margin' variable. Our sample ends in 2007, however, and for some PTAs the preferential tariffs are still falling at this time. We shall conduct robustness checks that exclude such PTAs from our reference sample.

For our empirical purposes, an *observation* in our *reference sample* is a combination tariff line $\times$ partner such that (i) the preference margin is positive, (ii) the MFN rate in the final year of observation is positive and (iii) the PTA was negotiated after 1994 (the conclusion of the Uruguay round). Our *reference sample* comprises 3,455 such tariff lines for a total of 28,454 observations. Condition (i) implies that we study the determinants of the preference margin *conditional* on a preference margin being granted to a given partner. If the MFN tariff rate is zero, the preferential rate cannot be lower than the MFN rate (trivially). This binding constraint on the preference margin, if not taken into account, would induce a bias in our regressions; however, there are no observations that simultaneously satisfy (i) and violate (ii) in our reference sample, so we may not worry further about this source of bias. We relax conditions (i) and (ii) in robustness checks below. Finally, we impose condition (iii) so as to eliminate an obvious source of reverse causality bias from our regressions.

The main sources for data are the UNCTAD-TRAINS and the WTO-CTS Bound Duty Rates databases. Both databases provide information at the legal tariff line level (8-digit the HS nomenclature) – what we refer to as *goods* below. There are 8,661 such tariff lines in the unrestricted database. The WTO-CTS database provides information on bound rates negotiated at the Tokyo round and those negotiated at the Uruguay round. Hence, the difference between these two rates corresponds to the effective reduction in bound tariffs negotiated during the Uruguay round. The database also provides information on the implementation period of bound tariff reductions that were negotiated during the Uruguay round. The implicit rule revealed by analysing the data indicates that the absolute yearly cut in the MFN tariffs corresponds to the overall negotiated cut divided by the number of years of implementation. For instance, if MFN rates are falling from 20 to 10 percent over a period of 10 years then, the implicit implementation framework would stipulate a reduction of 1 point percentage each year. Lines subject to non-ad valorem duties were not included in the original dataset; such tariff lines account for around 8% of HS 6-digit subheading in 2006 according to the World Tariff Profiles (2007).

The UNCTAD-TRAINS database includes MFN applied rates and preferential rates. The period informed is 1996-2007. Preferential Trade Agreements covered in the database are: the Generalized System of Preferences (1976), Israel (1985), Caribbean Basin Economic Recovery Act (1986), Andean Trade Preference Act (1992), NAFTA (1994), GSP for Least Developed Countries (1997), the African Growth and Opportunity Act (2000), Caribbean Basin Trade Partnership Act (2000), Jordan (2001), Chile (2004), Singapore (2004), Morocco (2006), Bahrain (2006), Australia (2005), Dominican Republic-Central American FTA (2006). Preferences margins are computed over the whole period of investigation, namely any period between 1996 and 2007. From the exhaustive list above, we exclude the PTAs that were negotiated before the end of the Uruguay round (1994). For the remainder of the paper, we refer to the set of countries belonging to one of these PTAs as a ‘partner’ or ‘partnership’ and we index such partnerships with the subscript $p \in \{1,\ldots,10\}$; note that ‘*p*’ is also a mnemonic for ‘PTA’. Table 1 reports the ten partnership included in our reference sample.

Table 1 about here

Concerning non-tariff measures (NTMs), we use information available in the TRAINS database. We focus on NTMs classified as Technical Measures (chapter 8) in the UNCTAD Coding System of Trade Control Measures. This covers *inter alia* both sanitary and phyto-sanitary (SPS) and technical barriers to trade (TBT) type of measures. Data are available only for the year 1999.

Table 2 provides summary statistics for the main variables used in estimations (detailed definitions for each variable are provided in Appendix Table A1). The mean MFN tariff agreed during the Tokyo round of multilateral trade negotiations (ending in 1979) is 9.81 percents for our sample mean; the mean MFN tariff for the Uruguay round is 6.63. The first cell of the table reports that the mean MFN cut is 3.18, which is the difference between the two previous numbers. The mean preference margin, which is the mean of the difference between the preferential rate in the final year of observation minus the Uruguay MFN rate, is equal to 6.17. In other words, preferential cuts are on average bolder than multilateral cuts. The second column reports the standard deviations and, together with the third and fourth columns, suggests that there is substantial variation to exploit. We should also notice that MFN cuts on tariff lines included in PTAs are strictly positive. This means that zero MFN cuts are perfectly associated with zero preference margins.

We also use sectoral output data for our two-step Heckman robustness check. The data are imported from the UNIDO Industrial Statistics database (INDSTAT4-Rev3). The use of sectoral output based on UNIDO data unfortunately limits the sectors covered to manufacturing according to the ISIC classification at the four-digit level.

Table 2 about here

Figure 2 illustrates the key information uncovered in Table 3, which breaks down the information from Table 2 in distinguishing between agricultural goods and non-agricultural ones.[12] The information summarised in this table reveals that agricultural goods are quite different from non-agricultural ones in two respects at least. First, by any of the measures reported in this table, agricultural goods are more protected than non-agricultural ones at any point in time; this is true both at the preferential and at the multilateral levels. Second, MFN tariff cuts and preferential margins are boldest for agricultural products by any measure. In a sense, there is some 'trade liberalisation' catching-up for these goods. We shall exploit this source of variation in some of our regressions below.

Table 3 about here

Figure 2 about here

Figure 3 reports the number of tariff lines with a preference margin broken down by PTA. We observe that the number of tariff lines subject to a PTA varies from a low 2,325 for the generalised system of preferences (GSP) for least developed countries (LDCs) to a maximum of 3,219 for Chile (CHL). The Duty Rate database reports that 3,455 (39.9%) tariff lines ('goods') out of 8,661 are subject to a PTA. Figure 4 reports the frequency analysis of tariff lines included in the various PTAs signed by the U.S. For instance, less then 1% of the tariff lines subject to a PTA (18 out of 3,455) appear in only one post-Uruguay round PTA and 40% of them (1,407 out of 3,455) appear in all of them. The last column and bottom lines of Table 4 (in appendix) encompass the information illustrated in Figure 3 and Figure 4, respectively. We shall control for these variations across partners and tariff lines in our empirical analysis.

Table 4 about here

Figure 3 about here

Figure 4 about here

[12] The distinction between agricultural and non-agricultural goods follows the classification used in trade negotiation at the WTO. See for instance the World Tariff Profiles (2006) for a detailed description based on the Harmonized System Nomenclature 2002.

4. ESTIMATION STRATEGY

As mentioned previously, estimates are for the U.S. trade policy; in other words, the U.S. is the importing country in all that follows.

Let p denote partners (individual countries or members of a specific PTA), g denote each good and G denote sectors (HS-2 in the benchmark regressions and HS-4 in robustness-check specifications).[13] We denote the set of all goods by Γ; then G is a partition of Γ and (with some slight abuse of notation) we use $G(g)$ to denote the sector in which good g is classified. Thus, G is also a mapping $G : \text{good} \rightarrow \text{sector}$.

Emulator effect: the intensive margin. In our baseline specification we run the following regression:

$$\Delta PT_{g,p} = f_p + f_{G(g)} + \beta \Delta MFN_g + \varepsilon_{g,p} \tag{3}$$

where $\Delta PT_{g,p} = MFN_{g,t} - PT_{g,p}$ is the preferential margin applied by the U.S. on imports of good g from country $p \in \{1,...,10\}$ in the final year t ($t \geq 1997$ in our sample); it is defined as the (non-negative) difference between the (applied) MFN tariff prevailing when the PTA enters into force (i.e. in some year $\tau \leq t$) and the preferential tariff being applied in the final year t that we observe in our sample. Notice that the MFN rate it is always larger or equal to the bound rate negotiated during the Uruguay round. In a similar way, the corresponding MFN tariff-cut, $\Delta MFN_g \equiv MFN_g^{\text{Tokyo}} - MFN_g^{\text{Uruguay}}$, is defined as the Tokyo-Round MFN tariff (as agreed in 1979) minus the Uruguay-Round MFN tariff (as agreed in 1994). We use the negotiated MFN tariff rates. Finally, $\varepsilon_{g,p}$ is an error term assumed to be independently and identically distributed from some distribution with finite first and second moments.

We must assume that ΔMFN_g is exogenous in order to obtain consistent and unbiased estimates from (3). The heart of our identification strategy rests on the timing of

[13] There are 97 HS-2 sectors and 636 HS-4 sectors in our benchmark sample. We use HS-2 dummies in our benchmark regressions to save degrees of freedom and to save on computational power. In this, we also follow Limão (2006). We later run our most basic regressions with HS-4 dummies as robustness checks.

events. We limit our sample to the ten PTAs, listed in Table 1, which entered into force after the conclusion of the Uruguay round in 1994. This sample selection is expeted to eliminate any reverse-causality bias for two main reasons: first between 1994 and 2000 no new trade agreements had been implemented by the U.S., with the unique exception of an extension of the GSP scheme granted to LDCs. Such an extension is pretty unlikely to have influenced negotiations during the Uruguay round. Second, no trade agreement signed in the post Uruguay round period had actually been negotiated in the pre Uruguay round period. The Clinton administration did undertake talks to form a Free Trade Area of the Americas (FTAA) and to sign a trade agreement with the Asian Pacific Economic Cooperation (APEC) country members in 1994. However, it is unlikely that those talks had any influence on the tariff reductions negotiated at the Uruguay round. On one hand, no agreement has been reached yet in the context of FTAA negotiations. On the other hand the APEC forum held in Bogor in 1994 signed a declaration to work toward free trade in the region by 2010 for developed countries and by 2020 for all member-countries. A sixteen-year time frame makes any influence of those talks on tariff cuts defined the Uruguay round quite implausible. What is usually recognized is that the APEC summit together with NAFTA helped "squeeze the European Union to complete the Uruguay round of GATT".[14]

We introduce sector fixed effects $f_{G(g)}$ in (3) to capture two dimensions of unobserved heterogeneity. Indeed, if an omitted variable affects ΔPT and ΔMFN simultaneously, then regressing the former on the later will cause a spurious correlation. Some political economy determinants (lobbying) of tariffs could constitute this omitted variable, as suggested in our theoretical discussion in the introduction. The determinants of comparative advantage are another source of omitted variable bias. Since we have a cross section of observations, we cannot include goods fixed effects in (3). Insofar such unobserved shocks are common to goods within sectors, then including $f_{G(g)}$ in (3) corrects for this source of omitted variable bias. It is also reasonable to assume that lobbying takes place at the sector

[14] Robert Zoellick (2001) statement then-USTR.

level. It is also reasonable to assume that the level of aggregation at which the forces of comparative advantage are operating is best approximated by the sector.

We also include partner dummies f_p in (3) to capture bargaining power of the partner and all other partner-specific considerations that arise during the negotiation of the PTA; such considerations can be trade- and non-trade related (Limão 2007). We expect a stronger bargaining position to translate into a larger average margin of preference. Partner dummies (together with sector dummies) can also reflect reciprocity of the agreement, that is, the preference margin given to US products. Indeed, the reciprocity 'rule' may vary across agreements; there is for instance no explicit reciprocity in the GSP scheme for LDCs (Romalis 2007). More reciprocity (larger preference margins in partner markets) should lead to larger preference margin. In addition, exchange of concessions may be implemented at different levels, from the most highly disaggregated level (i.e. goods) to the HS-2 sectoral level, or a combination of the two. This flexibility in the framing of exchange of concessions was not encountered in the GATT/Uruguay round negotiations as described in Finger and al. (1996). We do not disentangle reciprocity from bargaining power because it is unclear whether a low bargaining position also means more reciprocity. We assume that our partner and sector fixed effects together absorb these various effects. We later consider various extensions of our benchmark specification to account for various bargaining-power-reciprocity schemes and associations.

Pursuing the logic of removing any possible element of reverse-causality further, we only consider tariff lines with implementation starting in 1995 and no later. That is to say that we drop all negotiations post Uruguay round or that could be seen as a continuation of the round. This is the case of tariff reductions implemented under the agreement on trade in information technology products (ITA) concluded at the Singapore Ministerial Conference in December 1996. These tariff lines represent less than 1 % of observations in the reference sample and less than 2% in the whole sample; thus this attrition is minimal. To attenuate concerns related to the selection bias of our sample, we later run regressions that include all data available as a robustness check of our results. In order to eliminate issues possibly related to data truncation, we restrict our reference sample to strictly positive MFN cuts observations.[15] Overall, we reduce the size of our sample by around 10%.

To summarize, we run (3) for the observations that are subject to a preferential margin. That is, (3) being a conditional regression, it estimates the *intensive margin* of the emulator effect. However, in order to obtain an unbiased and efficient estimate of β, we must assume that the observations for which the preference margin is strictly positive are randomly generated. This is a strong assumption. Therefore, in an extension we run a two-step regression, where the first step consists in estimating:

$$\Pr\{\Delta PT_{g,p} > 0\} = F_p + F_{G(g)} + \alpha_0 MFN_g + \mathbf{X_g \alpha_1} + \mu_{g,p} \tag{4}$$

where $\mathbf{X_g}$ is a set of good-characteristics (like production and employment) and MFN_g is the relevant tariff *level*. The economic logic for the latter determinant of $\Pr\{\Delta PT_{g,p} > 0\}$ is the following: when the multilateral tariff is low to start with, there is not much preference margin to be granted. Thus, the likelihood that a preference margin is being granted should be increasing in the manoeuvre margin, i.e. the MFN tariff. Taken literally, no existing theory has anything to say about the probability that a good is included in a PTA but probabilistic interpretations of Ornelas (2005b) and Freund's (2000) models are consistent with this view. We then run a modified version of (3),

$$\Delta PT_{g,p} \big| \Delta PT_{g,p} > 0 = f_p + f_{G(g)} + \beta \Delta MFN_g + b\hat{m}_g + \varepsilon_{g,p} \tag{5}$$

where $\hat{m}_g$ is the inverse Mills ratio estimated from the first stage. Although this estimation strategy is attractive from an economic point of view, the statistical properties of US tariff data push us to consider it as a robustness check only. We mentioned previously that all tariff lines with a zero Uruguay MFN bound rate are never included in any PTA. As they represent more than 70% of the tariff lines not included in any PTA, correcting for the selection bias may result simply in accounting for the presence of MFN duty free lines which by definition could be excluded from any trade arrangement as they do not represent any potential for preferential access.

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[15] This is always the case for preferences margins by definition.

Emulator effect: the extensive margin. As underlined in section 3 no positive preference margin is observed on tariff lines where the MFN cut is zero. This suggests that zero MFN cuts are likely to predict almost perfectly the non inclusion of a tariff line in a given PTA. Hence, as a separate piece of evidence that the selection of a tariff line is not random we regress the *number* of times good g is being selected in any PTA on sector dummies and ΔMFN_g, namely

$$\#\{p : \Delta PT_{g,p} \geq 0\} = f_{G(g)} + \gamma \Delta MFN_g + \varepsilon_{g,p} \tag{6}$$

Estimating (6) by OLS would yield biased results, though, because the left hand side variable is a count variable. To address this issue, we use Poisson estimation techniques. The generic Poisson specification assumes that the relationship between the number of PTAs per line and its determinants is based on a constant semi-elasticity model which implies estimating,

$$\#\{p : \Delta PT_{g,p} \geq 0\} = \exp\left(f_{G(g)} + \gamma \Delta MFN_g \right) + \varepsilon_{g,p} \tag{7}$$

For the sake of comparison with (6), we also assume that the underlying relationship between the number of PTAs per line and its determinants remains strictly linear. That is, we estimate

$$\exp\left(\#\{p : \Delta PT_{g,p} \geq 0\}\right) = \exp\left(f_{G(g)} + \gamma \Delta MFN_g \right) + \varepsilon_{g,p} \tag{8}$$

The results of these regressions will inform us about the *extensive margin* of the emulator effect of multilateral trade agreements on preferential margins.

5. RESULTS

Baseline regressions. The results of our baseline regression are presented in the first column of Table 5. Specifically, we obtain (using '^' for estimated values):

$$\hat{\Delta} PT_{g,p} = \hat{f}_p + \hat{f}_{G(p)} + \underset{(11.53)}{0.216} \Delta MFN_g \tag{9}$$

Our estimate for β, $\hat{\beta} = .216$, is precisely estimated, for the robust t-statistics (in parenthesis) is above 11, thus the estimated β is positive well beyond the one percent level of confidence.[16] Putting this in perspective, consider the effect of reducing the MFN tariff rate by one standard error. Then, a 3.39% MFN tariff cut between the Tokyo and Uruguay Rounds is associated with an extra (.216)(3.39%) = .73% tariff cut at the preferential level, which represents 15% of the standard error of the preferential margin at the sample average. Our regression explains fully 31.1% of the variation in our sample (adjusted-R^2), which is quite substantial.[17] Therefore, we conclude that the emulator effect of multilateral trade liberalisation on regionalism finds strong support in the data, that is, *the past success of multilateralism is at least partly responsible for the current wave of U.S. regionalism*; but also that sector characteristics and partner attributes are responsible for a large part of it. We then run the regression excluding Morocco, Bahrain and DR/CAFTA from the sample on the grounds that these were not fully implemented by 2006; we obtain $\hat{\beta} = .274$ (not reported in Table 5).

Table 5 about here

The second column of Table 5 reports estimations including the goods for which the MFN rate is zero in the final year of observation. Since preferential rates cannot be larger than MFN rates, the corresponding preferential rates must be equal to zero, too. This introduces a downward bias in our estimates. This is confirmed by the fact that $\hat{\beta} = .204$ (t-stat. 11.85) in this case, which is statistically significantly lower than $\hat{\beta} = .216$ obtained in column (1). Despite this downward bias, the qualitative result whereby more MFN liberalisation is associated with more Preferential liberalisation is strongly supported.

Figure 5 about here

Partner-specific β. Partner bargaining power could not only affect the overall average margin of preferences but also the influence of the MFN cuts. Also, Foreign

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[16] As MFN rates by definition are the same for all country partners, intra country group correlation is likely to be present. With cluster-robust standard errors estimation the coefficent remains significant beyond the one-percent level although we observe an important drop of the t-statistic to 5.36.

[17] The adjusted R^2 we report is the 'true' one: if one excludes the constant and include ten fixed effects (one for each partner *p*) to the regression, then a software like STATA® (the one we use) reports an artificially inflated R^2, in our case: $R^2 = .74$. The true one, R^2=.31, is substantially lower.

lobbies might influence U.S. policy in differently (Gawande, Krishna and Robbins 2006). Thus, in the third column of Table 5, we allow the coefficient β to vary across partners; specifically, we run

$$\Delta PT_{g,p} = f_p + f_{G(g)} + \beta_p \Delta MFN_g + \varepsilon_{g,p} \tag{10}$$

Figure 5 illustrates the point estimates as well as the 95- and 99 percent confidence intervals; the dashed line reports the point estimate $\hat{\beta} = .216$ of the pooled regression (9) in order to ease comparison. The effect is strongest for GSP for LDCs ($\hat{\beta}_{\text{GSP}} = .708$). The effect is weakest for Singapore (the point estimate, though positive, is not different from zero in a statistical sense) and for Bahrain, Morocco and DR/CAFTA. We briefly study the determinants of the variation across partners in section 6.

Table 6 about here

Agriculture-specific β. Agricultural goods are distinct from others in many ways. Industrialised countries impose systematically higher tariffs on such goods. The U.S. is no exception, as can be seen from Table 3. Therefore, we allow the relationship between the MFN cuts and the preferential margins to be specific to the type of good. Table 6 reports the results. In column (1), we simply regress equation (3), adding an agricultural dummy, 'AGRI'. The point estimate for β, $\hat{\beta} = .214$, is not different from $\hat{\beta} = .216$ in (9) in a statistical sense. The AGRI dummy is negative, indicating that, on average, tariffs cuts have been lower for agricultural goods once unobserved sector and partner characteristics are controlled for in the regression. This contradicts the message conveyed by the (unconditional) first moments of the distribution of cuts of Table 3. We next allow both the intercept and the coefficient to differ for agricultural goods. Specifically, we run

$$\begin{aligned}\Delta PT_{g,p} &= f_p + f_{G(g)} + f_{AGRI} I_{AGRI(g)} \\ &\quad + \beta_{AGRI} I_{AGRI(g)} \Delta MFN_g + \beta_{NON-AGRI} \left(1 - I_{AGRI(g)}\right) \Delta MFN_g + \varepsilon_{g,p}\end{aligned} \tag{11}$$

where $I_{AGRI(g)} = 1$ if good g is classified as an agricultural product and $I_{AGRI(g)} = 0$ otherwise. The results, reported in column (2) of Table 6, indicate that the impact of MFN cuts is stronger for agricultural goods ($\hat{\beta}_{AGRI} = .652$ against $\hat{\beta}_{NON-AGRI} = .131$

for non-agricultural products) although the observed preferences margins are smaller ceteris paribus (the agricultural dummy has a negative coefficient which significantly undermines the impact of MFN cuts). Thus, including this distinction leaves the qualitative effect unaffected but it also shows that the emulator effect is sector-specific. That is to say, the political and economic forces that shape trade policy vary across sectors.

Figure 6 about here

Sector-specific β. The previous regression shows that the quantitative emulator effect varies across the type of goods. Therefore, we now allow the coefficients to vary across our 97 HS-2 sector and run

$$\Delta PT_{g,p} = f_p + f_{G(g)} + \beta_{G(g)} I_{G(g)} \Delta MFN_g + \varepsilon_{g,p} \tag{12}$$

where $I_{G(g)} = 1$ if good g belongs to sector G and $I_{G(g)} = 0$ otherwise. We do not report the 97 estimated coefficients here, for obvious reasons, but Figure 6 illustrates the results and reveals that the emulator effect operates in most HS-2 sectors; the emulator effect is rejected in only 13% of the sectors.[18] In Figure 6, we have ordered the 97 sector-specific estimates in increasing order of $\hat{\beta}_G$. We also report the 95 and 99 percent confidence intervals. Some $\hat{\beta}_G$'s are not precisely estimated, for some sectors have relatively few tariff lines affected by a PTA and/or little variation.

Extensive margin. A separate piece of evidence in favour of the emulator effect comes from estimating (6) . Specifically, we estimate the effect of the MFN cut on the number of times a tariff line appears in any of the ten PTAs of our reference sample. In other words, we assess the *extensive margin* of the emulator effect. Our Poisson estimations as specified in (7) and (8) give a positive coefficient significant at the 1% level, as shown respectively in columns 1 and 2 of table 7.[19] Therefore, all

[18] Out of the 97 HS-2 sectors, 49 sectors (51%) have an estimated β that is larger than zero at the one percent level (at the five percent level , this number rises to 51 sectors), 13 sectors have an estimated β that is negative at the one percent confidence level and a third of sectors (33 in total) with an estimated β that is not different from zero. Of these 33 sectors, 12 were dropped by STATA® (their β could not be estimated for lack of sufficient variation).

[19] As a check we also estimated (6) using standard robust OLS. We obtained a positive coefficient equal to 0.79 significant at the one percent level.

specifications indicate that the larger the MFN cut on a given good, the more often it will be included in a PTA. We interpret this result as evidence of *an extensive margin of the emulator effect*: tariff lines with bolder MFN cuts negotiated during the Uruguay round are more frequently included in PTAs.

Table 7 about here

6. EXTENSIONS AND ROBUSTNESS CHECKS

In this section we consider several extensions of regressions (3) and (10). As we shall see, the qualitative result whereby multilateral liberalisations emulated preferential tariff reductions is unaffected. The quantitative emulator effect also remains of the same order of magnitude.

Table 8 about here

Bargaining power. We first seek to refine our econometric specification to account for bargaining power and reciprocity. Results are reported in column (1) of Table 8. We assert that the partner's GDP (scaled by the U.S. one), is a reasonable proxy.[20] However, we have no prior as to the sign of the coefficient on GDP_p/GDP_{US} because this proxy captures effects that potentially work in opposite directions. The benchmark $\hat{\beta}$ in (9) and the explanatory power are both unchanged in this specification. The sign of the coefficient of the relative GDP variable is negative ($\hat{\beta}_{GDP} = -.503$) and precisely estimated. In plain English, preferences margins are smaller for bigger countries. This suggests that the likely interaction between bargaining power and reciprocity mechanisms dominates the effect of the bargaining power alone. Alternatively, it may also be the case that relative GDP may also proxy the competition threat of the trade partner for domestic industries. A larger partner GDP (smaller GDP difference) could reflect similar industrial structures and thus

[20] In the results we report in column (1), we use the GDP_p/GDP_{US} ratio prevailing the year before the signing of the trade agreement. In unreported regressions, we run alternative regressions using several Gross Domestic Product measures: Purchasing Power Parity (PPP) $US, in per capita units expressed in current and PPP $US. In any of these regressions, the coefficient on MFN cuts is statistically indistinguishable from .216 and the explanatory power of the specification remains identical to that of the benchmark. The sign of the coefficient of the relative GDP variable is negative throughout and the coefficient is always precisely estimated.

more competition from foreign firms for domestic ones. As a consequence, it would not be surprising to find lower overall average preferences margins for trade partners, which are closer on GDP measures ground. We also run a regression similar to column (1) with the addition of an interaction term between relative GDP and the MFN cut. Results (not reported in the table) show that this leaves the coefficients of column (1) virtually unaffected ($\hat{\beta} = .221$ and $\hat{\beta}_{GDP} = -.491$ in this case).

NTBs. We next control for the presence of NTBs (i.e. technical barriers in our database). The corresponding additional dummy does not affect the coefficient on MFN cuts as shown in column (2) of Table 8. The adjusted R^2 remains the same as in the benchmark regression. We obtain a positive and significant coefficient for the NTBs dummy ($\hat{d}_{NTB} = .181$). This indicates that the preference margin is larger in the presence of an NTB. In column (3), we also the NTB dummy with the coefficient on MFN cuts (not reported). The coefficients on MFN cuts are equal to $\hat{\beta}_{N-NTB} = .299$ and $\hat{\beta}_{NTB} = .097$ for tariff lines unaffected by technical barriers and tariff lines affected by a technical barrier, respectively. The coefficient of the NTB dummy remains positive ($\hat{d}_{NTB} = .083$). Thus, the average preference margin is larger on tariff lines affected by a NTB but the emulator effect is weaker. This suggests that such tariff lines got relatively large tariff cuts at the multilateral level, which is consistent with the view that technical barriers often replace tariff barriers. The adjusted R^2 remains equal to 0.31.

Figure 7 about here

Categories of goods. Column (4) of Table 8 reports the results of regressing (3) using HS-4 dummies instead of HS-2 sector dummies. There are 636 such sectors. The resulting estimated coefficient decreases ($\hat{\beta} = .175$) but remains positive. Column (5) replaces the sector dummies by dummies reflecting the type of goods. This classification counts 7 categories: Basic manufacturing, Consumption goods, Equipment goods, Intermediate goods, Mixed products, Primary goods, and a residual category (which includes only 38 observations). The coefficient on MFN cuts is equal to 0.289, which is larger than in our baseline regression. The adjusted R^2 falls to 15%. All category dummies are positive (the benchmark category is Basic manufacturing) except that corresponding to Equipment goods. This means that all categories but

Equipment Goods have a preferential margin larger than the one observed for the Basic Manufacturing category. We further interact the good category variable with the MFN cuts one to qualify the result above. We do not report the seven coefficients in the table but we report them in Figure 7 (we report both the point estimates of the β's and the 95 and 99 percent confidence intervals). An interesting pattern emerges: on the whole, *final goods* ('Consumption goods' and 'Basic manufacturing') get a lower $\hat{\beta}$ than *intermediate inputs* ('Intermediate goods', 'Equipment goods' and 'Primary goods').[21] Thus, the emulator effect is strongest for intermediate inputs. This is consistent with the view that downstream sectors lobby actively against protecting upstream sectors (Gawande and Bandyopadhyay 2000), extended to a dynamic sense.

Figure 8 about here

Quantile regressions. Our regressions by type of goods, by sector and by partner, suggest that β is not the same throughout the sample. If tariff lines subject to a PTA are not randomly assigned but vary systematically with the preference margin, then this calls for a quantile regression. Thus we run (3) not conditioning on the mean of the distribution but on deciles (Koenker and Basset 1978). Figure 8 illustrates the results (the decile-specific $\hat{\beta}$'s are reported on the vertical axis). A clear pattern emerges: the emulator effect is tiny (and marginally significant) for the lowest decile and rises monotonically to above .5 for the highest one ($\hat{\beta}_{.9} = .523$). In plain English, the emulator effect is strongest for the largest preference margins. This suggests that the emulator effect dampens the variation by MFN cut. It is all the more noticeable, then, to see this effect emerging in the pooled regression reported in (9). We also run distinct quintile regressions for each of the seven categories. Two patterns emerge, which are in line with our previous findings. First, the estimated β is positive for all of the 7x5 = 35 quintiles but two (the first two quintiles of 'consumption goods'). Second, the estimated β is rising monotonically with the quintile for all of the seven categories.

[21] Goods that belong to the 'Mixed' and 'NES' categories can not easily be categorised as either final goods or intermediate inputs.

Final MFN rate. In our estimations so far, the tariff rate retained for computing the preference margin in PTAs is the AMFN rate prevailing at the time of entry into force of the PTA. As a robustness check, we use instead the MFN rate that shall prevail at the end of its implementation period, as negotiated during the Uruguay round. After all, trade negotiators expect this rate to prevail eventually as they negotiate PTA rates. As should be expected, the coefficient of MFN cuts decreases but, at $\hat{\beta} = .123$, it remains positive and significantly different from zero. Details of this regression are reported in column (6) of Table 8.

Full sample. Results obtained for the whole original sample using the benchmark specification (3) are reported in the next column of Table 8. Our preferred specification uses only a subset of available data in order to avoid an upward bias: in some instances, the MFN rate is zero in the final year of the implementation of the PTA. When this happens, the PTA tariff that year is constrained from above and the preference margin is constrained from below. In column (7), $\hat{\beta} = .349$ is significantly larger than the one obtained in (9), confirming the existence of an upward bias.

Restricted sample. Finally, we exclude two PTAs from our sample, the GSP for LDCs (1997) and AGOA (2000) because such agreements have a large number of duty free tariff lines so as to ensure that the emulator effect is not being identified by a special case. The result of the regression on this restricted sample is reported in column (8) of Table 8. Our main effect remains positive and it is only slightly weakened ($\hat{\beta} = .181$).

Table 9 about here

Selection bias. Not all the tariff lines (forty percent) are affected by a PTA (Table 4, Figure 3 and Figure 4). In addition, two thirds of the tariff lines that are affected by some PTA are included in only a fraction of the ten PTAs of our sample. What are the determinants of including a tariff line in a PTA? To answer this question, we run a Probit on (4). Results are reported in Table 9. All the specifications we run (the result of some of them are not being reported here) suggest the existence of a systematic selection of tariff lines into PTAs: the $\hat{\alpha}$'s are all significant different from zero (the coefficient of the Inverse Mill's Ratio -IMR- variable is always negative and highly significant). We first run (4) using the full sample, excluding the UNIDO controls $\mathbf{X_g}$.

Results are reported in column (1). We find that the probability that a tariff line is selected into a PTA is positively associated with the MFN tariff level: that is to say, countries that negotiate a PTA are more likely to include goods that have a high MFN tariff to start with. These are the goods for which they have the largest margin of manoeuvre.[22] We also find that some observations have been dropped for some sector dummies perfectly predict the selection into not being part of any PTA (this happens for tariff lines with little variation and/or few observations). Two solutions are available when running the second step (5): either we drop such observations or we attribute them a predicted value for the first stage of $\hat{\Pr}\{\Delta PT_{g,p} > 0\} = .01$, i.e. almost zero[23]. We report the results of the latter only.[24] We run (5) on the full sample in column (1a), whereas we run (5) on our reference sample in column (1b). In both cases, $\hat{\beta} < .216$ in a statistical sense[25]; this reveals the existence of a systematic downward bias. Nevertheless, the emulator effect of multilateral trade liberalisation on preferential margins remains positive, even after controlling for the fact that negotiating partners chose to cut tariffs where there is the largest margin to grant 'concessions'. Columns (2a) and (2b) replicate the two-step procedure of columns (1a) and (1b), respectively, adding the UNIDO sector controls. Since the UNIDO variables (output, employment, wages, value added) are collinear, we use output only. Running (4), the estimated coefficient is negative, indicating that the U.S. is more likely to grant preferential access in the sectors where the stakes (in terms of output, employment, wages or value added) are lowest.[26] The estimated coefficients for (5) are $\hat{\beta} = .202$ and $\hat{\beta} = .189$ for the 'UNIDO-full' and 'UNIDO-reference' samples, respectively. The downward bias is much smaller, but one should recall that using

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[22] As far as the US are concerned, the average Uruguay MFN rate is equal to 2.06 for those lines not included in any PTA. The corresponding figure jumps to 6.63 for those lines included in at least one PTA. Standard deviations are respectively 6.37 and 6.86.

[23] A comparable approach is adopted for instance in Helpman, Melitz and Rubinstein (2008).

[24] When we chose the former the estimated coefficient for β falls to .13.

[25] Results reported in Table 9 are obtained with standard OLS robust estimation. We have also computed bootstrapped standard errors as the inverse Mill's ratio is an estimated variable. Standard errors based on 500 replications hardly vary from the ones reported in table 9 and do not affect any coefficient significance test, at any level.

[26] This is in line with the findings of e.g. Trefler (1993).

UNIDO controls requires one to scale down the samples – which may introduce a bias on its own.[27]

Alternative specifications. We conclude with a series of simple robustness checks on the benchmark regression that we do not report in tables. Specifically, we first estimate the following within-estimator:

$$\Delta_p^{\text{within}}\Delta PT_{g,p} = \tilde{f}_{G(g)} + \beta\Delta_p^{\text{within}}\Delta MFN_g + \varepsilon_{g,p}\,, \tag{13}$$

where

$$\Delta_p^{\text{within}}\Delta PT_{g,p} \equiv \Delta PT_{g,p} - \overline{\Delta PT}_p\,, \qquad \Delta_p^{\text{within}}\Delta MFN_g \equiv \Delta MFN_g - \overline{\Delta MNF}$$

and

$$\overline{\Delta PT}_p \equiv \sum\nolimits_{g=1}^{|g|_p}\Delta PT_{g,p}\,, \qquad \overline{\Delta MFN} \equiv \sum\nolimits_{g=1}^{|g|_p}\Delta MFN_g\ .$$

In words, $\Delta_p^{\text{within}}\Delta PT_{g,p} \equiv \Delta PT_{g,p} - \overline{\Delta PT}_p$ and $\Delta_p^{\text{within}}\Delta MFN_g \equiv \Delta MFN_g - \overline{\Delta MNF}$ are the difference between tariff cuts at the good-level and the sample average (preferential and multilateral, respectively). Also, $\overline{\Delta PT}_p$ is defined as the average preferential margins granted to partner *p* and $|g|_p$ is the number of goods included in the preferential trading agreement *p*; $\overline{\Delta MNF}$ is defined analogously. Clearly, if the ε's are *i.i.d.*, then (13) is equivalent to (3); indeed, running (13), we obtain $\hat{\beta} = .216$ (t-stat. 11.52), which are identical to the values obtained in (9). In a similar spirit, we also estimate the 'diff-in-diff':

$$\Delta_p^{\text{within}}\Delta PT_{g,p} - \Delta_p^{\text{within}}\Delta PT_{\gamma,p} = \beta\left(\Delta_p^{\text{within}}\Delta MFN_g - \Delta_p^{\text{within}}\Delta MFN_\gamma\right) + \varepsilon_{g,p} \tag{14}$$

for any *g* and γ such that $g \neq \gamma$ and $G(g) = G(\gamma)$. In this case, we find that $\hat{\beta} = .216$ (t-stat. 11.48).[28]

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[27] As a further robustness check (unreported), we run (3) adding UNIDO controls (i.e. output). The reported estimated coefficient rises to .246 (t-stat = 13.15). The coefficient on output is negative (t-stat = .3.66). The adjusted R^2 is equal to .31. We then run both (4) and (5) with the UNIDO control; the estimated β is equal to .202 (UNIDO-full sample) and .189 (UNIDO-restricted sample). The t-stats are 13.31 and 11.98, respectively.

Table 10 about here

A final robustness check consists of using annual rates cuts rather absolute ones. This is motivated by the fact underlined previously that the time length and path of implementation are pre-determined. Annual rate cuts are obtained by dividing the simple difference between two rates by the number of years it takes to obtain a flat tariff profile (the tariff rate becomes constant). A similar path can be attributed to PTAs implementation. Regression results are reported in Table 10. In the benchmark specification (column 1), the coefficient of the annual rate of the MFN cut is equal to 0.46 with a standard deviation of 0.025: larger annual cuts in MFN bound rates imply larger annual preference margins. Column (2) and (3) includes the AGRI dummy and its interaction with the speed of MFN cuts variable. The speed of preferential margins implementation is slower for agricultural goods and the impact of the speed of MFN cuts is comparable to any other good.

7. CONCLUSIONS

This paper investigates the empirical relationship between cuts in MFN bound rates negotiated during the Uruguay Round of the GATT (1986-1994) and preferences margins given to PTAs partners in the aftermath of its completion. Our empirical investigation focuses on the United States using official tariff line level data. Our results reveal that an ‘emulator effect’ of multilateral trade negotiations on preferential agreements is at work: multilateralism and regionalism appear as ‘dynamic complements’. Indeed, preferences margins are larger the larger the cut in MFN bound rates. Moreover, a tariff line is included in a larger number of PTAs the larger the cut in MFN bound rates. Therefore, we can state that the past success of multilateralism is at least partly responsible for the current wave of U.S. regionalism.

Our empirical results also indicate that although preference margins are on average smaller for agricultural products, the emulator effect for those products is strong relative to non agricultural products. When we control for the length of the

[28] Regressing deviations from the mean implies regressing values that have been computed. Therefore, we have also bootstraped the errors to obtain unbiased estimations of the standard errors. The estimated coefficient remains unchanged and it t-statistic varies around 11.50 depending to the number of replications.

implementation period, we find that the emulator effect is *stronger* for non-agricultural products. Together, these findings suggest that the emulator effect is *slower* for agricultural products. We conduct several other extensions. In one of them, we find that the emulator effect is stronger for intermediate inputs than for final goods.

To the best of our knowledge, our paper is unique in looking at the causal relationship from multilateralism to regionalism. The existing empirical literature, only burgeoning, is exclusively looking at the relationship running the other way (Limão 2006, Limão and Karacaovali 2007, Estevadeordal et al. *forthcoming*). This line of research is motivated by the view expressed in numerous theoretical contributions that regionalism has a 'stumbling block' effect on multilateral trade liberalisation (Bhagwati 1991). If the stumbling block hypothesis is correct, then the proliferation of PTAs involving at least one WTO member is guilty of slowing down and threatening the (by now interrupted) 'Doha round' of negotiations at the GATT/WTO. Trade negotiators should then urge to reach a final agreement in order to contain the stumbling impact of forthcoming PTAs (e.g. the FTA between the US and Malaysia whose negotiations are in active mode). This point of view is consistent with the argument that NAFTA and discussions with APEC countries gave the US leverage on the EU within negotiations at Uruguay round.

A related and pessimistic received wisdom, which runs in the other direction, is that the explosion of regionalism is a symptom of the difficulties encountered by the Doha round. However, the fact that multilateral trade negotiations may be seen a strong instigator of the current wave of regionalism could also lead us to think that delayed negotiations are not necessarily reflecting any stumbling block effect of PTAs. Rather it could reflect political/negotiations slowness due to a multilateral agenda which has become more ambitious than it would have been without growing regionalism. Moreover, trade negotiations appear to take place in a context where political leadership is lacking (Irwin, Mavroidis and Sykes 2008)..

This paper questions this received wisdom and shows that (recent) trade liberalisation negotiated between the U.S. and various small sets of countries was boldest for those goods for which the tariffs had been reduced most sharply during the Uruguay round of multilateral trade liberalisation that ended in 1994. Such dynamic complementarities between sequential rounds of trade liberalisation are consistent with the 'Juggernaut' theory of trade liberalisation put forth by Baldwin (1994) and formalised by Staiger (1995), Maggi and Rodriguez-Clare (2007), Baldwin and Robert-Nicoud (2007) and Robert-Nicoud (2008). This theory stresses the role of *domestic* sluggish adjustments to account for the systematic, monotonically decreasing trade barriers of the modern trading system.

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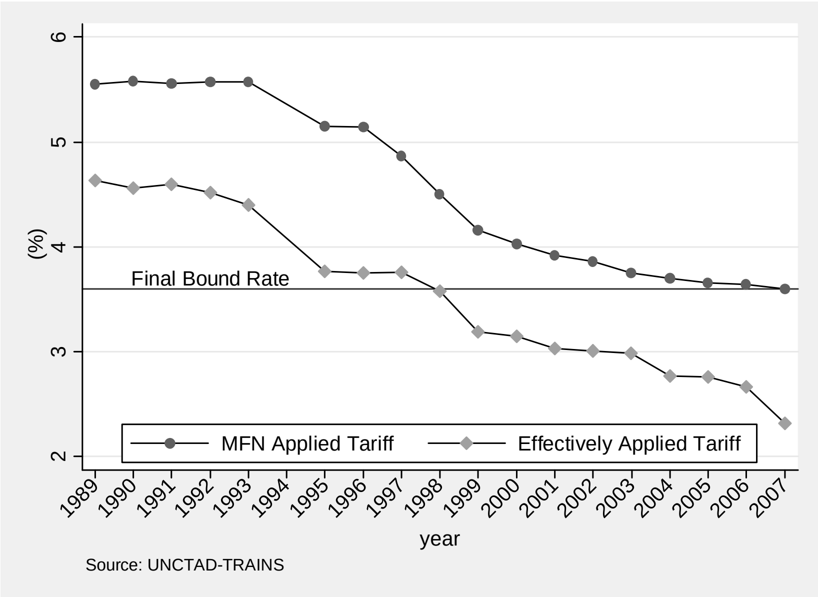
FIGURES

Figure 1. **Tariffs are falling over time (US data - simple averages)**



Note: The effectively applied tariff is defined as the lowest available tariff. If a preferential tariff exists, it is used as the effectively applied tariff. Otherwise, the MFN applied tariff is used.

Figure 2**. Average ad-valorem tariffs: Agri vs. non-agri.**

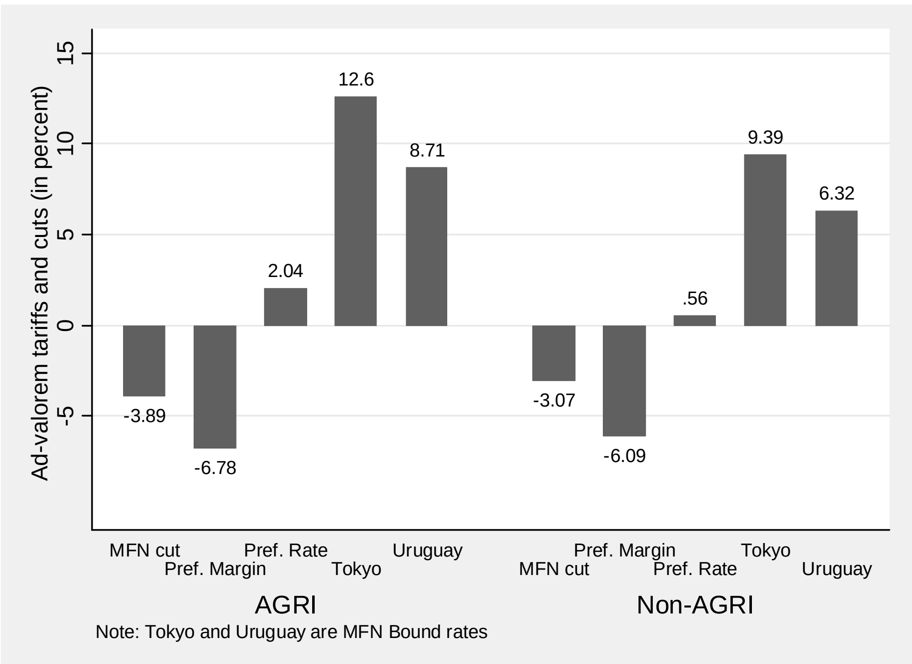


Figure 3**. Number of tariff lines included in PTAs, by PTA**

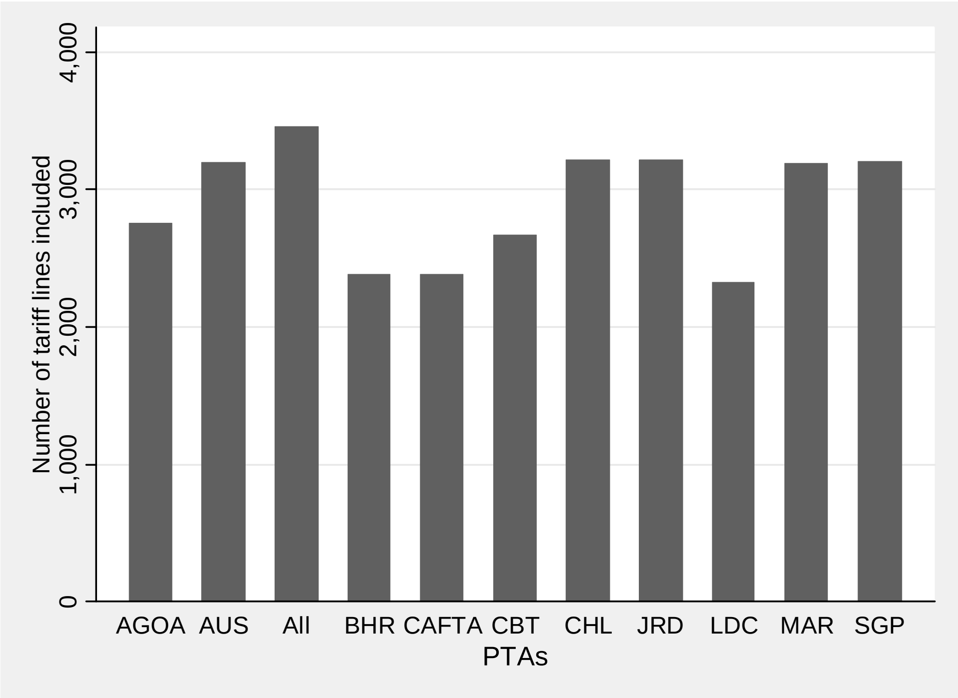


Figure 4. **Number of tariff lines included in PTAs, by frequency**

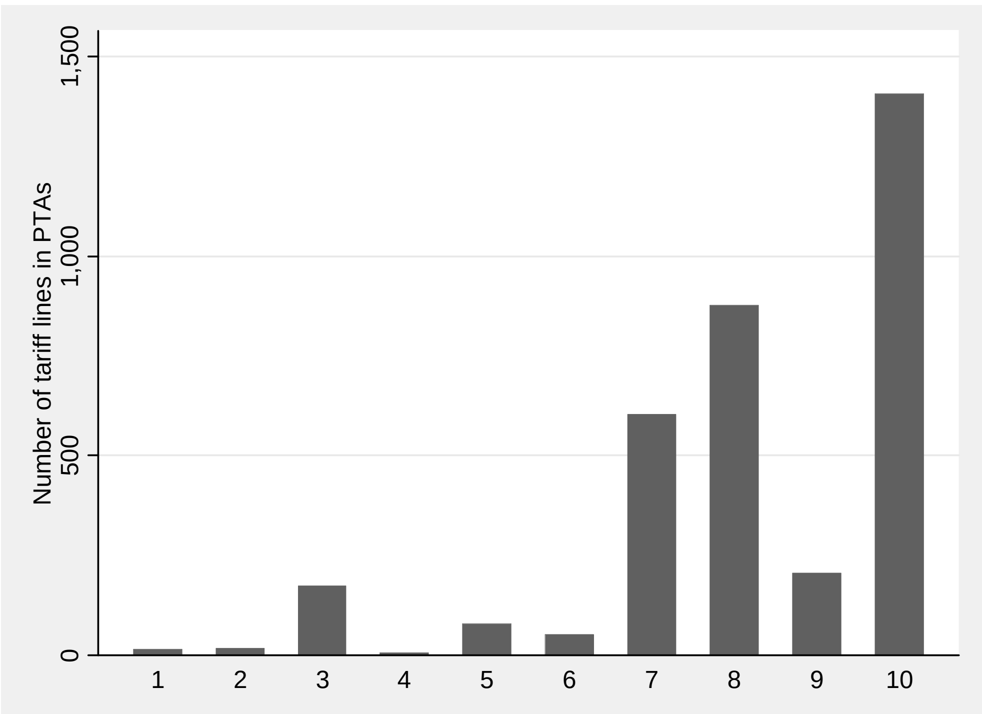


Figure 5**. Partner-specific coefficients**

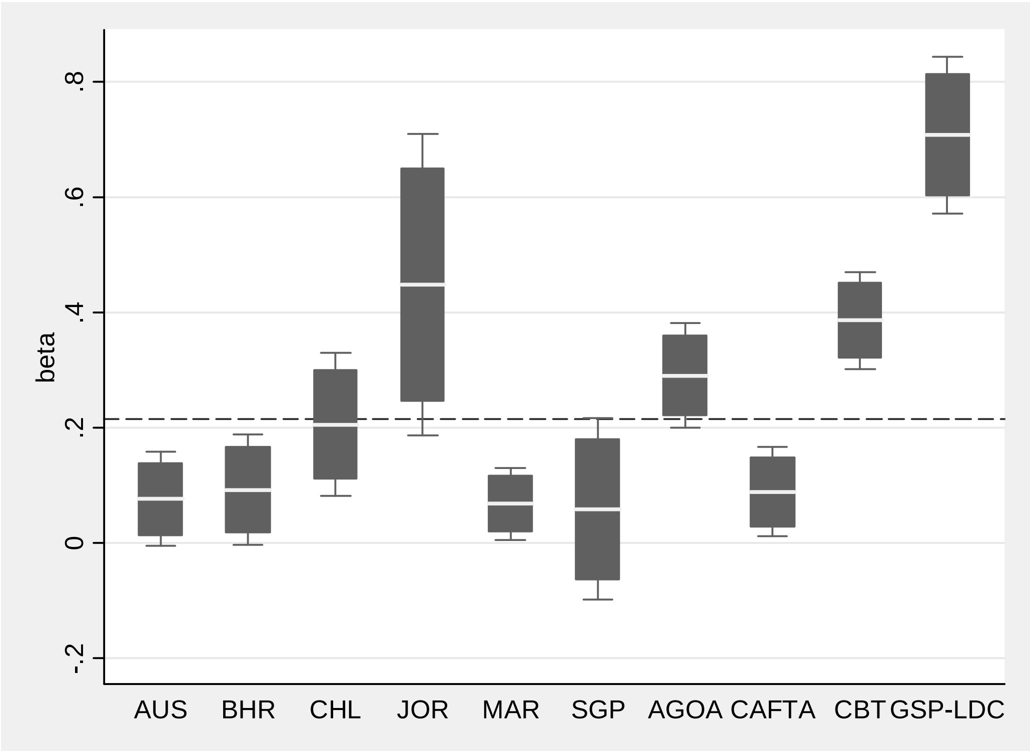


Figure 6**. Sector-specific coefficients**

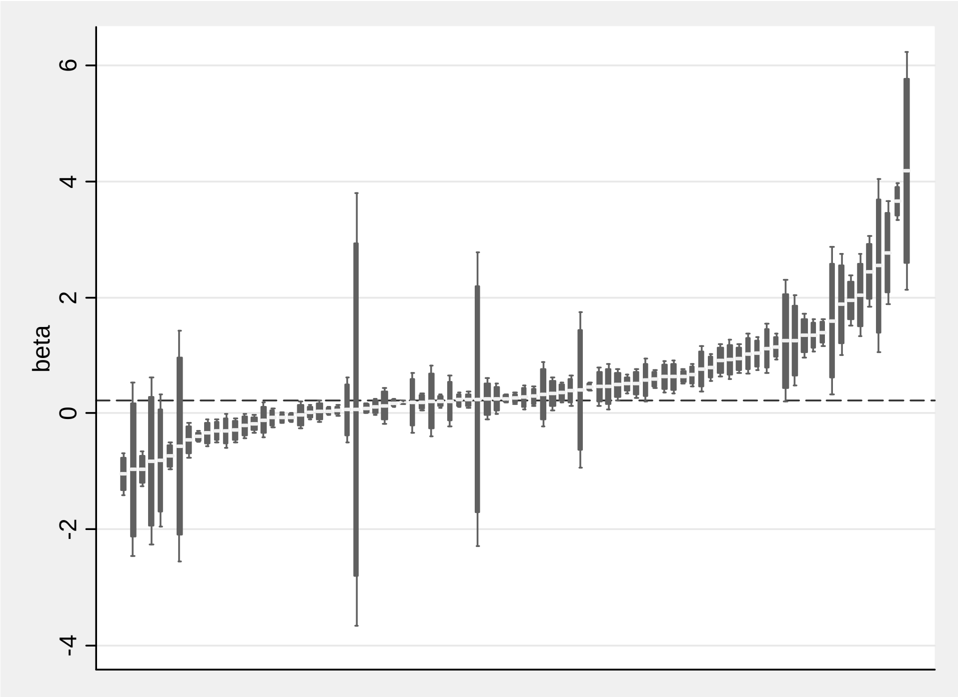


Figure 7**. Type-specific coefficients**

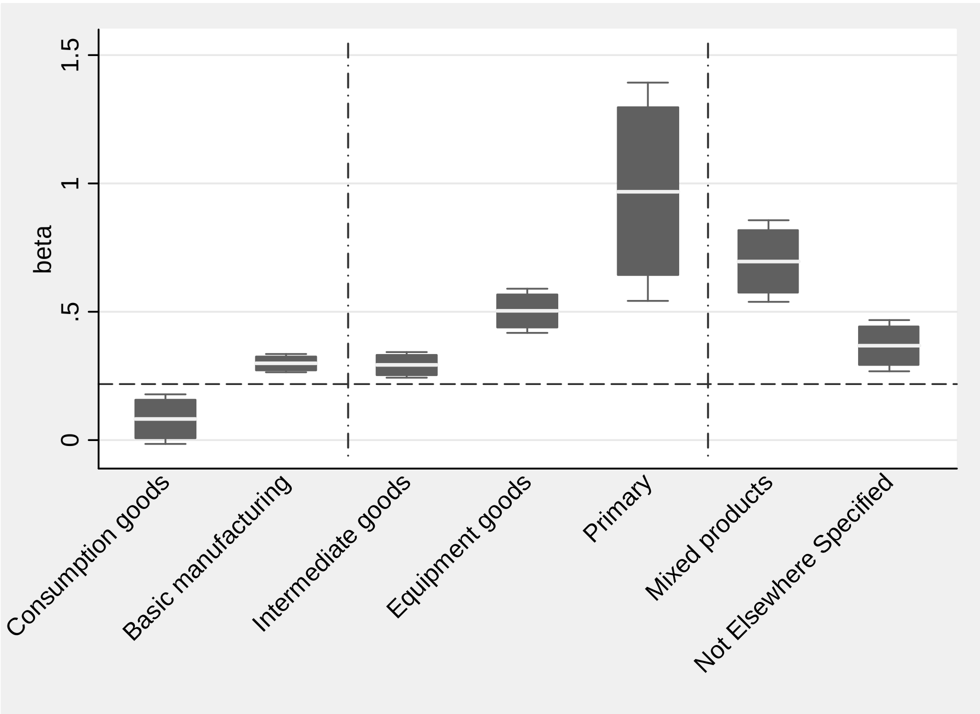
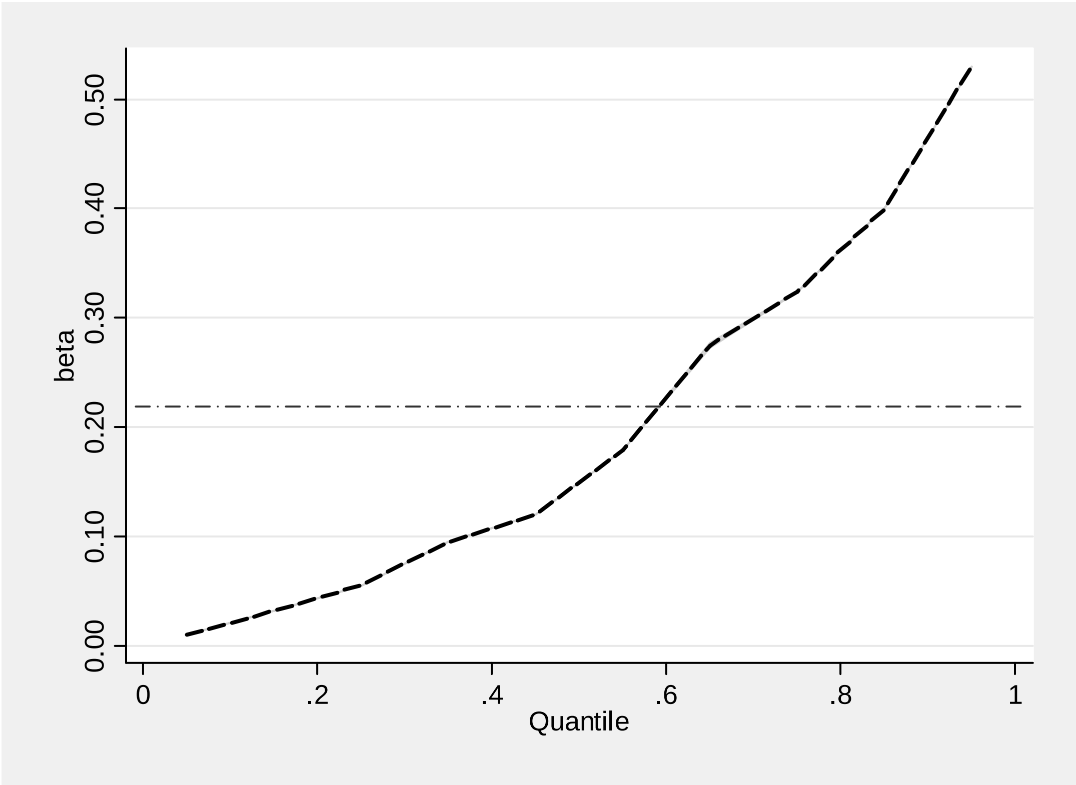


Figure 8**. Quantile regression: coefficients**



TABLES

Table 1**. Sample of PTAs**

p	Partner	Year of Implementation
1	Australia	2005
2	Bahrain	2006
3	Chile	2004
4	Jordan	2001
5	Morocco	2006
6	Singapore	2004
7	African Growth Opportunity Act (AGOA)	2000 (AGOAI)-2002 (AGOAII)-2004 (AGOAIII)
8	Domican Republic-Central America (DR-CAFTA)	2005
9	Caribbean Basin Trade Partnership act (CBT)	2000
10	Generalized System of Preferences for Least Developed Countries (GSP-LDCs)	1997

Table 2. **Summary Statistics (Reference Sample)**

	Mean	**Std. Dev.**	**Min.**	**Max.**
MFN-cut	**3.18**	3.39	0.1	31.5
Preference Margin	**6.17**	4.87	0.1	114.7
No. of Observations: 28,454				

Table 3. **Summary Statistics (Reference Sample): Agri. vs. Non-agri.**

	Agricultural goods				Non-agricultural goods			
	Mean [a]	**Median [a]**	**Mean [b]**	**Median [b]**	**Mean [a]**	**Median [a]**	**Mean [b]**	**Median [b]**
BMFN Uruguay	**9.52** (16.)	**6.4**	**8.71** (13.65)	**6.4**	**6.47** (4.94)	**5.5**	**6.32** (4.89)	**5.5**
BMFN Tokyo	**13.48** (18.71)	**10**	**12.6** (16.07)	**10**	**9.55** (6.27)	**7.5**	**9.39** (6.2)	**7.2**
MFN-cut	**3.96** (3.93)	**3**	**3.89** (3.78)	**3**	**3.07** (3.37)	**1.9**	**3.07** (3.32)	**2**
Preference Rate			**2.04** (9.85)	**0**			**.56** (2.08)	**0**
Preference Margin			**6.78** (6.53)	**5.4**			**6.09** (4.59)	**4.9**

[a] Number of agricultural goods: 384. Number of non-agricultural goods: 3,058. Each good (tariff line) gets an equal weight.

[b] Number of observations (agricultural goods): 3,428. Number of observations (non-agricultural goods): 25,026. Each tariff line × partner gets an equal weight.

Note. The figures for the Preference Rate and Margin in column ([a]) are not well defined since there can be up to 10 different Preference Rates and Margins in our sample of ten partnerships.

Table 4**. Frequency of tariff lines with a preferential tariff rate**

	Frequency of partners										
	1	2	3	4	5	6	7	8	9	10	Total
AUS	0	0	0	5	73	39	584	878	208	1,407	3,194
BHR	0	0	0	0	0	4	437	330	208	1,407	2,386
CHL	0	0	2	6	82	50	588	878	206	1,407	3,219
JRD	4	0	2	8	82	53	576	878	208	1,407	3,218
MAR	0	0	0	1	73	40	583	877	208	1,407	3,189
SGP	0	0	2	6	73	50	581	878	208	1,407	3,205
AGOA	4	8	173	3	9	20	148	781	203	1,407	2,756
CAFTA	0	0	0	0	0	1	437	330	208	1,407	2,383
CBT	7	18	173	4	9	47	153	646	208	1,407	2,672
LDC	3	14	173	3	9	20	141	548	7	1,407	2,325
All	18	20	175	9	82	54	604	878	208	1,407	**3,455**
(%)	0.5	0.6	5.1	0.3	2.4	1.6	17.5	25.4	6.0	40.7	100

Notes. The dataset contains 3,455 unique tariff lines. Each cell informs about the frequency of tariff line × partner pairs. For instance, the first column reveals that there are 4 tariff lines that are subject to the U.S.-JORDAN (JRD) PTA and no other. Likewise, there are respectively 3 tariff lines that are unique to the U.S. GSP for LDCs (LDC). The fourth column reveals that there are five tariff lines that are subject to the U.S.-Australia PTA as well three other PTAs (hence the column's name 4).

Table 5. **Benchmark regressions**

diffPREF	(1)	(2)	(3)
diffMFN	0.216	0.203	
	(11.53)**	(11.75)**	
AUSTRALIA*diffMFN			0.076
			(2.38)*
BARHAIN*diffMFN			0.093
			(2.46)*
CHILE*diffMFN			0.206
			(4.26)**
JORDAN*diffMFN			0.449
			(4.38)**
MOROCCO*diffMFN			0.068
			(2.77)**
SINGAPORE*diffMFN			0.059
			-0.95
AGOA*diffMFN			0.291
			(8.21)**
CAFTA*diffMFN			0.089
			(2.93)**
CBT/DR*diffMFN			0.387
			(11.69)**
GSP-LDCs*diffMFN			0.708
			(13.23)**
Industry dummies	Yes	Yes	Yes
Partner dummies	Yes	Yes	Yes
Observations	28454	31239	28454
Parameters	96	96	105
R-squared Adjusted	0.31	0.33	0.33

Robust t statistics in parentheses;* significant at 5%; ** significant at 1%

Table 6. **Regressions by type of products (agri. vs. non-agri)**

diffPREF	(1)	(2)
diffMFN	0.214	
	(11.38)**	
AGRI	-1.547	-2.997
	(10.28)**	(-10.26)**
AGRI*diffMFN		0.652
		(6.59)**
NON-AGRI*diffMFN		0.131
		(12.38)**
Industry dummies	Yes	Yes
Observations	28454	28454
Parameters	97	98
R-squared Adjusted	0.31	0.33

Robust t statistics in parentheses ;* significant at 5%; ** significant at 1%

Table 7. **Regressions on the extensive margin (number of PTAs per tariff line)**

	(1)	(2)
diffMFN	0.11	0.12
	(17.71)**	(15.14)**
Industry dummies	Yes	Yes
Observations	8,661	8,661
Parameters	98	98
Pseudo R-squared	0.25	0.28

Robust t statistics in parentheses ;* significant at 5%; ** significant at 1%

Table 8. **Extensions and robustness checks (I)**

diffPREF	GDP (1)	NTB (2)	NTB (3)	HS-4 sector dummies (4)	Stage of production (5)	Uruguay rates (6)	Whole sample (7)	Restricted sample (8)
diffMFN	0.216	0.216		0.175	0.289	0.123	0.349	0.181
	(11.5)**	(11.52)**		(3.74)**	(15.95)**	(6.51)**	(19.09)**	(8.70)**
GDP_p/GDP_{US}	-.503							
	(-21.63)**							
Dummy NTB (d_{NTB})		0.181	0.833					
		(2.24)*	(6.31)**					
d_{NTB} * diffMFN			0.097					
			(2.87)**					
(1-d_{NTB}) * diffMFN			0.299					
			(14.59)**					
Industry	Yes	Yes	Yes	Yes	No	Yes	Yes	Yes
Partner	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	28454	28454	28454	28454	28454	28454	49303	23413
Parameters	97	97	97	653	17	96	102	94
R-squared Adjusted	0.31	0.31	0.32	0.44	0.15	0.28	0.33	0.30

Table 9. **Selection**

	First Stage	
	(1)	(2)
$MFN_{Uruguay}$	0.587	0.617
	(154.29)**	(153.37)**
UNIDO		- 6.54e-12
		(19.76)**
Observations	52,510	44,050
Pseudo-R^2	0.72	0.67

	Second Stage			
	(1a)	**(1b)**	**(2a)**	**(2b)**
diffMFN	0.194	0.173	0.202	0.189
	(10.71)**	(9.63)**	(13.31)**	(11.98)**
IMR	-7.887	-8.131	-7.675	-7.558
	(56.99)**	(56.91)**	(95.63)**	(88.11)**
Observations	34,118	28,454	24,845	20,824
R-squared Adjusted	0.39	0.38	0.47	0.45

Robust t statistics in parentheses ;* significant at 5%; ** significant at 1%

Table 10. **Extensions and robustness checks (II)**

SdiffPREF	(1)	(2)	(3)
SdiffMFN	0.46	0.458	
	(10.09)**	(10.05)**	
AGRI		-0.324	-0.251
		(-5.38)**	(-3.27)**
AGRI*diffMFN			0.359
			(6.19)**
NAGRI*diffMFN			0.491
			(8.64)**
Industry dummies	Yes	Yes	Yes
Observations	20408	20408	20408
Parameters	96	97	98
R-squared Adjusted	0.45	0.45	0.45

Robust t statistics in parentheses;* significant at 5%; ** significant at 1%