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THE 'EMULATOR EFFECT' OF THE URUGUAY ROUND ON UNITED STATES REGIONALISM

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Abstract

Using a detailed data set at the tariff line level, we find an emulator effect of multilateralism on subsequent regional trade agreements (RTAs) involving the United States. We exploit the variation in the frequency with which the United States grants immediate duty free access (IDA) to its RTA partners across tariff lines. A key finding is that the United States grants IDA status especially on goods for which it has cut the multilateral most favoured nation (MFN) tariff during the Uruguay Round the most. Thus, the Uruguay Round (multilateral) "concessions" have emulated subsequent (preferential) trade liberalization. We conclude from this that past liberalization may sow the seeds of future liberalization.

Key words: Regionalism, multilateralism, stumbling bloc, Uruguay Round

JEL Classification: F13, F14, F15, N70

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Any mistakes or errors remain the authors' own.

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1. Introduction

Many preferential trade agreements came to light since the completion in 1994 of the Uruguay Round of multilateral trade negotiations under the auspices of the General Agreement on Tariffs and Trade (GATT). The United States is no exception. These agreements involving the United States vary in scope – the number of goods included in the agreement varies across agreements – and breadth – the United States tariff on some goods goes to zero immediately upon implementing the agreement but the imports of many others are fully liberalized only gradually. In this paper, we shed light on the causes of these cross-good variations and show that they are best thought of as the continuation of a process that includes multilateral liberalizations. Specifically, we find that the imports of goods that the United States liberalizes swiftly the most frequently on a preferential basis are also the goods for which it granted the boldest tariff *cuts* during the Uruguay Round and/or those for which the current MFN tariff *levels* are low. Both findings are robust to a variety of specifications. The quantitative effects are also quite large. We interpret the former finding as evidence that past trade agreements are dynamic complements, or emulator, to consecutive agreements. The latter finding is consistent with the idea that the benefits of new regional trade agreements are especially large when multilateral tariffs are low.

By design, the paper addresses on an important and little studied question in international trade: whether multilateral trade agreements (MTAs) drive, in any way, the proliferation of regional trade agreements (RTAs). Since there is some concern and evidence that RTAs block or slow down the formation of MTAs (Limão, 2006), it is important to understand whether the success of one MTA may actually lead to the failure of the following MTA negotiations as a result of increased regionalism. Several theoretical papers have demonstrated that a multilateral reduction in tariffs can increase the formation and the self-enforceability of RTAs (Ethier, 1998; Freund, 2000a). Other papers have shown in a dynamic framework that existing trade agreements erode the resistance to liberalization and hence pave the way for further liberalization in the future (Staiger, 1995; Maggi and Rodriguez-Clare, 2007). These papers are motivated by the fact that in many countries tariffs are declining over time (figure 1 illustrates this pattern for the United States).[1] We take this feature of the data seriously in our analysis.

While existing empirical papers on the subject focus on the determinants of RTA *formation* (e.g. Baier and Bergstrand, 2004; Egger and Larch, 2008; Mansfeld and Reinhardt, 2003), the current paper strives to explain the impact of an MTA on the *characteristics* of RTAs that follow it. Our contribution is threefold. We start by examining which products are liberalized most swiftly in an RTA, taking its existence as given. In particular, as we explain in section 3, we focus on RTAs signed by the United States after the Uruguay Round in 1994 and, in section 4, we show that the products that were most likely to be liberalized quickly in an RTA were precisely the ones that had the largest tariff *reduction* in the Uruguay Round and, for a given Uruguay Round tariff reduction, those that had high Uruguay round tariff *levels*. The former finding supports the claim that MTAs and RTAs are dynamic complements and it constitutes our first contribution. The latter finding provides original evidence that is consistent with Freund's (2000b) theoretical argument suggesting that RTA liberalization is more likely when MFN tariffs are low.[2] This is our second contribution. Our third contribution is to analyse the effect of MTAs on RTAs using

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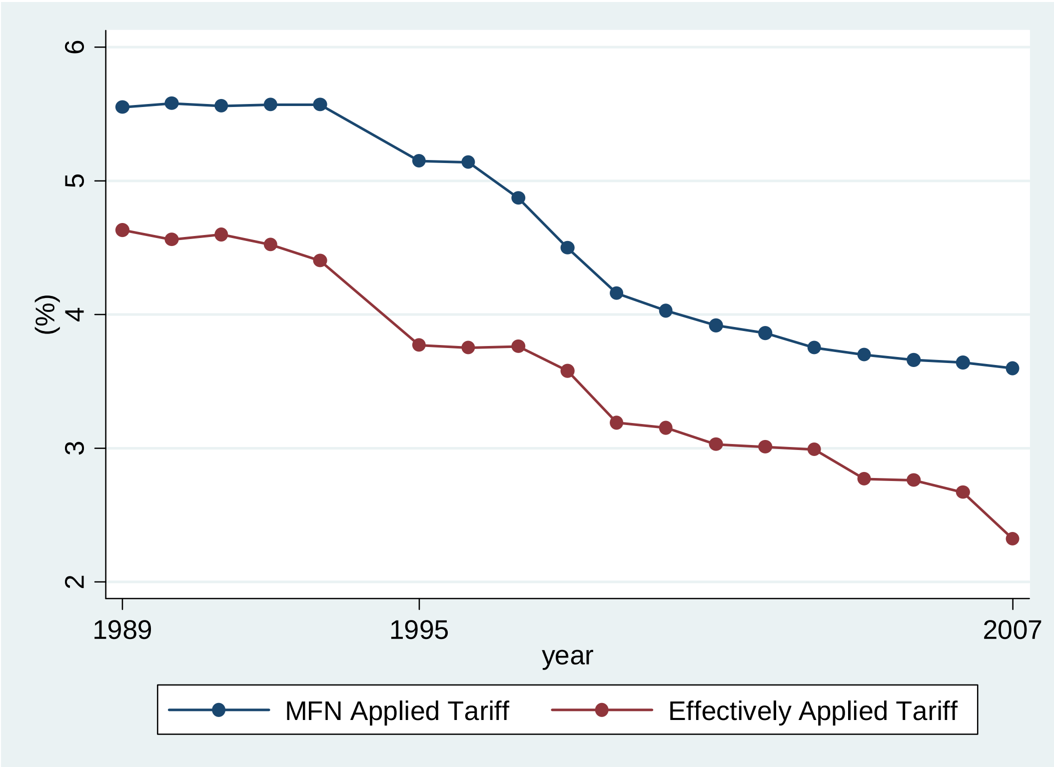
[1] This theory is usually cast in a two-country framework and thus is silent about the multilateral vs. preferential liberalization issue. It can thus guide the dynamic flavour of our empirical analysis.

[2] A lower MFN tariff expands imports from trading partners to which the United States applies the MFN tariff and reduces its imports from all others. As a result, the new RTA optimal tariff is lower. Since institutional constraints require the actual RTA tariff to be either 0 (if the good is excluded from the RTA) or equal to the MFN tariff (if the good is included in the RTA), the empirical counterpart of this theoretical effect is to increase the likelihood of this good being included in the list of goods that are liberalized swiftly.

differences rather than levels, thus avoiding the assumption of stationary tariffs (made previously in the literature), which is clearly incorrect for United States tariffs during that period (see figure 1).

Figure 1: United States Tariffs (Simple Means)[3]



Note: At the tariff line level, the effectively applied tariff corresponds to the lowest available tariff. Whenever it exists, the lowest preferential tariff is the effectively applied tariff. Otherwise it is the MFN applied tariff.

Section 5 then undertakes to establish the causal effect of MTA on RTAs. This is achieved by, first, choosing only RTAs signed by the United States whose negotiations were initiated well after the Uruguay Round of 1994 (a total of seven agreements signed between 2001 and 2006). Second, we add sector dummies to control for any political economy or other unobserved factors common to all goods in a sector which can affect the swiftness of liberalization across sectors. Third, we demonstrate that goods with high trade barriers due to non-tariff measures had no significant impact of MTA tariff-reduction on the swiftness of RTA liberalization. Fourth, we show that the emulator effect is weaker for goods with preference margins that do not bind because of prohibitive rules of origin. Fifth, we classify the goods in the sample according to the different stages of production in the value chain. Following the logic of the "Protection for sale" framework (Grossman and Helpman, 1994), downstream sectors oppose tariffs in upstream sectors from which they source (Gawande, Krisha and Olarreaga, 2009); conversely, upstream sectors support tariffs in downstream sectors to which they sell. Consistently, we find some evidence that the positive impact of MTA liberalization on the swiftness of RTA liberalization is weaker for Consumption and stronger for Equipment and Intermediate goods. Finally, we use hypothetic Uruguay Round

[3] In Figure 1, the "effectively applied tariff" series is a simple average of MFN and preferential tariffs across tariff lines. For institutional reasons specific to the United States, most of the preferential tariffs are zero.

tariff cuts to instrument for actual cuts: the overall aim of this Round was to achieve a 30 per cent cut in tariffs on manufacturing goods. Taken together, our results provide strong evidence that higher MTA tariff cuts increase the likelihood of immediate RTA liberalization and demonstrate that the effect is far from linear and apparently stronger for products that were swiftly liberalized in all seven RTAs considered in the paper versus those that were liberalized swiftly in six or less.

While we are able to demonstrate a clear link between greater MTA tariff reduction and a higher probability of immediate duty-free trade in more RTAs, we also exploit other RTA characteristics related to the timing of liberalization which could potentially be just as interesting. This is done, among other extensions, in section 6. For instance, we find that short MTA implementation periods are associated with swift RTA liberalizations: this is another manifestation of the emulator effect.

2. Related literature

Our findings speak to two different theoretical arguments put forth in the literature.

2.1. Are RTAs and MTAs substitutes or complements?

The first class of models studies the welfare effects of preferential versus multilateral trade liberalization and, on the positive side, whether liberalizing on a preferential basis first, by changing the status quo ante, undermines multilateralism. 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; Grossman and Helpman, 1995; Levy, 1997; Bagwell and Staiger, 1998; Krishna, 1998; Cadot, De Melo and Olarreaga 1999; McLaren, 2002; Saggi, 2006; Limão, 2007).[4] Limão (2006) finds empirical support for the stumbling block hypothesis for the United States case; Estevadeordal, Freund and Ornelas (2008) find a "building block" effect in a sample of 10 Latin American countries; Freund and Ornelas (2010) provide an excellent review of this abundant literature.[5] We complement it by asking the causality question in the opposite direction, as Ethier (1998) and Freund (2000a), but from an empirical angle.[6]

The models in this literature are essentially static: the supply side of the economy is exogenously given and the only dynamic thought experiment is an application of the agenda-setting game, a classic in political science. Aghion, Antràs and Helpman (2007) study this canonical game in a trade liberalization context explicitly. Freund (2000b) emphasizes that the

[4] Also, not one month elapses without the economic press worrying about this issue. Editorial lines predominantly echo the "stumbling block" hypothesis. For economic and political mechanisms consistent with the "building block" hypothesis, see e.g. Kennan and Riezman (1990), Richardson (1993), Bagwell and Staiger (1999) and Ornelas (2005a).

[5] Limão and Karakaovali (2008) find a stumbling block effect for the EU. Baldwin and Seghezza (2008) find a negative correlation between MFN tariffs and preference margins in their sample of 23 large countries. They conclude from this that the stumbling block mechanism, if it exists, is not of first order importance.

[6] Ethier (1998) analyses whether a multilateral trade agreement between developed countries promotes bilateral agreements between a developed country, which is part of the multilateral agreement, and a developing country, which is not part of the multilateral agreement. However, in the present analysis, all subsequent partners of the seven bilateral agreements are also members of the World Trade Organization (WTO), so Either's analysis is not suited to guide our empirical work.

same type of logic also entails that the incentives to form an RTA are shaped by the state of multilateral tariff levels. In an oligopolistic setting, she finds that the profit-shifting effect of regionalism, whereby discriminatory trade agreements expand output and profits in the participating countries at the expense of the countries left out, is especially strong when multilateral tariffs are low. She concludes from her analysis that "each Round of multilateral tariff reduction should lead to a new wave of RTAs" (Freund, 2000a: 359). Our results vindicate her conclusion. In a "Protection for sale" setting, Ornelas (2005a) points out that preferential trade liberalization erodes the rents from protection, which encourages participating countries to lower their external tariff. Insofar as this line of reasoning also applies in the opposite direction, our results are consistent with Ornelas' theoretical findings. A similar line of analysis asks whether the conditions under which RTAs are enforceable are affected by the multilateral trading environment (Freund, 2000b and Ornelas, 2005b). In these models, the static costs and benefits from protection are time-invariant by construction, so that natural solution to this kind of dynamic problem is a stationary tariff. In this, these papers are no different from existing theoretical studies on the complements-vs.-substitutes issue. Yet, if anything, tariffs fall over time and hence this line of explanation misses an important dimension of the real world.

2.2. Are past and current liberalization episodes complements or substitutes?

The "Juggernaut theory" of trade liberalization implies that current liberalization, by eroding protectionist forces and hence resistance to future trade reforms, is sowing the seeds of future liberalization (Baldwin, 1994; Staiger, 1995; Maggi and Rodrìguez-Clare, 2007; Baldwin and Robert-Nicoud, 2007). Our regression results provide (to the best of our knowledge: original) evidence consistent with the Juggernaut theory. A central insight in these papers recognizes that some sector-specific factors of production like (human) capital depreciate gradually over time; as a result, the politically optimal tariff is thus also decreasing over time. Freund (2000b) and McLaren (2002) also combine dynamic aspects of trade liberalization with the regionalism versus multilateralism issue but their focus (the hysteretic effects of preferential trade barriers) is different.

2.3. Relations with the empirical literature

From an empirical point of view, the main strand of the literature that relates to our research is on the determinants of RTAs formation. Several papers study the economic determinants of RTAs. The main identifying assumption remains that RTA-related trade gains are closely linked to the standard gravity covariates. Baier and Bergstrand (2004) find that the likelihood of an RTA is larger, the closer the two countries are to each other, the more remote they are from the rest of the world, the larger their GDPs, and so on. Building on Baier and Bergstrand (2004), Egger and Larch (2008) find evidence consistent with Baldwin's (1995) Domino theory of regionalism, whereby pre-existing RTAs increase the likelihood that two countries participate in a common RTA. In a separate but no less interesting line of research, Martin, Mayer and Thoenig (2009) find that multilateralism causes peace-motivated regional trade agreements.[7] The macro-level empirical evidence in these papers complements our micro-level evidence.

[7] The logic goes as follows: countries that have fought wars in the distant past tend to sign RTAs as a way of increasing the opportunity cost of a bilateral war, thereby reducing the probability that possible bilateral

Importantly, whereas we take the existence of the Free Trade Agreement as given, and aim to find out which tariff lines are liberalized the most swiftly, the three aforementioned papers aim to explain the formation of RTAs.

3. Definition of variables, data and summary statistics

In the case of the United States (and others), the legally binding and the applied MFN tariffs coincide exactly (by definition the latter may not be higher than the former), so we refer to them as the MFN tariff for short (World Tariff Profiles 2007). All United States MFN tariffs are non-increasing in the post-Uruguay Round period. Our key explanatory variable is a good-specific measure of the intensity of multilateral trade liberalization. We denote it by CUT_g with the subscript g referring to good g. CUT_g is defined as the (non-negative) difference (or tariff "cut") between the Tokyo and Uruguay MFN rates, i.e. $CUT_g \equiv MFN_g^{Tokyo} - MFN_g^{Uruguay}$. The stated aim of the Uruguay Round was to cut tariffs by about 30 per cent for industrial goods and bind the MFN tariff rate for all agricultural goods but in the end Canada, the EU, Japan, and the United States achieved a larger reduction on average (Baldwin, 2009).

Our main sources are the UNCTAD Trade Analysis and Information System (TRAINS) and the WTO Consolidated Tariff Schedules (CTS) Bound Duty Rates databases. Both databases provide information at the legal tariff line level (8-digit in the Harmonized System (HS) nomenclature), what we refer to as *goods*. They do not include goods subjected to non-ad valorem duties.[8] This leaves 9,303 goods. The WTO-CTS database provides information on bound rates negotiated at both the Tokyo and the Uruguay Rounds. Hence, CUT_g 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

In our analysis, we want to understand to what extent past multilateral trade liberalization is a factor towards current regional trade liberalization. A measure of the intensity of the regional trade liberalization similar in spirit to CUT_g is the *preference margin* $PM_{g,p}$, defined as the (non-negative) difference between the MFN tariff and the preferential tariff, or $PM_{g,p} \equiv MFN_g^{Uruguay} - PT_{g,p}$, where $PT_{g,p}$ is the good- and partner-specific *preferential tariff*. We exclude tariff lines for which the Uruguay MFN tariff was already zero, since no preference margin can be granted to such goods by definition. This leaves 7,419 goods in our reference sample.

The UNCTAD-TRAINS database includes MFN applied rates and preferential rates. The informed period is 1996–2008. This exhaustive database covers 15 trade agreements, from which we exclude trade agreements that were negotiated before the end of the Uruguay Round (1994) so as to eliminate an obvious source of reverse causality bias from our regressions (more on this in the next section); we also exclude unilateral trade agreements, for the focus of our analysis is on preferential trade liberalization or RTAs, not on unilateral ones. We are thus left with seven RTAs: Jordan (2001), Chile (2004), Singapore (2004), Morocco (2006), Bahrain (2006), Australia (2005)

conflicts might escalate into wars. In previous work (Martin et al. 2008), the same authors show that multilateral trade reduces the opportunity cost of a bilateral war. Taken together, this line of reasoning and these results imply that an increase of multilateralism raises the probability of bilateral war among old foes and they then enter bilateral or regional trade deals as an endogenous response to the threat it poses to bilateral peace.

[8] Such tariff lines account for around 8 per cent of the HS-6 subheadings of the World Tariff Profiles (2007).

and the Central American-Dominican Republic Trade Agreement (2006).[9] In our analysis, an "observation" is a good-and-partner entry for $PT_{g,p}$. Our reference sample has 51,814 observations, which is slightly lower than 7 x 7,419 = 51,933, because not all goods are included in all RTAs. Table 1 (panel a) breaks down the number of tariff lines included in our reference sample by partner. Table 1 (panel b) presents the summary statistics of our quantitative variables. For instance, the sample mean of CUT_g is 4,22 percentage points and the sample mean of $MFN_g^{Uruguay}$ is 6.2 percentage points.

Table 1: Descriptive Statistics

Panel (a) Tariff Lines in Trade Agreements

Partner	Tariff Lines Status: Immediate duty free	Gradual duty free	Total included	Excluded
Australia	5,319	1,591	**6,910**	509
Bahrain	5,306	2,113	**7,419**	None
Chile	6,651	733	**7,384**	35
Jordan	4,420	2,557	**6,977**	442
Morocco	5,397	1,979	**7,376**	43
Singapore	5,033	1,735	**6,768**	651
CAFTA	5,394	2,025	**7,419**	None

Panel (b) Variables

	Mean	Median	Standard Deviation	Min	Max
MFN tariff CUT, in pp (Tokyo minus Uruguay)	4.22	2.1	4.34	0	31.5
MFN tariff rate, in pp (Uruguay)	6.2	4.19	5.02	0.1	48
Share of imports (total) from RTA partners	.45	.23	.51	.005	1.31
Share imports (tariff line) from RTA partners	.21	0	2.63	0	100
Share imports from NAFTA partners	13.15	.73	24.09	0	100
Share exports to RTA partners	.91	.44	.89	.04	2.25

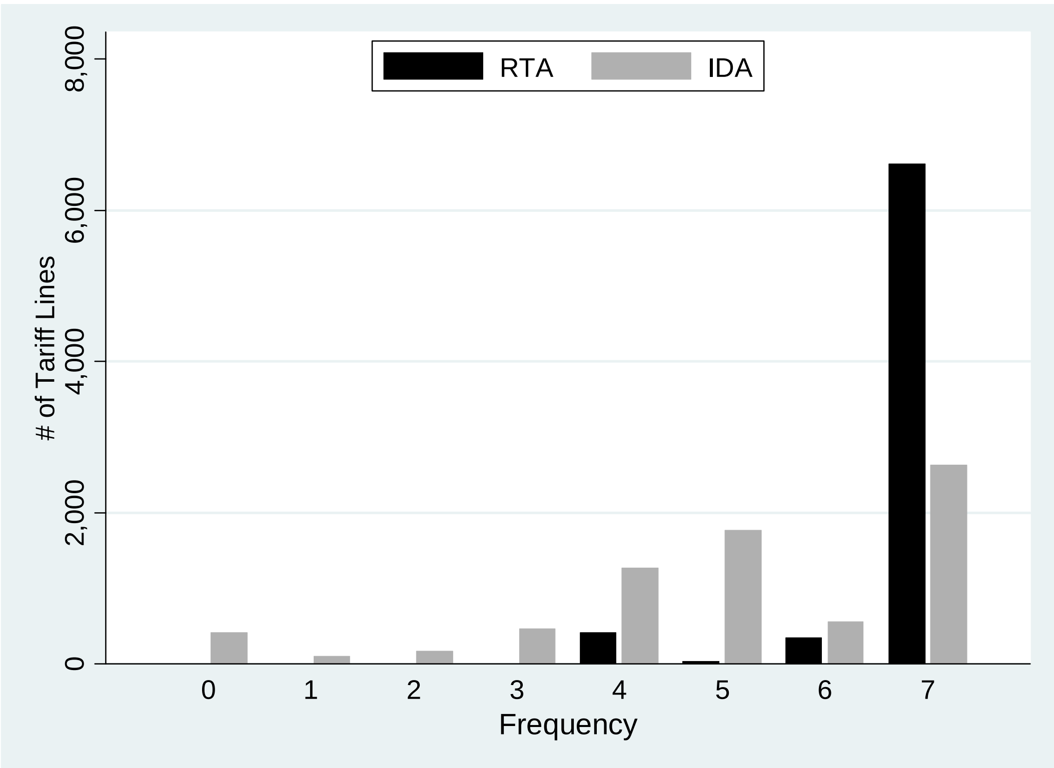
Note: All shares are calculated for the year 2000.

[9] That is, we exclude the Generalized System of Preferences (1976), Israel (1985), the Caribbean Basin Economic Recovery Act (1986), the Andean Trade Preference Act (1992), NAFTA (1994), the Generalized System of Preferences (GSP) for Least Developed Countries (1997), the African Growth and Opportunity Act (2000, 2001 and 2002), and the Caribbean Basin Trade Partnership Act (2000). See Romalis (2007).

It turns out that, in the United States case, each RTA is in fact a free trade agreement (FTA) de jure, namely, all tariffs on included goods eventually go to zero.[10] However, there is considerable variation in the timing of the implementation of this free trade policy about both goods and partners: overall, 69 per cent of our observations are fully liberalized at the start of the implementation of the RTA, whereas goods that are included in any of the RTAs but that are liberalized only gradually represent 27 per cent of our observations; the rest consists of good-partner pairs that are excluded from the corresponding RTA altogether (fewer than 4 per cent of observations).

Figure 2 illustrates various cross-RTA features of the sample. No tariff line has been included in fewer than four RTAs and the majority of them is part of all agreements (dark bars). Variation is clearly higher when considering the implementation of duty-free access (light bars). Many tariff lines (35 per cent) are set to zero on the date of entry into force of each and every trade agreement. Conversely, 6 per cent of some tariff lines are set to zero only gradually in all trade agreements. The remaining tariff lines are set to zero gradually in at least one but fewer than seven trade agreements.

Figure 2: Tariff lines in RTAs



Note: The RTA histograms refer to the number of tariff lines included in an RTA by frequency; "frequency" refers to the number of RTAs in which a given tariff line is being included. The IDA histograms refer to number of tariff lines granted IDA (Immediate Duty-free Access) status (i.e. tariff lines that are liberalized as an RTA enters into force).

[10] In a separate and fascinating line of research, Conconi, Fachini and Zanardi (2008) dig deeper into another peculiarity of the United States trade policy institutional setting: the fast track authority.

We also use the information available in the TRAINS database for non-tariff measures (NTM). We focus on NTMs classified as Technical Measures in the UNCTAD Coding System of Trade Control Measures (chapter 8). 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. Our control variables include imports at the tariff line; this information is also provided by UNCTAD-TRAINS. Table 1 (panel b) reports the summary statistics of the share of imports at the tariff line level that are covered by a preference margin as well as of the other controls.

4. Estimation strategy and estimation results

At a very general level, we would ideally like to regress the preference margin on both the multilateral CUTg (in first differences) and the most-favoured-nation tariff rate MFNg (in levels), that is, estimate an equation of the form

$$PM_{g,p} = \alpha + \beta_1 CUT_g + \beta_2 MFN_g + \varepsilon_{g,p} . \qquad (1)$$

The null hypothesis is that RTA liberalization is independent of MTA liberalization ($\beta_1 = 0$) and of MFN tariff levels ($\beta_2 = 0$). Our alternative hypothesis on β_1, which we dub as the "emulator effect", predicts a positive coefficient, whereas $\beta_1 < 0$ would be consistent with a *dynamic* version of the competing "money-left-on-the-table hypothesis".[11] Freund (2000a) guides our alternative hypothesis on β_2 (positive) and this competes with a *static* version of the "money-left-on-the-table hypothesis" ($\beta_2 < 0$) whereby there is more room to include a tariff line in an RTA if the MFN rate is relatively high to start with. Let us emphasize that MFN_g is orthogonal to CUT_g (the correlation is -.01 in our reference sample) so our empirical analysis is able to discriminate between competing hypotheses in *both* levels and first differences. This somewhat surprising feature of the data is also helpful for our identification strategy and we return to it in section 5.4.

The problem with a naïve estimation of the **intensive margin** of the emulator effect in (1) is that the United States institutional setting is such that a Preferential Trade Agreement is *de jure* a Free Trade Agreement. This makes using the intensive margin of preferential trade liberalization as the dependent variable problematic (at the end of the implementation period $PT_{g,p} = 0$, hence $PM_{g,p}$ boils down to $MFN_g^{Uruguay}$ by definition). For this reason we exploit instead its **extensive margin** and the timing of the preferential liberalization. Our first cut through the data is to set goods that are granted duty free access to the United States market immediately upon implementation of each of the seven RTAs in the sample apart from other goods. The idea is that these goods turned out to be the easiest to liberalize on a preferential basis and we want to understand the dimensions that make such goods special. Inspection of Figure 2 also shows that the most frequent number of times a good is granted "immediate duty-free access" (IDA) to the United States market is the maximum (seven). For these reasons, we create a binary variable for each good *g*, $SEVEN_g$, with $SEVEN_g = 1$ if good *g* is granted IDA status in the seven RTAs and 0 otherwise (i.e. if the good is granted only *gradual* duty-free access in, or excluded altogether from, at least one RTA); formally,

[11] Many commentators believe that the surge of regional trade agreements is the result of the stalling of the Doha Round: frustrated parties conclude bilateral agreements to substitute for the lack of a multilateral deal. The weekly *The Economist* is a major proponent of this thesis.

$SEVEN_g \equiv \mathrm{I}_7\left\{\# p : PT_{g,p}^{impl} = 0\right\}$, where *impl* denotes the implementation year and $\mathrm{I}_7\{\cdot\}$ denotes an indicator function that takes value 1 if its component is equal to seven and 0 otherwise.[12] We also create two additional measures along those lines, the binary variable ONE_g that takes value 1 if good g gets IDA status in at least one RTA and 0 otherwise and the count variable NTL_g that counts the number of RTAs in which good g gets IDA; these being mostly robustness checks, we postpone the regression results for ONE_g and NTL_g to Section 6.

As our second measure of the extensive margin of preferential trade liberalization, we define a good- and partner-specific measure of preferential trade liberalization for our central specification that takes value 1 if imports of good g from partner p are granted the IDA status upon implementation of the RTA in question and zero otherwise.

4.1. Evidence at the good level: Logit

We start by running the following logit:

$$\Pr\{SEVEN_g = 1\} = \Lambda\left(f_{G(g)} + \beta_1 CUT_g + \beta_2 MFN_g^{Uruguay} + \mathbf{X_{g,p}\beta}\right), \tag{2}$$

where $\Lambda(\cdot) \equiv \exp(\cdot)/\left[1+\exp(\cdot)\right]$ is the logistic cumulative distribution function, $f_{G(p)}$ is a sector dummy, CUT_g is the reduction in the MFN tariff negotiated over the course of the Uruguay Round (in percentage points), $MFN_g^{Uruguay}$ is the ad-valorem Uruguay MFN tariff rate (in percentage points) and $\mathbf{X_{g,p}}$ is a set of additional controls; β_1 is our coefficient of interest. Denote the set of all goods by $\Gamma = \left\{1,\ldots,N_g\right\}$; then G is a partition of Γ and we use $G(g)$ to denote the HS-2 sector in which good g is classified. Thus, G is also a mapping $G : \text{good} \to \text{sector}$.

Though we view (2) as a reduced form relationship between $SEVEN_g$ and CUT_g, we must assume that CUT_g is exogenous in order to obtain consistent and unbiased estimates of the coefficients. Our strategy to rid ourselves of the reverse causation bias rests partly on the timing of events. We limit our sample to the seven RTAs that entered into force after the conclusion of the Uruguay Round in 1994. This sample selection is expected to eliminate any reverse-causality bias for two main reasons: first, no new multilateral trade agreements had been implemented by the United States between 1994 and 2000. This buffer is likely to be long enough to ensure that these trade agreements to come did not influence the Uruguay Round trade negotiators. The second reason reinforces this point: no trade agreement signed in the post-Uruguay Round period had actually been negotiated during 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, no agreement has yet been reached in the context of FTAA negotiations. In addition, the APEC forum held in Bogor in 1994 signed a declaration to work toward free trade in the region by 2010 for

[12] A comment about goods-partner pairs that do net get the IDA status is in order here. Goods g that are included in the RTA p but that are liberalized only gradually and goods that are excluded from that RTA altogether are both coded the same way. This is because the frequency of the latter in the data is very low (less than 5 per cent of good-partner pairs). Our qualitative results do not change if we drop these observations from the sample.

developed countries and by 2020 for all member-countries. A 16-year time frame makes any influence of those talks on tariff cuts defined the Uruguay Round quite implausible.[13] Note that the absence of correlation between CUT_g and $MFN_g^{Uruguay}$ is also helpful: it implies that the past determinants of trade liberalization (at the good level) that cumulated to give rise to the Tokyo tariff *level* are different from those that led to the Uruguay Round tariff *cut*: in line with the Juggernaut hypothesis, this suggests that the sectoral determinants of tariffs are not as long-lived as one might think. However, if an omitted variable affects $SEVEN_g$ and CUT_g simultaneously, then regressing the former on the later will cause a spurious correlation. We thus introduce sector dummies $f_{G(g)}$ in (2) to capture sector invariant sources of unobserved heterogeneity such as the political economy determinants of tariffs or the sources of comparative advantage.[14] Insofar as such unobserved shocks are common to goods within sectors, then including $f_{G(g)}$ in (2) corrects for this source of omitted variable bias in our cross section exercise.[15] Together, these three working assumption constitute our maintained identification hypothesis. We complement them with additional approaches in Section 5.

We use sector fixed effects at a relatively high degree of aggregation so that our sample has a large number of observations for each partner p and for each sector G; as a result, the β's in the conditional logit in (3) are consistently estimated.

Table 2 presents the results. We report odds ratios throughout (standard errors clustered at the tariff line in parenthesis). The odds ratio associated to β_j is defined as exp β_j (j = 1,2,...) and has the meaning that a one extra percentage point in CUT_g raises the probability of granting IDA status to all partners for the good in question by a factor exp β_j *relative to* not including the tariff line or delaying setting this preferential tariff to zero. The two independent variables of interest, CUT_g and $MFN_g^{Uruguay}$, are significant beyond the one per cent level in all specifications and the results are stable across specifications. The regression in Column (1) includes the two independent variables and Column (2) adds sector dummies. The findings are consistent with the emulator hypothesis: the odds ratio implies that one extra percentage point of CUT_g raises the probability that good g gets IDA treatment for all of the United States' RTA partners by almost a fourth (1.227 – 1 = .227) relative to getting it only for a subset of those. By contrast, the "money-left-on-the-table hypothesis" is rejected by the data: raising $MFN_g^{Uruguay}$ by one percentage point *decreases* the odds that good g gets IDA status by a third (1 – .657 = .343). This result is thus empirical evidence in favour of Freund (2002a).

[13] What is usually recognized is that the APEC summit together with NAFTA helped "squeeze the European Union to complete the Uruguay round of GATT" in the words of Robert Zoellick's (2001) statement as United States Trade Representative.

[14] The chosen level of disaggregation corresponds to standard practice in the literature (Limão, 2006). However, we also check the robustness of our benchmark specifications using HS-4 sectoral dummies. Coefficients vary only slightly indicating that HS-2 sectoral level dummies are sufficient to capture major sources of unobserved heterogeneity. We opted for the HS2- level dummies in order to gain computational flexibility in implementing our sensitivity analysis presented in Section 6.

[15] See also Broda, Limão and Weinstein (2008) on this.

Table 2: LOGIT "Seven"

	Dependant variable: *SEVEN* (Probability that tariff line *g* is granted IDA to United States market to all 7 partners)				
	(1)	(2)	(3)	(4)	(5)
Tariff CUT (Tokyo minus Uruguay)	1.140[a] (0.00826)	1.227[a] (0.0109)	1.330[a] (0.0158)	1.331[a] (0.0158)	1.313[a] (0.0159)
MFN tariff rate	0.668[a] (0.0127)	0.657[a] (0.0165)	0.612[a] (0.0174)	0.612[a] (0.0175)	0.611[a] (0.0175)
DIFF0 (no Uruguay Round cut)			4.375[a] (0.459)	4.378[a] (0.459)	4.253[a] (0.446)
Share imports from RTA partners				1.019 (0.0351)	1.010 (0.0341)
Share imports from NAFTA partners					0.992[a] (0.00162)
Sector FE	No	Yes	Yes	Yes	Yes
Observations	7419	6822	6822	6822	6822
Pseudo R^2	0.209	0.294	0.318	0.318	0.321
LL	-3815.2	-3206.3	-3099.7	-3099.5	-3085.6

Notes: **Coefficients**: Exponentiated (odds ratios); **Robust standard errors** in parentheses.
[a] $p < 0.01$, [b] $p < 0.05$.

In Column (3), we add a good-specific dummy $DIFF0_g$ that takes value $DIFF0_g = 1$ if the United States did not liberalize good g during the Uruguay Round (i.e. if $CUT_g = 0$) and zero otherwise.[16] That is, we estimate

$$\Pr\left\{SEVEN_g = 1\right\} = \Lambda\left(f_{G(g)} + \beta_1 CUT_g + \beta_2 MFN_g^{Uruguay} + \beta_3 DIFF0_g\right).$$

The fact that goods that were not liberalized during the Uruguay Round – because these sectors are better organized and successfully fought to be left out of the Uruguay Round entirely, say – might be quite different from other goods motivates this specification. The coefficient β_3 is positive at the 1 per cent level, implying that goods that were not liberalized at the multilateral level were more likely to be liberalized at the preferential level: on its own, this result is consistent with a dynamic version of the "money-left-on-the-table hypothesis". Adding this control also raises the odds ratio of CUT_g to 1.33. Overall, the effect of CUT_g on IDA treatment thus seems to be non monotonic: the United States grants IDA status more frequently for goods for which the Uruguay Round tariff cut was zero as well as for those that had a large CUT_g

In order to quantify this non-monotonic effect, we replace *DIFF0* by a quadratic term to (2) and we compute the marginal effect of CUT_g for all observations/goods. Specifically, we first run

[16] This is verified for 21.8 per cent of the tariff lines in our reference sample.

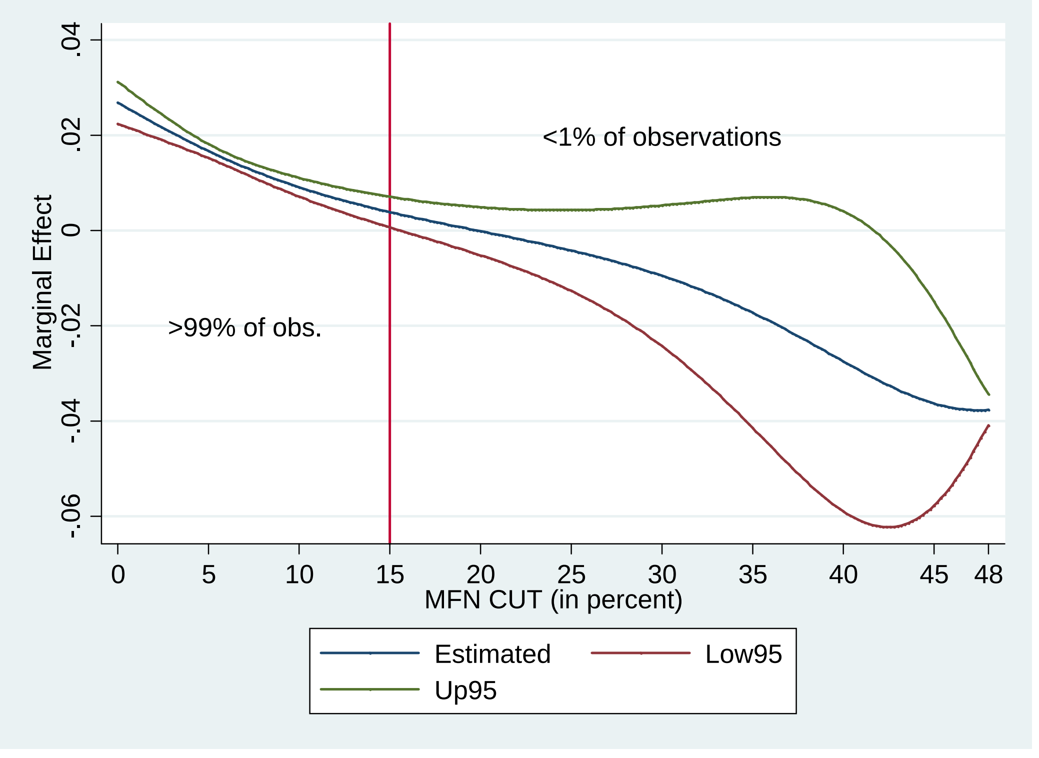
$$\Pr\left\{SEVEN_g = 1\right\} = \Lambda\left(f_{G(g)} + \beta_0 CUT_g^2 + \beta_1 CUT_g + \beta_2 MFN_g^{Uruguay}\right)$$

and we obtain that the odd ratios associated with β_0, β_1 and β_2 are 0.997 (0.007), 1.144 (0.012) and 0.931 (0.004), respectively (t-statistics in parenthesis), which confirms the non-monotonicity uncovered in the previous specification. In order to quantify this non-monotonic effect, we compute the marginal effect as

$$\frac{\partial}{\partial CUT_g}\Pr\left\{SEVEN_g = 1\right\} = \Lambda_g^{'}(\cdot)\left[2\beta_0 CUT_g + \beta_1\right],$$

where $\Lambda_g^{'}(\cdot)$ is the density of the logistic distribution $\Lambda(\cdot)$ evaluated at the explanatory variables pertaining to observation *g*. Figure 3 plots the estimated values of the marginal effect as well as the 95 per cent confidence interval against CUT_g. As is obvious from the figure, the dynamic version of the "money-left-on-the-table hypothesis" is rejected in 99.99 per cent of cases (i.e. for *all but seven* observations out of 51,814). By contrast, the data are consistent with the emulator effect (i.e. the net effect is statistically significantly positive) in 99.41 per cent of cases. We conclude from this that the data provide strong support for the emulator hypothesis and reject the "money-left-on-the-table hypothesis". To sum up, the average, median and net effects are all consistent with the emulator hypothesis.

Figure 3: Estimated Marginal Effects with Quadratic CUT_g Term



Note: The net marginal effect of the CUT variable is significantly positive for 99.4% of observations. It is not significantly different from zero for 0.7 per cent of the observations.

The results reported in Columns (4) and (5) show that these qualitative findings are robust to the inclusion of several controls. Column (4) introduces the import share of all seven partners in United States total imports of good g, defined as $SM_g \equiv \sum_p M_{g,p} / M_g$ (where M denotes the value of imports observed in the year 2000), to control for the possibility that the United States might be granting IDA access to prominent exporters more easily. The estimated coefficient in Col. (4) is statistically insignificant: thus, the United States does not seem to discriminate between large and small exporters when granting IDA status.[17]

Column (5) adds $SNAFTA_g$ to the set of controls, with $SNAFTA_g$ being defined as the good-specific import share of NAFTA products in year 2000, i.e. $SNAFTA_g \equiv M_{g,NAFTA} / M_g$. Its coefficient is statistically negative at the one-percent level (its odds ratio is lower than unity), implying that the United States is less likely to grant IDA status from markets that NAFTA already penetrates widely. This suggests that NAFTA and ensuing RTAs are substitutes, that is, NAFTA worked as a 'stumbling block' to post-Uruguay Round regionalism. Col. (5) forms our baseline specification henceforth.

4.2. Evidence at the good-partner level: Logit

The evidence so far indicates that CUT_g and $MFN_g^{Uruguay}$ influence the extensive margin of preferential trade liberalization. The evidence portrayed is at the good level. However, we can address a more demanding question to the data: given some other good characteristics (observable or not), how do CUT_g and $MFN_g^{Uruguay}$ influence the likelihood that the United States grants IDA status to partner p's exports of good g to the United States? For this purpose, we create a good-partner indicator variable $IDA_{g,p} \equiv \mathrm{I}\left\{PT_{g,p}^{impl} = 0\right\}$ that takes value 1 if partner p gets immediate duty-free access to the United States market for good g and zero otherwise. We then estimate the following logit:

$$\Pr\left\{IDA_{g,p} = 1\right\} = \Lambda\left(f_p + f_{G(g)} + \beta_1 CUT_g + \beta_2 MFN_g^{Uruguay} + \mathbf{X_{g,p}\beta}\right), \tag{3}$$

where f_p is a partner dummy and the other right-hand side variables are as in (2).[18] Running (3) is similar to running (2) at the good-partner level. The implicit assumption in (3) is that the functional form that maps the right-hand-side variables into $IDA_{g,p}$ is symmetric for each partner. As we shall see, though, the effect of CUT_g on $IDA_{g,p}$ is non-linear. For this reason, we consider running (3) as a conservative robustness check that provides a lower bound for the emulator effect.

With this caveat in mind, turn to Table 3, which reports the results (standard errors clustered at the tariff line in parenthesis). The qualitative results are in line with those of Table 2. The coefficients for CUT_g, $MFN_g^{Uruguay}$, $DIFF0_g$ and $SNAFTA_g$ are still precisely estimated and they have the expected sign.

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[17] We use a qualitative variable to discriminate between goods with zero imports and those with positive imports in Section 6.5 below.

[18] Preferential trade agreements can be motivated by non trade objectives as argued in Limão (2007). The inclusion of partner dummies in specification (3) absorbs any effect possibly related to such non trade objectives. Also, our set up is cross-sectional while trade agreements were signed in different years; the inclusion of partner dummies also absorbs factors that may cause a specific sequence in bilateral trade talks.

Table 3: p-g LOGIT

	Dependant variable: Pr{IDA = 1} (Probability that tariff line g is granted IDA to United States market to partner p)					
	(1)	(2)	(3)	(4)	(5)	(6)
Tariff CUT	1.064[a]	1.099[a]	1.125[a]	1.126[a]	1.115[a]	1.115[a]
(To. minus Ur.)	(0.0162)	(0.0197)	(0.0221)	(0.0220)	(0.0212)	(0.0213)
MFN tariff	0.922[a]	0.931[a]	0.926[a]	0.925[a]	0.930[a]	0.930[a]
level	(0.0119)	(0.0125)	(0.0139)	(0.0136)	(0.0134)	(0.0134)
DIFF0 (no Uruguay			1.683[a]	1.688[a]	1.623[a]	1.623[a]
Round cut)			(0.316)	(0.316)	(0.296)	(0.298)
Partner's				1.039[a]	1.039[a]	1.041[a]
share of M_g				(0.0144)	(0.0152)	(0.0128)
Share imports					0.996[a]	0.996[a]
from NAFTA partners					(0.00103)	(0.00103)
SALL: Partner's						0.951
share of United States X+M						(0.160)
Sector FE	No	Yes	Yes	Yes	Yes	Yes
Partner FE	No	Yes	Yes	Yes	Yes	No
Observations	51814	51814	51814	51814	51814	51814
Pseudo R^2	0.044	0.115	0.119	0.120	0.085	0.086
LL	-29248.8	-27064.3	-26942.2	-26909.6	-28003.2	-27973.3

Notes: **Coefficients**: Exponentiated (odds ratios); **Robust standard errors** (clustered by tariff line) in parentheses. [a] $p < 0.01$, [b] $p < 0.05$.

Running (3) enables us to control explicitly for partner and good-partner characteristics. Thus, let $SM_{g,p} \equiv M_{g,p} / M_g$ define the share of good-*g* imports that are sourced in country *p*. What are our priors on the sign of its coefficient? In Grossman and Helpman's (1994) 'protection for sale' (PFS) framework, keeping the elasticity of imports and the domestic production constant (both vary across goods but are constant across partners), protection decreases in the volume of imports (which does vary across partners) in organized sectors. In non-organized sectors, the opposite is true. Estimation of

$$\Pr\left\{IDA_{g,p} = 1\right\} = \Lambda\left(f_p + f_{G(g)} + \beta_1 CUT_g + \beta_2 MFN_g^{Uruguay} + \beta_3 DIFF0_g + \beta_4 SM_{g,p}\right)$$

includes neither domestic production nor import elasticities. The former omission is harmless: for each good, there are several import sources (the partners) and possibly a different preferential tariff for each of them; this enables us to estimate β_4 via the cross-sectional variation of $SM_{g,p}$ along the *p*-dimension. The latter, however, introduces measurement error in the estimation of β_4. Also, the left-hand side of the structural PFS model is different from the LHS of (3). With these caveats in mind, the estimated coefficient in column (5) of Table 3 is statistically positive at the one-percent level. This is consistent with the PFS qualitative prediction for *organized* sectors. This finding is important for the interpretation of the emulator effect as evidence of the juggernaut mechanism. The estimated odds ratio corresponding to β_4 is equal to 1.04, which implies that an increase in the import penetration ratio of the pair (g, p) by 1 per cent increases the odds of the United States granting IDA status to p's exports of good g by 4 percentage points. In other words, the United States grants IDA status disproportionately to important import sources. The estimated coefficient is stable across specifications.

We might also expect the United States to grant tariff-free access to important trading partners as part of broader foreign and trade policy objectives. To check whether this intuition is verified in the data, we introduce the Partner's share of imports across all tariff lines as a an additional control in (3), namely $SMALL_p \equiv \sum_g M_{g,p} / M$, as well as the United States' share of exports towards p, defined as $SXALL_p \equiv \sum_g X_{g,p} / X$, where X denotes exports observed in the year 2000. In the same spirit, we also create $SALL_p$ as $SALL_p \equiv \sum_g (M_{g,p} + X_{g,p}) / (M + X)$ as an overall measure of the importance of p as a trading partner for the United States. $SALL_p$, $SMALL_p$ and $SXALL_p$ are defined at the partner level, so we drop the partner dummy in these regressions. Column (6) reports the results for $SALL_p$ (the results for $SMALL_p$ and $SXALL_p$ are similar and so we omit them). The estimated coefficient is statistically indistinguishable from zero, rejecting the hypothesis that the United States grants free access to its markets disproportionately to large partners.

Table 4: g-Logit on Partner-Specific Sub-sample

	Dependant variable: Pr{IDA = 1} (Probability that tariff line g is granted IDA to United States market to partner *p*)						
	(AUS)	(BHR)	(CHL)	(JOR)	(MAR)	(SGP)	(CAFTA)
Tariff CUT (To. minus Ur.)	1.075[b] (0.0313)	1.261[a] (0.0411)	1.120[a] (0.0448)	1.197[a] (0.0318)	1.090[b] (0.0369)	1.175[a] (0.0309)	1.273[a] (0.0449)
MFN tariff rate	0.815[a] (0.0342)	0.956[a] (0.0142)	0.895[a] (0.0277)	0.687[a] (0.0418)	0.878[a] (0.0282)	0.640[a] (0.0720)	0.968 (0.0207)
DIFF0 (no Uruguay Round cut)	2.110[a] (0.577)	2.440[a] (0.715)	1.862 (0.710)	2.902[a] (0.997)	3.097[a] (1.099)	2.389[b] (0.817)	2.410[a] (0.715)
Share imports from RTA partners	1.017 (0.0176)	38.49 (115.5)	0.971[b] (0.0112)	1.083 (0.151)	1.057 (0.0351)	0.998 (0.00926)	1.019[b] (0.00970)
Share imports from NAFTA partners	0.995[b] (0.00210)	0.995 (0.00242)	0.997 (0.00306)	0.992[a] (0.00222)	0.997 (0.00220)	0.995[b] (0.00211)	0.996 (0.00256)
Sector FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	6929	7287	6420	7332	6474	6771	7246
Pseudo R^2	0.463	0.180	0.207	0.343	0.453	0.341	0.184
LL	-2278.5	-3589.8	-1845.3	-3254.6	-2006.0	-2889.9	-3494.1

Notes: **Coefficients:** Exponentiated (odds ratios); **Robust standard errors** in parentheses.
[a] $p < 0.01$, [b] $p < 0.05$.

Finally, we re-run (3) for each partner separately (more precisely, the specification corresponding to Table 3, Col. 5). Table 4 reports the results. The coefficients of CUT_g and $MFN_g^{Uruguay}$ have the expected signs. The emulator effect is economically and statistically weakest for Australia and Morocco and especially large for CAFTA. The "money-left-on-the-table hypothesis" is rejected in all cases, albeit only in a weak sense in the case of CAFTA.[19]

5. Identification of the "Emulator Effect"

The "emulator" effect seems to be a robust feature of the data, unlike the "money-left-on-the-table" argument. We have so far relied mostly on the timing of events to identify the effect. In this section, we use the interaction between our variable of interest (CUT_g) and non-tariff measures (Section 5.1), the rules of origin (Section 5.2) or the type of goods (Section 5.3) to interpret the positive correlation between CUT_g and IDA in a causal way. Finally, we instrument for CUT_g (Section 5.4).

5.1. Non-tariff measures

We start by controlling for the presence of non-tariff measures, or "NTM", at the tariff line.[20] The idea is that the presence of such non-tariff measures should weaken the effect of CUT_g on preferential liberalization: a multilaterally agreed tariff cut is less effective if the imports of that good are impeded by other measures. We thus expect the CUT_g coefficient to be larger for NTM-free goods than for goods with some NTM. To test this idea, we create a dummy variable NTM_g that takes value one if the tariff line g has some NTM and zero if good g is NTM-free.

We first re-run (2), adding the NTM_g dummy and its interaction with CUT_g. Table 5, Col. (2) reports the results; these have to be compared with Col. (1), which reports the odds ratios of our baseline specification (Table 2, Col. 5). As expected, the CUT_g coefficient for NTM-free goods is (much) larger than for NTM goods; the difference is significant at any conventional level. The coefficient for CUT_g in goods with non-tariff measures is insignificant (the odds ratio is one). This finding is exactly what we should expect if multilateral and preferential tariff cuts are dynamic complements and if the presence of NTMs prevents the emulator effect from playing its role. We repeat this exercise for the good-partner specification (3) and the results, reported in Table 5, Col. (4), do not affect these conclusions.[21] These findings thus vindicate our emulator hypothesis further. By contrast, the odds ratio of MFN is reduced in this specification, weakening further the "money-left-on-the-table hypothesis".

[19] We also run (3) on the full sample and with country specific *CUT* coefficients to be estimated. Results indicate that they are pairwise statistically different. The larger estimate of the odds ratio is obtained for the CAFTA agreement (1.20), followed by Bahrain (1.18), Jordan (1.14), Australia (1.11), Morocco (1.09), Chile (1.08) and Singapore (1.06).

[20] There are 19 per cent of tariff lines with an NTM in our reference sample.

[21] Table 5, Col. (3) reproduces Table 3, Col. (5) to ease comparison.

Table 5: Non-tariff Measures (NTM)

	Dependant variables: SEVEN		Pr{IDA = 1}	
	(1)	(2)	(3)	(4)
Tariff CUT (To. minus Ur.)	1.313[a] (0.0159)		1.115[a] (0.0212)	
NTM * cutMFN		1.010 (0.0375)		0.993 (0.00689)
(1-NTM) * cutMFN		1.310[a] (0.0155)		1.140[a] (0.00455)
MFN tariff rate	0.611[a] (0.0175)	0.603[a] (0.0173)	0.930[a] (0.0134)	0.924[a] (0.00261)
DIFF0 (no Uruguay Round cut)	4.253[a] (0.446)	4.173[a] (0.431)	1.623[a] (0.296)	1.700[a] (0.0583)
NTM dummy	No	Yes	No	Yes
PartnerFE	N.A.	N.A.	Yes	Yes
Observations	7419	7419	51814	51814
Pseudo R^2	0.327	0.329	0.124	0.129
LL	-3056.2	-3046.0	-26810.9	-26652.3

Notes: **Coefficients**: Exponentiated; **Robust standard errors** in parentheses. [a] $p < 0.01$, [b] $p < 0.05$. All regressions include sector dummies and the controls *Share imports from RTA partners* and *Share imports from NAFTA partners*. The dummy NTM takes value one whenever a NTM is applied at the tariff line. NTM*CUT represents the interaction between the NTM dummy and the variable Tariff CUT.

5.2. Unused rules of origin

It is well known that the compliance costs of rules of origin (RoO henceforth) can be prohibitive (Krishna, 2006). Specifically, when the preference margin is low, foreign exporters might not bother with complying with rules of origin. In our setting, the preference margin is the MFN tariff rate. If the emulator effect is the manifestation of an actual economic mechanism whereby trade agreements are dynamic complements, then we expect the coefficient of CUT_g to be higher for the goods where the rules of origin are actually exploited by foreign exporters. Preference margins are irrelevant when below 2 to 3 percentage points (Estevadeordal *et al.*, 2008). To identify this differential effect in the data, we construct a dummy variable RoO_g that takes value 1 if $MFN_g > 2.5$ (when foreign exporters are expected to use the preference and thus to comply with the rules of origin) and zero otherwise and we re-run (2) and (3) with this dummy as an additional control variable. We expect the CUT_g coefficient to be larger for RoO-goods than for goods that have irrelevant rules of origin.

Table 6: Unused Rules of Origin (RoO)

	Dependant variables: SEVEN		Pr{IDA = 1}	
	(1)	**(2)**	**(3)**	**(4)**
Tariff CUT (Tokyo minus Uruguay)	1.321[a] (0.0165)		1.120[a] (0.00411)	
RoO * CUT		1.374[a] (0.0181)		1.169[a] (0.0107)
(1-RoO) * CUT		1.309[a] (0.0328)		1.113[a] (0.00425)
MFN tariff rate	0.551[a] (0.0216)	0.553[a] (0.0228)	0.927[a] (0.00270)	0.928[a] (0.00269)
DIFF0 (no Uruguay Round cut)	4.358[a] (0.453)	4.239[a] (0.439)	1.666[a] (0.0580)	1.636[a] (0.0571)
RoO dummy	Yes	Yes	Yes	Yes
PartnerFE	N.A.	N.A.	Yes	Yes
Observations	6822	6822	51814	51814
Pseudo R^2	0.329	0.329	0.121	0.122
LL	-3049.1	-3046.0	-26876.9	-26861.0

Notes: **Coefficients**: Exponentiated; **Robust standard errors** in parentheses. [a] $p < 0.01$, [b] $p < 0.05$. All regressions include sector dummies and the controls *Share imports from RTA partners* and *Share imports from NAFTA partners*. The dummy RoO takes value 1 when MFN values are above or equal to the 2.5% threshold and zero otherwise. RoO*CUT represents the interaction between the RoO dummy. and the variable Tariff CUT.

Table 6, Col. (2) reports the results for (2), which have to be compared with those of the baseline specification, reproduced in Col. (1). The results are supportive of the emulator hypothesis: as expected, the CUT_g coefficient is larger for the goods for which it matters than for goods with an irrelevant preference margin. By contrast, the coefficient and the odds ratio for $MFN_g^{Uruguay}$ shrink noticeably, rejecting the "money-left-on-the-table hypothesis" further.

Table 6, Col. (4) reports the results for (3), which have to be compared with those of Col. (3). Here, the results are as again supportive; the Wald statistics rejects the hypothesis that the coefficients are the same at the one-percent level. We have also re-ran (2) and (3) with 2 and 3 percentage points as thresholds (results not reported); the qualitative results were not affected.

In sum, the differential effect of CUT_g on granting *IDA* status for goods affected by rules of origin or non-tariff measures that we find in the data confirms this set of predictions of the emulator hypothesis.

5.3. The role of intermediate goods

As we shall see in Section 6, the emulator effect is non-linear. Specifically, the largest emulator effect is between granting this preferential access to *all* partners or not, rather than between *some* partners or none. This in turn suggests that the *type* of goods might be more important than the partners' characteristics; also, when we include sector dummies in our regressions, the coefficients of interest tend to rise in a significant way, suggesting that unobserved sector-invariant characteristics are indeed important. Therefore, we split the sample among the following categories of goods that correspond to different stages of production in the value chain: Basic manufacturing, Consumption goods, Equipment goods, Intermediate goods, Mixed products and Primary goods and we estimate one β_1 for each category in our baseline regression (with $MFN_g^{Uruguay}$ and $DIFF0_g$ as controls).[22] Table 7 reports the results. The estimated coefficients are positive and significant at the one-percent level in all cases but for consumption and primary goods, for which it is insignificant. It is particularly strong for equipment and intermediate goods and weakest for consumption and primary goods.

Table 7: LOGIT "Seven" by Type of Goods

	Dependant variable: SEVEN (Probability that tariff line g is granted IDA to United States market to all 7 partners)					
	Basic-manufacturing	**Consumption-goods**	**Equipment-goods**	**Intermediate-goods**	**Mixed-products**	**Primary**
Tariff CUT (To. minus Ur.)	1.423[a] (0.0433)	1.181[a] (0.0572)	1.306[a] (0.0426)	1.343[a] (0.0404)	1.404[a] (0.0613)	1.061 (0.102)
MFN tariff rate	0.561[a] (0.0301)	0.494[a] (0.0407)	0.838[a] (0.0368)	0.445[a] (0.0344)	0.808[a] (0.0353)	0.201[a] (0.0632)
DIFF0 (no Uruguay Round cut)	18.62[a] (5.180)	1.675 (0.529)	3.080[a] (0.667)	2.493[a] (0.711)	5.951[a] (1.785)	2.53e-09 (0.00031)
Share imports from RTA partners	1.018 (0.0716)	1.085 (0.103)	1.366 (0.260)	0.676 (0.229)	0.679[b] (0.125)	1.257[b] (0.121)
Share imports from NAFTA partners	0.996 (0.00352)	0.986[a] (0.00468)	0.976[a] (0.00519)	0.994 (0.00434)	0.995 (0.00385)	0.990 (0.0134)
Sector FE	Yes	Yes	Yes	Yes	Yes	Yes
Observations	1598	1031	859	1029	691	132
Pseudo R^2	0.313	0.480	0.226	0.361	0.222	0.669
LL	-726.4	-335.9	-457.6	-437.3	-335.2	-28.68

Notes: **Coefficients**: Exponentiated (odds ratios); **Robust standard errors** in parentheses. [a] $p < 0.01$, [b] $p < 0.05$.

These results are less helpful in the quest of identifying the emulator hypothesis. To see why, recall that in our interpretation of the dynamic complementarity between trade agreements, past trade liberalization in a given sector undermines its current resistance to trade openness because trade liberalization decreases the (quasi) rents associated with the (quasi) fixed factors that

[22] The most represented categories of goods are Equipment, Consumption and Intermediate goods categories. The goods that are liberalized the most systematically belong to the Equipment goods followed by Basic and Primary goods. The largest average MFN cut is obtained for Primary goods followed by Basic and Equipment goods. When considering only the set of goods included in all seven RTAs ($SEVEN_g$=1), Equipment, Intermediate and Basic goods categories show the largest average MFN cuts. *Source:* United Nations statistical division (2007).

fight for protection. Over time, these factors depreciate and with them the resistance to trade liberalization. By the same logic, downstream sectors oppose tariffs in upstream sectors from which they source, and this opposition increases as downstream tariffs fall; also, upstream sectors that sell domestically have an interest in keeping downstream tariffs high (Gawande, Krishna and Olarreaga, 2009). As a result, we expect the emulator effect to be strongest in upstream sectors, i.e. for primary products, intermediate goods and capital goods ("equipment goods"), and weakest in downstream sectors ("consumption goods").[23] With the noticeable exception of primary products, the results in Table 7 are in line with those priors.

5.4. IV estimation

Finally, we use some exogenous sources of variation for our key right-hand side variable, CUT_g, to identify the causal effect of CUT_g on the dependent variable ($SEVEN_g$ or $\Pr\{IDA_{g,p} = 1\}$, depending on the specification). We have initially experimented with three instruments: the corresponding EU's MFN tariff cut, the share of EU and Japanese exports in United States imports at the tariff line and a theoretical MFN cut. Standard redundancy tests led to the exclusion of the first two instruments. Their exclusion is likely to lead to more reliable estimation (Hahn and Hausman, 2002).

In constructing our theoretical MFN tariff cut, we exploit the objective of the Uruguay Round which was to obtain an overall reduction target of thirty per cent in non-agricultural products (Baldwin 2009). We thus construct a hypothetical CUT_g variable, denoted by $HCUT_g$, where the base tariff rate of each industrial product is reduced by thirty per cent (thus $HCUT_g \equiv 0.7$ x Tokyo MFN).[24] $HCUT_g$ is a valid instrument for CUT_g insofar as it is correlated with CUT_g and does not influence $SEVEN_g$ or $\Pr\{IDA_{g,p}=1\}$ directly: the correlation between the actual and the hypothetical CUT_g variables is equal to 0.67; the latter hypothesis is warranted because this 30 per cent reduction was across the board, i.e. it was meant to affect all manufacturing goods *and all countries*. We then run (2) and (3) by TSLS (two-stage least squares) and by two-step IV Probit with the actual CUT being instrumented by its hypothetical value. The second-stage results are reported in Table 8 in the (IV) columns. Note that the TSLS and IV-Probit coefficients are not readily comparable to those of the logit regressions so far. Therefore, we also run (2) and (3) by OLS and Probit as a benchmark and the results are reported in Table 8, in the (Non Instrumented) columns.[25] Two results are noteworthy. First, the coefficients of the variable of interest, CUT_g, are positive, which is in line with our findings so far in all specifications. Second, except for the IV Probit estimation results obtained in the good-specific case and reported in panel (a), the IV and Non-IV coefficients are quantitatively similar, suggesting that the endogeneity bias is not severe in the first place. More importantly, we conclude that the effect of bold Uruguay Round tariff cuts on the likelihood of posterior (preferential) trade liberalization is *causal*.

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[23] "Basic manufacturing" is a mixed-bag category: it includes beverages, spirits and vinegar as well as iron, steel and other base metals *inter alia*.

[24] Note that the correlation between *HCUT* and *MFN* is almost zero. That is to say, and perhaps surprisingly, the Tokyo and Uruguay *MFN* rates are almost uncorrelated.

[25] The estimated coefficients could be compared to those obtained by Logit estimation by using a simple rule of thumb as suggested by Cameron and Trivedi (2005), that is, $\hat{\beta}_{Logit} \cong 4\hat{\beta}_{OLS}$ and $\hat{\beta}_{Logit} \cong 1.6\hat{\beta}_{Probit}$. Applying that rule of thumb to our estimates we find that the correspondence is small enough to relate the IV findings to the Logit ones.

Table 8: IV Estimation

Panel (a): *Good Specific*

	Dependant variable: *SEVEN* (Probability that tariff line *g* is granted IDA to United States market to all 7 partners)			
	Linear Probability		Probit	
	(IV)	(Not Instrumented)	(IV)	(Not Instrumented)
Tariff CUT	0.0307[a]	0.0362[a]	0.0379[a]	0.159[a]
(Tokyo minus Uruguay)	(0.00202)	(0.00147)	(0.010)	(0.00695)
MFN	-0.0162[a]	-0.0168[a]	-0.2184[a]	-0.253[a]
Tariff rate	(0.00160)	(0.00147)	(0.0105)	(0.0147)
DIFF0 (no Uruguay	0.0982[a]	0.129[a]	0.0420	0.825[a]
Round cut)	(0.0181)	(0.0158)	(0.0844)	(0.0622)
Observations	6822	6822	6822	6822
R^2	0.299	0.301	-	-
LL	-	-	-	-26952

Notes: **Robust Standard errors** in parentheses. [a] $p < 0.01$, [b] $p < 0.05$.
All regressions include sector dummies and the controls 'Share imports from RTA partners' and "Share imports from NAFTA partners".

Panel (b): *Country-Good Specific*

	Dependant variable: Pr{IDA = 1} (Probability that tariff line *g* is granted IDA to United States market to partner p)			
	Linear Probability		Probit	
	(IV)	(Not Instrumented)	(IV)	(Not Instrumented)
Tariff CUT	0.0185[a]	0.0204[a]	0.0683[a]	0.0651[a]
(Tokyo minus Uruguay)	(0.00122)	(0.000926)	(0.0026)	(0.00324)
MFN	-0.0125[a]	-0.0127[a]	-0.0484[a]	-0.0439[a]
Tariff rate	(0.000957)	(0.000894)	(0.0021)	(0.00229)
DIFF0 (no Uruguay	0.0845[a]	0.0952[a]	0.320[a]	0.283[a]
Round cut)	(0.0109)	(0.00949)	(0.0345)	(0.0298)
Observations	51814	51814	51814	51814
R^2	0.097	0.097	-	-
LL	-	-	-	-3132

Notes: **Robust Standard errors** in parentheses (clustered by tariff line). [a] $p < 0.01$, [b] $p < 0.05$.
All regressions include sector and partner dummies as well as the additional controls of Table 3. We carried out the usual series of tests (Hansen-J and Kleibergen-Paap rk ML statistics) that assess the validity of the instrumental variables and none of these tests indicates a problem at the usual confidence levels. The Kleibergen-Paap rk Wald F-statistics for weak instruments in the presence of clustered standard errors indicate with 95% confidence a maximum TSLS size well below 10%, implying that our instrument is strong. The values of the test are 314,96 and 251,12 for column (1) of panels (a) and (b) respectively. The 10% maximal IV size test critical value is 19.93. With IV-Probit estimation, the set of available tests to assess the validity of instrumental variables remains limited. First-step statistics point to a very strong predictive power of our excluded instrument. Exogeneity is strongly rejected according to the standard Wald test in the IV probit specifications reported in panel (a). In panel (b) specification, it is only rejected at the 10-percent level.

6. Sensitivity analysis

In this section we subject our results to a variety of robustness checks. We start by running alternative specifications to (2); we further test the relevance of the "money-left-on-the-table hypothesis"; we account for the high incidence of zero imports, for the influence of stalling multilateral trade talks and for the length of the implementation period of tariff cuts. As we shall see, these essentially establish that the emulator effect is non-linear.

6.1. Evidence at the good level: Alternative Logit

In our quest for the effects of CUT_g on the IDA status of goods, specification (2) with $SEVEN_g$ as the dependent variable is quite conservative insofar as it lumps together goods that are excluded from all RTAs altogether with goods that are granted IDA status in all but one RTA. Other categorizations of the data are possible.

Our first robustness check is to run a logit that is the mirror image of (2):

$$\Pr\left\{ONE_g = 1\right\} = \Lambda\left(f_{G(g)} + \beta_1 CUT_g + \beta_2 MFN_g^{Uruguay} + \mathbf{X_{g,p}\beta}\right), \qquad (4)$$

where ONE_g takes value one if the specific good gets IDA status into the United States market in *at least one* RTA and zero otherwise (i.e. $ONE_g \equiv 1 - \mathrm{I}_0\left\{\# p : PT_{g,p}^{impl} = 0\right\}$, where $\mathrm{I}_0\{\cdot\}$ denotes an indicator function that takes value 1 if its component is equal to zero and value 0 otherwise).

We report the results in Table A1, which is symmetric to Table 2 (same set of controls, same estimator). Qualitatively, all the findings are similar to those of Table 2. Quantitatively, the positive effect of CUT_g and the negative effects of $MFN_g^{Uruguay}$, $DIFF0_g$ and $SNAFTA_g$ in (4) are smaller in absolute value than in (2). The odds ratio corresponding to the coefficient of interest β_1 is ranges from 1.13 in the baseline specification to 1.17 with the $DIFF0_g$, SM_g and $SNAFTA_g$ controls, implying that an additional one-percentage point multilateral tariff cut is associated with a 13–17 per cent increase in the odds of including the good in the group of IDA goods. Though quite strong, the effect of CUT_g on ONE_g is weaker than its effect on $SEVEN_g$. This suggests that the domestic resistance to preferential trade liberalization is decreasing in the number of IDA statuses being granted at the margin.

6.2. Evidence at the good level: Poisson

A natural alternative to (2) and (4) is to regress the *number* of times good g is being granted IDA status, defined, as $NTL_g \equiv \#\{p : PT_{g,p}^{impl} = 0\}$, on our list of control variables. This alternative measure of the extensive margin of the 'emulator effect' is a count variable, so we run the constant semi-elasticity model (Poisson regression)

$$\mathrm{E}\left[NTL_g \middle| CUT_g, MFN_g^{Uruguay}, \mathbf{X_{g,p}\beta}\right] = \exp\left(f_{G(g)} + \beta_1 CUT_g + \beta_2 MFN_g^{Uruguay} + \mathbf{X_{g,p}\beta}\right) \qquad (5)$$

with one observation per good g.

Table A2 presents our findings. The results are consistent with those of Tables 2 and 8. Columns (1) and (2) report the results of specification (5), respectively excluding and including the sector dummies $f_{G(g)}$, excluding any other control. The coefficients are precisely estimated. In column (2), the Poisson incidence rate ratio (PIRR = exp β_1) is equal to 1.02, implying that an extra one percentage point CUT_g increases the expected number of times that the good in question is granted IDA status by 2 per cent. The PIRR rises to 1.03 when we add the additional controls of columns (3) and (4) (our preferred specification). The effect is not strong quantitatively but it is statistically significant and robust. Again, this suggests that the domestic resistance to preferential trade liberalization is decreasing in the number of IDA statuses being granted at the margin. We confirm this in the immediate sequel.

6.3. Evidence at the good level: Hurdle

We verify that the effect of CUT_g on the extensive margin of preferential trade liberalization as captured by the IDA status is non-linear by implementing a two-stage Hurdle regression. The first step is a logit that is the mirror image of (2),

$$\Pr\{SEVEN_g = 0\} = \Lambda\left(f_{G(g)} + b_1 CUT_g + b_2 MFN_g^{Uruguay} + \mathbf{X_{g,p}b}\right), \tag{6}$$

and the second step is the conditional Poisson regression:

$$\mathrm{E}\left[7 - NTL_g \middle| SEVEN_g = 0;\cdot\right] = \exp\left(f_{G(g)} + c_1 CUT_g + c_2 MFN_g^{Uruguay} + \mathbf{X_{g,p}c}\right). \tag{7}$$

For instance, b_1 informs us about the extent to which one extra percentage point of CUT_g for good g is associated with a *reduction* of the likelihood of that good of being granted IDA status to all seven partners and, failing this, c_1 says how this extra percentage point cut reduces the likelihood of good g being included in one extra RTA. In line of our previous findings, we expect b_1 to be negative (and b_2 to be positive).

The results of the first step (6) are reported in Table A3, panel (a). As expected, the exponentiated coefficients are the mirror image of those of Table 2 (the values of $\left|\beta_j - 1\right|$ in tables 2 and 5 are comparable for all j = 1,2,…) and thus require no further discussion. Likewise, the results for the second step (7) – reported in Table A3, panel (b) – are comparable to those of (5). They also confirm our priors, in line with our earlier finding, that most of the emulator effect is captured by $SEVEN_g$. The economic significance of the coefficients is small (though all coefficients are statistically significant at the one per cent level with the exception of SM_g, which is significant at the five per cent level).

Taken together, the findings of Tables 2 and Appendix tables A1 to A3 imply that the manifestation of the emulator effect is non-linear and most strongly felt between granting 7 IDA statuses and 6 IDA statuses or fewer.

6.4. Interaction between CUT and MFN

In order to further distinguish between the "money-left-on-the-table hypothesis" and the emulator effect, we now interact CUT_g with $MFN_g^{Uruguay}$ in all the above specifications. The motivation for doing this is the following. If the dynamic complementarity between past (multilateral) cuts and current (preferential) liberalization that we have uncovered so far hid a *static* substitution between multilateralism and regionalism, then we should expect the effect of CUT_g on IDA treatment to be stronger where there is more room for manoeuvre, that is, where $MFN_g^{Uruguay}$ tariff rates are larger. This is not what we find, however.

Table A4 reports the results, two of which are noteworthy. First, the coefficient of $MFN_g^{Uruguay} * CUT_g$ is strongly negative (its odds ratio is lower than unity), which rebukes the hypothesis whereby multilateralism and regionalism are substitutes. Second, the addition of this interaction term increases the coefficient on CUT_g and reduces the coefficient on $MFN_g^{Uruguay}$. Results obtained with the Hurdle estimation strategy largely confirm these patterns.

We interpret all these results as adding extra pieces of evidence if favour of the emulator hypothesis and against the alternative "money-left-on-the-table hypothesis".

6.5. Zero Imports

The share of zero imports is a prominent feature in the dataset (see Table 1). We exploit this feature of the data as an additional test for the emulator hypothesis in two ways that are in line with political economy arguments presented above. First, the emulator effect is economically meaningful only if it holds for goods for which import competition bites. We thus introduce a dummy ZM_g to signal zero imports (ZM_g = 1 if imports are zero and ZM_g = 0 if imports are strictly positive) and we expect the coefficient of the interaction term $(1-ZM_g)*CUT_g$ to be positive. Second, we expect the coefficient of ZM_g*CUT_g to be larger than the former coefficient because zero imports are synonymous with no competitive threat by foreign firms in the import competing segment. In mercantilist terms, granting preferential access to such imports is a cheap "concession" to make.

The results are reported in Table A5. In line with our priors, the coefficient of CUT_g remains positive and the coefficient of the interaction term ZM_g*CUT_g is larger than the coefficient on $(1-ZM_g)*CUT_g$. That is to say, the emulator effect is stronger for goods where imports are zero than for goods where imports are strictly positive. Import liberalization is thus likely to take place more systematically for products not exposed to competition from the partner country exporters.

6.6. Implementation period

Tariff cuts negotiated during the Uruguay Round have not been implemented uniformly either across countries or across products. Only 40 per cent of United States tariff cuts for industrial products have been implemented the year after the end of the Round while 30 per cent of tariff cuts took 10 years to be implemented.[26] Differences in implementation periods are likely to add to the non-linearity of the emulator effect. They also provide a different metric to capture the boldness of multilateral trade liberalization. We thus control for the implementation period of the MFN tariff cuts using two alternative approaches. The first one consists in adding a control variable, *(Implementation period)$_g$*, defined as the number of years taken for implementing the

[26] The proportions are different for agricultural products: 30 per cent of the tariff cuts have been implemented within the first year following the completion of the Uruguay round it took six years to implement the remaining 70 per cent of tariff cuts.

cuts. The second approach combines the two measures of the boldness of multilateral liberalization, CUT_g and (*Implementation period)*$_g$, into a new one, (*Speed of CUT*)$_g$, defined as the ratio of the two. The results, reported in Table A6, are all consistent with the emulator hypothesis: the coefficient of (*Implementation period)*$_g$ in the first two columns is negative (the odd ratio is smaller than unity), implying that a good for which the tariff cut takes more time to be effective is less likely to be granted IDA status later on; note that the coefficient of our initial measure of multilateral liberalization, CUT_g, remains statistically positive and of the same order of magnitude. Consistently, the coefficient of the *(Speed of CUT)*$_g$ variable indicates that the emulator effect increases with the speed of the implementation of the MFN tariff cut.[27]

7. Summary and concluding remarks

This paper investigates the empirical relationship between cuts in MFN bound rates negotiated during the Uruguay Round of the GATT (1986–1994) and the depth and breadth of Preferential Trade Agreements signed in the aftermath of its completion. Our empirical investigation focuses on the United States using official tariff line level data. To the best of our knowledge, our paper is unique in looking at the causal relationship from multilateralism to regionalism. The existing empirical literature is looks exclusively at the relationship running the other way. This line of research is motivated by the view expressed in numerous theoretical contributions that regionalism may have a “stumbling block” effect on multilateral trade liberalization (Bhagwati, 1991). If the stumbling block hypothesis is correct, then the proliferation of RTAs involving at least one WTO member is guilty of slowing down and threatening the Doha Round of negotiations at the GATT/WTO.

The main findings of the paper are that (a) the imports of goods that the United States liberalizes swiftly the most frequently on a preferential basis are also the goods for which it granted the largest MFN tariff reductions during the Uruguay Round; (b) this effect is robust qualitatively but varies across the types of goods, being stronger for intermediate and capital goods; (c) it holds only for goods that have no alternative import restrictions in the form of Non Tariff Measures; (d) it is weaker for goods with prohibitively costly Rules of Origin.

We interpret these findings as evidence that multilateral tariff “concessions” are dynamic complements to preferential treatment of RTA partners. Such dynamic complementarities between sequential Rounds of trade liberalization are consistent with the “Juggernaut” theory of trade liberalization. This theory stresses the role of domestic sluggish adjustments to account for the systematic, monotonically decreasing trade barriers of the modern trading system.

Crossing our results with those of Limão (2006) – who finds that preferential trade liberalization prior to the completion of the Uruguay Round acted as a stumbling block to multilateralism in the United States case – we may thus conclude that the difficulties encountered by the Doha Round might in part be the indirect result of the success of the Uruguay Round.

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[27] We also ran a regression with (*Speed of RTA liberalization*)$_{g,p}$ as the dependant variable (and defined in a similar way as *(Speed of CUT)*$_g$) on the controls of columns 3 and 4 of Table A6. The results (not reported) are also strongly consistent with the emulator hypothesis.

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Appendix tables

Table A1: LOGIT "One'"

	Dependant variable: ONE (Probability that tariff line is granted IDA to United States market to at least one partner)				
	(1)	(2)	(3)	(4)	(5)
Tariff CUT (Tokyo minus Uruguay)	1.054[a] (0.0124)	1.133[a] (0.0179)	1.178[a] (0.0226)	1.178[a] (0.0227)	1.169[a] (0.0234)
MFN tariff level	0.976[a] (0.00644)	0.954[a] (0.00543)	0.946[a] (0.00581)	0.947[a] (0.00581)	0.948[a] (0.00590)
DIFF0 (no Uruguay Round cut)			2.275[a] (0.378)	2.279[a] (0.379)	2.217[a] (0.371)
Share imports From RTA partners				1.037 (0.0671)	1.031 (0.0675)
Share imports from NAFTA partners					0.995[b] (0.00202)
Sector FE	No	Yes	Yes	Yes	Yes
Observations	5756	5756	5756	5756	5756
Pseudo R^2	0.019	0.132	0.140	0.140	0.141
LL	-1662.1	-1355.6	-1343.0	-1342.8	-1340.6

Notes: **Coefficients**: Exponentiated (odds ratios); **Robust standard errors** in parentheses.
[a] $p < 0.01$, [b] $p < 0.05$.

Table A2: POISSON regressions

	Dependant variable: *NTL* (Number of times that tariff line *g* is granted IDA to United States market)				
	(1)	**(2)**	**(3)**	**(4)**	**(5)**
Tariff CUT (Tokyo minus Uruguay)	1.015[a] (0.000949)	1.021[a] (0.00102)	1.028[a] (0.00129)	1.028[a] (0.00129)	1.026[a] (0.00133)
MFN tariff rate	0.971[a] (0.00122)	0.975[a] (0.00134)	0.974[a] (0.00137)	0.974[a] (0.00137)	0.974[a] (0.00137)
DIFF0 (no Uruguay Round cut)			1.152[a] (0.0152)	1.153[a] (0.0152)	1.150[a] (0.0152)
Share imports from RTA partners				1.011[b] (0.00500)	1.010[b] (0.00494)
Share imports from NAFTA partners					0.999[a] (0.000201)
Sector FE	No	Yes	Yes	Yes	Yes
Observations	7419	7419	7419	7419	7419
Pseudo R^2	0.029	0.045	0.048	0.048	0.048
LL	-15775.5	-15505.6	-15469.7	-15468.0	-15459.7

Notes: **Coefficients**: Exponentiated (Poisson Incidence Rate Ratios, or PIRR); **Robust standard errors** in parentheses. [a] $p < 0.01$, [b] $p < 0.05$.

Table A3: HURDLE regressions

Panel (a) *Logit*

	Dependant variable: 1- *SEVEN* (Probability that tariff line *g* is not granted IDA to United States market to all 7 partners)				
	(1)	(2)	(3)	(4)	(5)
Tariff CUT	0.877[a]	0.815[a]	0.752[a]	0.751[a]	0.761[a]
(Tokyo minus Uruguay)	(0.00636)	(0.00727)	(0.00892)	(0.00894)	(0.00924)
MFN	1.496[a]	1.522[a]	1.635[a]	1.635[a]	1.637[a]
tariff rate	(0.0286)	(0.0382)	(0.0466)	(0.0467)	(0.0469)
DIFF0 (no Uruguay			0.229[a]	0.228[a]	0.235[a]
Round cut)			(0.0240)	(0.0240)	(0.0247)
Share imports				0.981	0.990
from RTA partners				(0.0338)	(0.0334)
Share imports					1.008[a]
from NAFTA partners					(0.00165)
Observations	7419	7419	7419	7419	7419
LL	-12392.7	-11372.1	-11254.2	-11253.8	-11238.9

Notes: **Coefficients**: Exponentiated (odds ratios); **Robust standard errors** in parentheses. [a] $p < 0.01$, [b] $p < 0.05$.

Panel (b) *Conditional Poisson*

	Dependant variable: 7 – *NTL*, conditional on *NTL* < 7 (Number of times that tariff line *g* is *not* granted IDA to United States market)				
	(1)	(2)	(3)	(4)	(5)
Tariff CUT	0.995	0.982[a]	0.977[a]	0.977[a]	0.977[a]
(Tokyo minus Uruguay)	(0.00248)	(0.00281)	(0.00311)	(0.00312)	(0.00315)
MFN	1.004[a]	1.011[a]	1.012[a]	1.012[a]	1.012[a]
tariff rate	(0.000331)	(0.00144)	(0.00150)	(0.00150)	(0.00151)
DIFF0 (no Uruguay			0.871[a]	0.871[a]	0.873[a]
Round cut)			(0.0242)	(0.0242)	(0.0244)
Share imports				0.993	0.994
from RTA partners				(0.00762)	(0.00765)
Share imports					1.001
from NAFTA partners					(0.000344)
Sector FE	No	Yes	Yes	Yes	Yes
Observations	7419	7419	7419	7419	7419
LL	-12392.7	-11372.1	-11254.2	-11253.8	-11238.9

Notes: **Coefficients**: Exponentiated (Poisson Incidence Rate Ratios, or PIRR); **Robust standard errors** in parentheses. [a] $p < 0.01$, [b] $p < 0.05$.

Table A4: Interacting CUT and MFN

	Specification:					
	LOGIT Seven	**p-g LOGIT**	**LOGIT One**	**POISSON**	**HURDLE I [*] (logit)**	**HURDLE II [*] (trunc. poisson)**
Tariff CUT (To. minus Ur.)	1.443[a] (0.0419)	1.172[a] (0.0115)	1.187[a] (0.0269)	1.033[a] (0.00209)	0.693[a] (0.0201)	0.983[a] (0.00339)
MFN tariff rate	0.669[a] (0.0255)	0.953[a] (0.00414)	0.970[b] (0.0118)	0.979[a] (0.00172)	1.494[a] (0.0568)	1.020[a] (0.00186)
MFN*Tariff CUT	0.979[a] (0.00541)	0.993[a] (0.000880)	0.998 (0.00117)	0.999[a] (0.000260)	1.021[a] (0.00564)	0.999[a] (0.0000832)
DIFF0 (no Uruguay Round cut)	3.891[a] (0.406)	1.567[a] (0.0783)	2.126[a] (0.348)	1.145[a] (0.0151)	0.257[a] (0.0268)	0.864[a] (0.0242)
Share imports from RTA partners	1.012 (0.0331)	1.039[a] (0.00818)	1.033 (0.0674)	1.010[b] (0.00489)	0.988 (0.0323)	0.993 (0.00771)
Sector FE	Yes	Yes	Yes	Yes	Yes	Yes
Partner FE	N.A.	Yes	N.A.	N.A.	N.A.	N.A.
Observations	6822	51814	5756	7419	7419	7419
Pseudo R^2	0.324	0.089	0.143	.		-
LL	-3072.3	-27870.2	-1338.2	-15450.5		-11215.5

Notes: **Coefficients**: Exponentiated ; **Robust standard errors** in parentheses. [a] $p < 0.01$, [b] $p < 0.05$.
All regressions include sector FE and "Share imports from NAFTA partners". MFN*CUT represents the interaction between the variable MFN tariff rate and the variable Tariff CUT. [*] Columns (5) and (6) report results from Hurdle estimation and should then be considered jointly. Column (5) shows results obtained in the first step (a logit estimation). Column (6) shows results obtained in the second step (a truncated Poisson estimation). Note that we expect the coefficients of the Hurdle regressions to be the opposite of the coefficients in Col. (1) to (4) because the Hurdle regressions are specified as the mirror image of the logit and Poisson regressions.

Table A5: Zero Imports

	Specification:	
	LOGIT Seven	**p-g LOGIT**
(1-ZM)* Tariff CUT	1.211 [a]	1.068 [a]
(To. minus Ur.)	(0.0181)	(0.0188)
ZM * Tariff CUT	1.363 [a]	1.123 [a]
	(0.0219)	(0.00698)
MFN	0.599[a]	0.926 [a]
tariff rate	(0.0175)	(0.00387)
ZM	1.047	0.767 [a]
(Zero Imports)	(0.13)	(0.0397)
Controls[c]	Yes	Yes
Sector FE	Yes	Yes
Partner FE	No	Yes
Observations	6822	51814
Pseudo R^2	0.333	0.121
LL	-3028.1	-26894.4

Notes: **Coefficients**: Exponentiated; **Robust standard errors** in parentheses. [a] $p < 0.01$, [b] $p < 0.05$; [c] all regressions include "DIFF0" and "Share imports from NAFTA partners" as controls. ZM is a dummy variable taking the value one if imports at the tariff line level are zeros. ZM*CUT represents the interaction between the ZM dummy and the variable Tariff CUT.

Table A6: Implementation Period

	Specification:			
	LOGIT Seven	**p-g LOGIT**	**LOGIT Seven**	**p-g LOGIT**
Tariff CUT	1.211 [a]	1.096 [a]		
(To. minus Ur.)	(0.0157)	(0.00749)		
MFN	0.62 [a]	0.929 [a]	0.598 [a]	0.925 [a]
tariff rate	(0.179)	(0.00386)	(0.0185)	(0.00391)
Implementation period	0.773[a]	0.935 [a]		
(Number of years)	(0.0136)	(0.00721)		
CUT speed			1.361 [a]	1.181 [a]
(CUT over Impl. period)			(0.0154)	(0.00804)
Controls[c]	Yes	Yes	Yes	Yes
Sector FE	Yes	Yes	Yes	Yes
Partner FE	No	Yes	No	Yes
Observations	6822	51814	6822	51814
Pseudo R^2	0.343	0.125	0.362	0.144
LL	-3028.1	-26894.4	-2897.6	-26190.8

Notes. **Coefficients**: Exponentiated; **Robust standard errors** in parentheses. [a] $p < 0.01$, [b] $p < 0.05$; [c] all regressions include "DIFF0" and "Share imports from NAFTA partners" as controls. "Implementation period" is defined as the number of years taken to implement the Tariff CUT. "CUT speed" is the Tariff CUT divided by the number of years of the "Implementation period".

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