THE 'EMULATOR EFFECT' OF THE URUGUAY ROUND ON US REGIONALISM *

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This version: January 2010

ABSTRACT

Using a detailed data set at the tariff line level, we find an *emulator effect* of multilateralism on subsequent regional trade agreements involving the US. We exploit the variation in the frequency with which the US has granted immediate duty free access (IDA) to its Free Trade Area partners across tariff lines. A key finding is that the US has granted IDA status especially on goods for which it had cut the multilateral MFN tariff during the Uruguay round the most. Thus, the Uruguay Round (multilateral) 'concessions' have emulated subsequent (preferential) trade liberalisation. We conclude from this that past liberalisation sows the seeds of future liberalisation and that multilateral and preferential trade agreements are dynamic complements.

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

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* We are especially grateful to Alessandro Nicita, Marcelo Olarreaga and Emanuel Ornelas for detailed feedback and comments as well as to Richard Baldwin, Emily Blanchard, Caroline Freund, Jaya Krishnakumar, Ugo Panizza and Esteban Rossi-Hansberg and participants in seminars at the WTO, Glasgow University and the Workshop on "The New Political Economy of Trade" at the European University Institute for useful discussions and suggestions. Robert-Nicoud is also a Peter B. Kenen Visiting Scholar at Princeton University and gratefully acknowledges the hospitality of its International Economics Section.

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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 GATT. The US is no exception. These agreements involving the US vary in scope – the number of goods included in the agreement varies across agreements – and breadth – the US tariff on some goods goes to zero immediately upon implementing the agreement but the imports of many other are fully liberalised only gradually. In this paper, we shed light on the causes of these cross-good variations and show that they are best though as the continuation of a process that includes multilateral liberalisations. Specifically, we find that the imports of goods that the US liberalises swiftly the most frequently on a preferential basis are also the goods for which it granted the boldest tariff cuts during the Uruguay Round. This finding is robust to a variety of specifications. The quantitative effect is also quite large. We interpret these findings as evidence that past multilateral (or non-discriminatory) trade agreements are a dynamic complement, or emulator, to consecutives regional (or preferential) agreements.[1]

Our results matter for three reasons at least. First, one striking feature of the current world trading 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 but 421 PTAs had been notified to the GATT/WTO and 230 of them were in force as of December 2008. What is driving this growing proliferation of PTAs? In this paper, we make ours Wilfred Ethier's assertion that 'regionalism is an endogenous response to the multilateral trading system (Ethier 1998: 1216)'. Our research question can thus be summarized as asking the question "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 rare, perhaps unique, exceptions. Our paper studies this question from an empirical perspective, focusing on the United States.

[1] We also find some interesting and systematic deviation from this pattern, to which we return below.

Second, our paper contributes to the large research agenda that asks whether 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, Grossman and Helpman 1995, Levy 1997, Bagwell and Staiger 1998, Krishna 1998, McLaren 2002).[2] Limão (2006) finds empirical support for the stumbling block hypothesis for the US case; Estevadeordal, Freund and Ornelas (2008) find a 'building block' effect in a sample of ten Latin American countries; Freund and Ornelas (2009) provide an excellent review of this abundant literature.[3] We complement it by asking the causality question in the opposite direction, as Ethier (1998) and Freund (2000a), but from an empirical angle.

Third, with few exceptions, existing theoretical studies on the complements-vs-substitutes issue address this question using either one-shot games or dynamic games that exhibit stable steady-state equilibrium tariffs. Therefore, these models are ill-suited to address the stylised fact illustrated in Figure 1: US tariffs, both preferential and multilateral, keep falling over time.[4] Consequently, in addressing the question as to whether there exists any (causal) link between the two series, we ask whether multilateral tariff *cuts* are associated with more preferential tariff *cuts*: in noticeable departure from the existing literature, we don't run our regressions in level. Our regression results reveal that the US' preferential tariff cuts are a dynamic complement to its multilateral cuts. This provides (to the best of our knowledge: original) evidence in favour of the 'Juggernaut theory' of trade liberalisation, whereby current

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[2] Also not one month elapses without the economic press worrying about this issue. Editorial lines predominantly echoe 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).

[3] Limao 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.

[4] In Figure 1, the 'effectively applied tariff' series is a simple average of MFN and preferential tariffs across tariff lines. For reasons that will become clear below, most of the preferential tariffs are zero.

liberalisation, by eroding protectionist forces and hence resistance to future trade reforms, is sowing the seeds of future liberalisation (Baldwin 1994, Maggi and Rodrìguez-Clare 2007, Baldwin and Robert-Nicoud 2007).

FIGURE 1 ABOUT HERE

The explanatory variable that is the focus of our interest is the multilateral tariff cut that we label 'CUT'. Our measure for CUT is the difference between the Tokyo Round and the Uruguay Round MFN tariffs. We want to relate this to a measure of the intensity of preferential trade liberalisation subsequent to the completion and much of the implementation of the Uruguay Round. In the US, resistance to preferential trade liberalisation (conditional on it taking place) cannot take the form of positive preferential tariffs for institutional reasons, as we explain in the data section of the paper. It can only take the form of delayed liberalisation. Therefore, our measure of the intensity of post-Uruguay Round preferential trade liberalisation (or 'PTL') for each good is the frequency at which the US grants immediate duty-free access to its market to its FTA trading partners.[5] For instance, the US grants immediate duty free access to all seven partners in our sample for 35% of the goods (2,627 goods out of 7,419), to none for 6% of the goods and to between one and all but one partners for 59% of the goods (See Figure 2).

FIGURE 2 ABOUT HERE

We find that an increase in the tariff CUT of one percentage point is associated with an increase in the probability of the US granting immediate duty-free access to its market to all trade partners by about twenty-five percent at the sample mean. Given that the standard error for CUT in the sample is 4.34 percentage points, this is a large effect.

An alternative interpretation for our results is also possible: the dynamic complementarities between the Uruguay Round and the preferential tariff cuts might just reflect dynamic complementarities between past and current liberalisations – regardless of the level (preferential or multilateral) at which they are conducted. Perhaps the US grants these

[5] A free trade area, or 'FTA', is a special kind of PTL: its preferential tariffs are zero.

‘concessions’ at the preferential level because the Doha Round of multilateral trade talks is currently stalling. This latter hypothesis, which we label ‘the money left on the table hypothesis’, is quite popular among many pundits or in the press (The Economist is a particularly ardent propagator of this view of the world). Note that the two explanations are not mutually exclusive. We control for this hypothesis in two ways. First, we introduce the Uruguay Round MFN tariff rate as a control in all our regressions. The estimated coefficient is negative, implying that the US disproportionately grants duty free access to its market on a preferential basis for goods that have a *low* MFN tariff rate already. This rejects the money left on the table hypothesis. Second, it turns out that the US did not cut MFN tariffs at the Uruguay Round on about 22% of goods in our sample; so, we introduce a dummy variable for such goods as an additional control, recognizing that these might be different for some reason. The estimated coefficient of this control is statistically significant and positive, implying that the Uruguay Round and the ensuing preferential tariff cuts are dynamic *substitutes* for these goods. The presence of this control among the independent variables also increases the estimated coefficient of CUT, which reinforces our emulator finding for the remaining 78% of tariff lines.

Several explanations may explain this emulator effect but not all of them imply that past (multilateral) trade liberalisation is a force behind current (preferential) trade liberalisation. We pursue several routes in order to interpret the positive correlation between multilateral tariff cuts and preferential liberalisation in causal way. As we explain in Section 4, we rely on the timing of events to rule out *reverse causation*. Dealing with the presence of *omitted variables* like political economics forces is more involved (Baldwin and Seghezza 2008, Estevadeordal et al. 2008). We start by introducing 2-digit sector dummies to control for characteristics that are common across goods of the same industry. Our results show that this improves the identification of the emulator hypothesis. We then estimate a different CUT coefficient for goods that are protected by non-tariff measures (NTM) and/or prohibitively costly rules of origin (RoO). If third factors were to explain the correlation between CUT and preferential trade liberalisation in full, then the conditional CUT coefficients should not systematically differ across goods categories. By contrast, if multilateral tariff cuts cause preferential tariff cuts, then our identifying assumption for the emulator effect is that it be strongest when it maters the most, namely, for goods that have no NTMs or prohibitive RoOs.

The results are consistent with this assumption: there is *no* emulator effect for goods with NTMs; the emulator effect is stronger for goods with prohibitively costly RoOs.

We also use existing theoretical results as an alternative way of identifying the emulator effect. We construct our argument by combining two ingredients. Our first ingredient is dynamic: Maggi and Rodriguez-Clare (2007) and others postulate that past trade liberalisation in a given sector undermines its current resistance to trade openness because trade liberalisation decreases the (quasi) rents associated with the (quasi) fixed factors that fight for protection. Over time, these factors depreciate and with them the resistance to trade liberalisation. Thus, over the years, this logic repeats and the past trade liberalisation feeds current and future liberalisation; once started, like a juggernaut, it keeps rolling. Our second ingredient is static: in the Protection For Sale (PFS) framework due to Grossman and Helpman (1994), Gawande, Krishna and Olarreaga (2009) formalize the idea that downstream sectors *oppose* protection of domestic upstream sectors from which they source. By a symmetric argument, upstream sectors favour protection in the domestic downstream sectors they sell to. Taken together, the PFS and the juggernaut logics imply that the emulator effect is strongest in upstream sectors and weakest in downstream ones. Consistent with this prior, the data reject the alternative hypothesis whereby there should be no differential effect.

Finally, we also experimented with instrumenting for MFN tariff cuts and levels with the corresponding EU tariff cuts and levels. This strategy is not faultless, but EU tariffs were too weakly correlated with their US counterparts to make them valid instruments anyway (this came as a surprise to us). We therefore do not discuss these issues or the results further.

The rest of the paper is organised as follows. Section 2 further discusses work related to ours. Section 3 defines the variables and the data. Section 4 introduces our estimation strategy and displays the baseline empirical results; Section 5 reports various identification strategies of the emulator effect while robustness checks are relegated to Section 6. Section 7 concludes.

2. RELATED LITERATURE

Our findings are consistent with two different arguments put forth in the theoretical literature. The first class of models studies the welfare effects of preferential versus multilateral trade liberalisation and, on the positive side, whether liberalising on a preferential basis first, by changing the status quo ante, undermines multilateralism (see Bhagwatti 1991 and the subsequent literature). Even in this case, though, the models 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 liberalisation context explicitly. Freund (2002b) emphasizes that the same type of logic also entails that the incentives to form an FTA 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 PTAs' (Freund 2002b: 359). Our results vindicate her conclusion. In a PFS setting, Ornelas (2005a) points out that preferential trade liberalisation 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 as the one above asks whether the conditions under which PTAs are enforceable are affected by the multilateral trading environment (Freund 2002b 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. However, if anything, tariffs fall over time and hence this line of explanation misses an important dimension of the real world.

The second strand of the theoretical literature that is related to our empirical work focuses on the dynamic aspect of trade agreements, putting aside the dimension of regionalism versus multilateralism, and seeks to explain why tariffs tend to fall over time. Maggi and Rodriguez-Clare (2007) is a key contribution here. Recognising that some sector-specific factors of production like (human) capital depreciate gradually over time, they stress that the politically optimal tariff is thus also decreasing over time as a result. See also Baldwin (1994), Staiger (1995) and Baldwin and Robert-Nicoud (2007). The central finding for our purpose is that past liberalisation sows the seeds of current liberalisation by eroding the rents from protection. Freund (2000a) and McLaren (2002) also combine dynamic aspects of trade liberalisation with the regionalism versus multilateralism issue but their focus (the hysteretic effects of preferential trade barriers) is different.

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 (Baier and Bergstrand 2004, Egger and Larch 2008). 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, the smaller the difference between their GDPs, the larger their relative factor endowment difference, and the wider the (absolute) difference between their and the rest of the world capital-labor ratios. Building on Baier and Bergstrand (2004), Egger and Larch (2009) find evidence consistent with Baldwin's (1995) Domino theory of regionalism, whereby pre-existing PTAs increase the likelihood that two countries participate in a common PTA. In a separate but no less interesting line of research, Martin, Mayer and Thoenig (2009) find that multilateralism causes peace-motivated regional trade agreements (RTA). 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 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 this threat it poses to bilateral peace. The macro-level empirical evidence in Martin et al. (2009), which is supportive of this argument, complements our micro-level evidence.

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

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.[6] All US MFN tariffs are non-increasing in the post-Uruguay round period. Our key explanatory variable, denoted by *CUT*, is defined as the (non-negative) difference (or tariff 'cut') between the Tokyo and Uruguay MFN rates, i.e. $CUT \equiv MFN^{Tokyo} - MFN^{Uruguay}$. *CUT* is our good-specific measure of the intensity of multilateral trade liberalisation, so we may write CUT_g to be more explicit (with the subscript g denoting the good). The stated aim of the Uruguay Round was to cut tariffs by about 30% but in the end Canada, the EU, Japan, and the US achieved a larger reduction on average (Baldwin 2009).

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 in the HS nomenclature), what we refer to as *goods*. They do not include goods subjected to non-ad valorem duties.[7] 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 liberalisation is a factor towards current preferential trade liberalisation. A measure of the intensity of the preferential trade liberalisation similar in spirit to *CUT* is the *preference margin PM*, defined as the (non-negative) difference between the MFN tariff and the preferential tariff, or $PM_{g,p} \equiv$

[6] See the World Tariff Profiles (2007).

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

$MFN_g^{Uruguay} - PT_{g,p}$, 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 for such goods by definition. This leaves us with our reference sample that includes 7,419 goods. Figure 2 illustrates various features of the sample. No tariff line has been included in fewer than four PTAs 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%) are always set to zero on the date entry into force of the trade agreement. However, we also find some tariff lines (6%) which are set to zero only gradually in all trade agreements.

The UNCTAD-TRAINS database includes MFN applied rates and preferential rates. The informed period is 1996-2008. This exhaustive database covers fifteen PTAs, from which we exclude the PTAs 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 the unilateral PTAs, for the focus of our analysis is not unilateral but preferential trade liberalisation or 'regionalism'. We are thus left with seven PTAs: Jordan (2001), Chile (2004), Singapore (2004), Morocco (2006), Bahrain (2006), Australia (2005), and the Central American-Dominican Republic FTA (2006).[8] 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 PTAs. 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 *CUTg* is 4,22 percentage points.

TABLE 1 ABOUT HERE

[8] 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 Generalised System of Preferences (GSP) for Least Developed Countries (1997), the African Growth and Opportunity Act (2000, 2001, 2002), and the Caribbean Basin Trade Partnership Act (2000). See Romalis (2007).

It turns out that in the US case, each PTA is in fact a free trade agreement (FTA) *de jure*, namely, the tariffs of all included goods all eventually go to zero. In our notation, this implies that PT = 0 at the end of the so-called 'implementation period' (specified in the agreement). By contrast, there is considerable variation in the timing of the implementation of this free trade policy about both goods and partners: overall, 69% of our observations are fully liberalised at the start of the implementation of the FTA, whereas goods that are included in any of the FTAs but that are liberalised only gradually represent 27% of our observations; the rest consists of good-partner pairs that are excluded from the corresponding FTA altogether (fewer than 4% of observations).

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 the multilateral CUT, that is, estimate an equation of the form

$$PM_{g,p} = \alpha + \beta CUT_g + \varepsilon_{g,p} \qquad (1)$$

The 'emulator effect' predicts a positive β, whereas a negative β would be consistent with a *dynamic* version of the 'money left on the table hypothesis'.

The problem with a naïve estimation of (1) is that the US 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 liberalisation as the dependent variable problematic (at the end of the implementation period PT = 0, hence PM boils down to $MFN^{Uruguay}$ by definition). For this reason we exploit instead its extensive margin and the timing of the preferential liberalisation. Our first cut through the data is to set goods that are granted duty free access to the US market immediately upon implementation of each of the seven FTAs in the sample apart from other goods. The idea is that these goods are the easiest to liberalise 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 US market is the maximum (seven). For these reasons, we create a binary variable for each good *g*, *SEVEN*, with $SEVEN_g = 1$ if good *g* is granted IDA status in all the seven FTAs and 0 otherwise (i.e. if the good is granted only *gradual* duty-free access in, or excluded altogether from, at least one FTA); formally, $SEVEN_g \equiv \mathrm{I}_7\left\{\# p : PT_{g,p}^{impl} = 0\right\}$, 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.[9] 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 FTA and 0 otherwise and the count variable NTL_g that counts the number of FTAs in which *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 liberalisation, we define a good- and partner- specific measure of preferential trade liberalisation 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 FTA in question and zero otherwise.

We include the most-favoured-nation (MFN) tariff rate in application MFN_g in all specifications. The motivation for doing this has to do with testing the static version of the 'money left on the table hypothesis', whereby there is more room to include a tariff line in a PTA 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 including it or not does not affect

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

the estimated coefficient β_1. This somewhat surprising feature of the data is also helpful for our identification strategy and we return to it shortly.

4.1. Evidence at the good level: Logit

We start by running the following logit:

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

where $\Lambda(\cdot) \equiv \exp(\cdot)/\left[1+\exp(\cdot)\right]$ is the logistic cumulative distribution function, $f_{G(p)}$ is sector dummy, *CUT* is the reduction in the MFN tariff negotiated over the course of the Uruguay round (in percentage points), MFN_g 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 $\Gamma = \left\{1,...,N_g\right\}$; 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 $G:\text{good} \rightarrow \text{sector}$.

Though we view (2) as a closed form relationship between *PTL* and *CUT*, 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 on the timing of events. We limit our sample to the seven PTAs 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 US 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 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.[10] Note that the absence of correlation between *CUT* and *MFN* is also helpful: it implies that the past determinants of trade liberalisation (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 *PTL* and *CUT* 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, like the political economy determinants of tariffs (e.g. lobbying), as suggested in our theoretical discussion in the introduction, or the determinants of comparative advantage. 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. 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 ABOUT HERE

Table 2 presents the results. We report odds ratios throughout. The odds ratios associated to β_j is defined as $\exp \beta_j$ (j = 1,2,...) and has the meaning that a one extra percentage point in *CUT* raises the probability of granting IDA status to all partners for the good in question by a factor $\exp \beta_1$ *relative to* not including the tariff line or delaying setting this preferential tariff to zero. The two independent variables of interest, *CUT* and *MFN*, are significant beyond the one percent level in all specifications and the results are stable across specifications. The

[10] 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 US Trade Representative.

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* raises the probability that good *g* gets IDA treatment for all US's FTA 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 the *MFN* tariff by one percentage point *decreases* the odds that good *g* gets IDA status by a third (1 – .657 = .343).

In Column (3), we add a good-specific dummy *DIFF*0 that takes value $DIFF0_g = 1$ if the US did not liberalise good *g* during the Uruguay Round (i.e. if $CUT_g = 0$) and zero otherwise.[11] That is, we estimate

$$\Pr\left\{SEVEN_g = 1\right\} = \Lambda\left(f_{G(g)} + \beta_1 CUT_g + \beta_2 MFN_g + \beta_3 DIFF0\right).$$

The fact that goods that were not liberalised during the Uruguay round – because these sectors are better organised 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 one percent level, implying that goods that were not liberalised at the multilateral level were more likely to be liberalised at the preferential level: this is consistent with a dynamic version of the 'money left on the table hypothesis.' Adding this control also raises the odds ratio of *CUT* to 1.33. Thus, the effect of CUT on IDA is 'non-linear': the US grants IDA status more frequently for goods for which the Uruguay Round CUT was zero as well as for those that had a large CUT. The *net effect* is consistent with the emulator hypothesis by our finding reported in Table 2, Col. (2).

The results reported in Columns (4) and (5) show that these qualitative findings are robust to the inclusion of several controls. Column (3) introduces the import share of all seven partners in the US' total imports of good *g*, defined as $SM_g \equiv \sum_p M_{g,p} / M_g$ (where *M* denotes the value of imports), to control for the possibility that the US might be granting IDA access to prominent exporters more easily. The estimated coefficient in Col. (4) is statistically

[11] This is verified for 21.8% of the tariff lines in our reference sample.

insignificant: thus, the US does not seem to discriminate between large and small exporters when granting IDA status.

Column (5) adds *SNAFTA* to the set of controls, with *SNAFTA* being defined as the good-specific import share of NAFTA products, 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 US is less likely to grant IDA status from markets that NAFTA already penetrates widely. This suggests that NAFTA and ensuing FTAs are substitutes, that is, NAFTA worked as a 'stumbling block' to post-Uruguay Round regionalism.

4.2. **Evidence at the good-partner level: logit**

The evidence so far indicates that *CUT* and *MFN* influence the extensive margin of preferential trade liberalisation. 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* and *MFN* influence the likelihood that the US grants IDA status to partner *p*'s exports of good *g* to the US? 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 US market for good *g* and zero otherwise. We then estimate the following logit:

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

where f_p is a partner dummy and the other right-hand side variables are as in (2). 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 is symmetric for each partner. As we shall see, though, the effect of CUT on IDA is non-linear. For this reason, we consider running (3) as a conservative robustness check that provides a lower bound for the emulator effect and the other effects we control for.

TABLE 3 ABOUT HERE

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 Tables 2. The coefficients for *CUT*, *MFN*, *DIFF*0 and *SNAFTA* are still precisely estimated and they have the expected sign.

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 organised sectors. In non-organised sectors, the opposite is true. Estimation of

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

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 *PT* for each of them; this enables us to estimate β_4 via the cross-sectional variation of *SM* 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 *organised* 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 percent increases the odds of the US granting IDA status to *p*'s exports of good *g* by 4 percentage points. In other words, the US grants IDA status disproportionately to important import sources. The estimated coefficient is stable across specifications.

We might also expect the US 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 US' share of exports towards *p*, defined as $SXALL_p \equiv \sum_g X_{g,p} / X$, where *X* denotes exports. 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 US. *SALL*, *SMALL* and *SXALL* are defined at the partner level, so we drop the partner dummy in these regressions. Column (6) reports the results for *SALL* (the results for *SMALL* and *SXALL* are similar so we omit them). The estimated coefficient is statistically indistinguishable from zero, rejecting the hypothesis that the US grants free access to its markets disproportionately to large partners.

TABLE 4 ABOUT HERE

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* and *MFN* 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.

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) 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 and IDA in a causal way.

5.1. Non-tariff measures

We start by controlling for the presence of non tariff measures, or 'NTM', at the tariff line.[12] The idea is that the presence of such non-tariff measures should weaken the effect of *CUT* on

[12] There are 19% of tariff lines with an NTM in our reference sample.

preferential liberalisation: a multilaterally agreed tariff cut is less effective if the imports of that good are impeded by other measures. We thus expect the *CUT* 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 g is NTM-free.

TABLE 5 ABOUT HERE

We first re-run (2), adding the *NTM* dummy and its interaction with *CUT*. 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* coefficient for NTM-free goods is (much) larger than for NTM goods; the difference is significant at any conventional level. The coefficient for *CUT* 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.[13] These findings thus vindicate our emulator hypothesis further. By contrast, the odds ratio of MFN falls, weakening further the ‘money left on the table hypothesis’.

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* 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). We expect the *CUT* coefficient to be larger for RoO-goods than

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

for goods that have irrelevant rules of origin. 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.

TABLE 6 ABOUT HERE

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* coefficient is larger for the goods for which it maters than for goods with an irrelevant preference margin. By contrast, the coefficient and the odds ratio for *MFN* 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 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* 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* and *DIFF*0 as controls). 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 ABOUT HERE

These results also are 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 liberalisation in a given sector undermines its current resistance to trade openness because trade liberalisation decreases the (quasi) rents associated with the (quasi) fixed factors that fight for protection. Over time, these factors depreciate and with them the resistance to trade liberalisation. 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, weakest in downstream sectors and somewhere in between for 'Mixed' goods. With the exception of the ranking of Mixed goods, this is what we find in Table 7. The emulator effect is weakest for Primary, Consumption and Basic manufacturing goods; it is statistically much stronger for Equipment and Intermediate goods.

6. SENSITIVITY ANALYSIS

In this section we subject our results to a variety of robustness checks. We start by running alternative specifications to (2); as we shall see, these 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* on the IDA status of goods, specification (2) with *SEVEN* as the dependent variable is quite conservative insofar as it lumps together goods that are excluded from all FTAs altogether with goods that are granted IDA status in all but one FTA. Other categorizations of the data are possible.

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

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

where *ONE* takes value one if the specific good gets IDA status into the US market in *at least one* FTA 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).

TABLE 8 ABOUT HERE

We report the results in the Table 8, 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* and the negative effects of *MFN*, *DIFF*0 and *SNAFTA* 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 *DIFF*0, *SM* and *SNAFTA* controls, implying that an additional one-percentage point multilateral tariff CUT is associated with a 13 – 17 % increase in the odds of including the good in the group of IDA goods. Though quite strong, the effect of *CUT* on *ONE* is weaker than its effect on *SEVEN*. This suggests that the domestic resistance to preferential trade liberalisation 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)

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

with one observation per good *g*.

TABLE 9 ABOUT HERE

Table 9 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* increases the expected number of times that the good in question is granted IDA status by two percents. 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.

6.3. **Evidence at the good level: Hurdle**

We verify that the effect of *CUT* on the extensive margin of preferential trade liberalisation 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),

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

and the second step is the conditional Poisson regression:

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

For instance, b_1 informs us about the extent to which one extra percentage point of *CUT* 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 FTA. In line of our previous findings, we expect b_1 to be negative (and b_2 to be positive).

TABLE 10 ABOUT HERE

The results of the first step (6) are reported in Table 10, 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) are comparable to those of (5) by the same token.

They also confirm our priors, in line with our earlier finding, that most of the emulator effect is captured by *SEVEN*. The economic significance of the coefficients is small (though all coefficients are statistically significant at the one percent level with the exception of *SM*, which is significant at the five percent level).

Taken together, the findings of Tables 2 and 8 to 10 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**

Finally, we interact *CUT* with *MFN* in all the specifications above. The motivation for this exercise is to further distinguish between the 'money left on the table hypothesis' and the emulator effect. Indeed, it could be said that *current* preferential IDA is a substitute to *current* multilateral liberalisation; put differently, it could be that the dynamic complementarity between past (multilateral) cuts and current (preferential) liberalisation that we have uncovered so far hides a *static* substitution between multilateralism and regionalism. If that 'static substitution' hypothesis was true, then we should expect the effect of *CUT* on *IDA* to be strongest where there is more room for manoeuvre, that is, where MFN tariff rates are largest. In order to verify this empirically, we re-run all the baseline specifications above with an interaction term.

TABLE 11 ABOUT HERE

Table 11 reports the results. The first thing to note is that the coefficient of *MFN * CUT* is strongly negative (its odds ratio is lower than unity), which rebukes this hypothesis. Second, comparing the results of Table 11, Col. (1), (2), (3) and (4) to Col. (5) in Tables 2, 3, 8 and 9, respectively, adding this interaction term increases the coefficient on *CUT* and reduces the coefficient on *MFN*. 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.

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 exclusively looking 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 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 'Doha round' of negotiations at the GATT/WTO. 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.

The main findings of the paper are that (i) the imports of goods that the US liberalises 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, (ii) this effect is robust qualitatively but varies across the types of goods being stronger for goods in upstream sectors and weaker for goods in downstream sectors, (iii) it holds only for goods that have no alternative import restrictions in the form of Non Tariff Measures, (iv) 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 FTA partners. We can state that the past success of multilateralism is at least partly responsible for the current wave of US regionalism.

The dynamic complementarities between sequential rounds of trade liberalisation brought to the fore by our empirical results are consistent with the 'Juggernaut' theory of trade liberalisation. 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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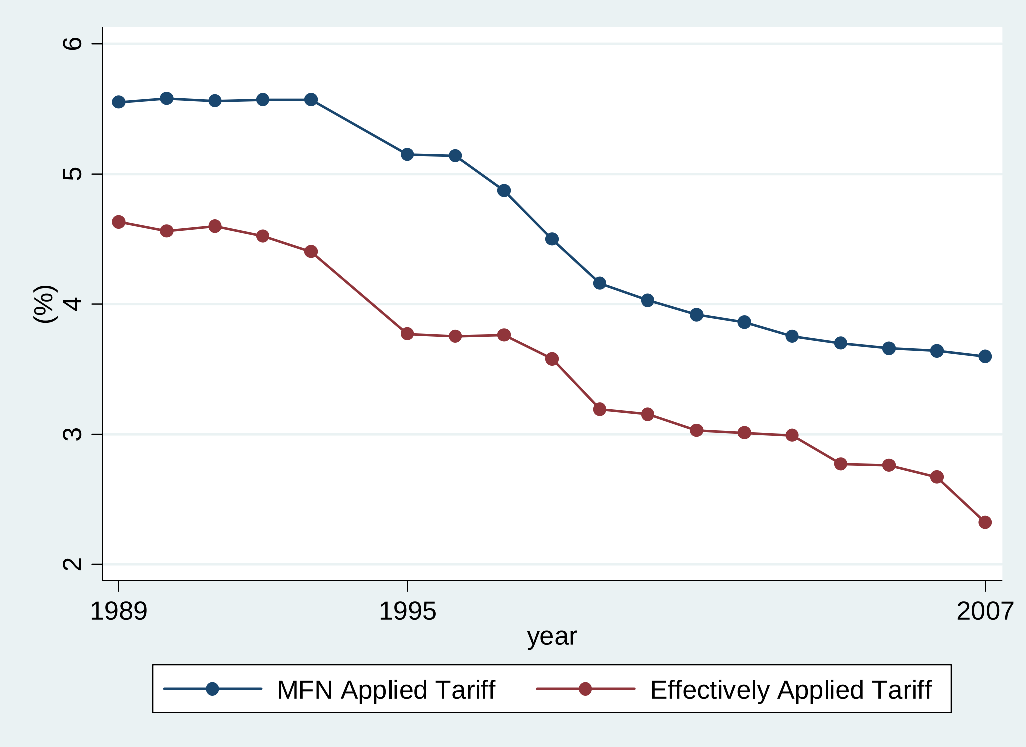
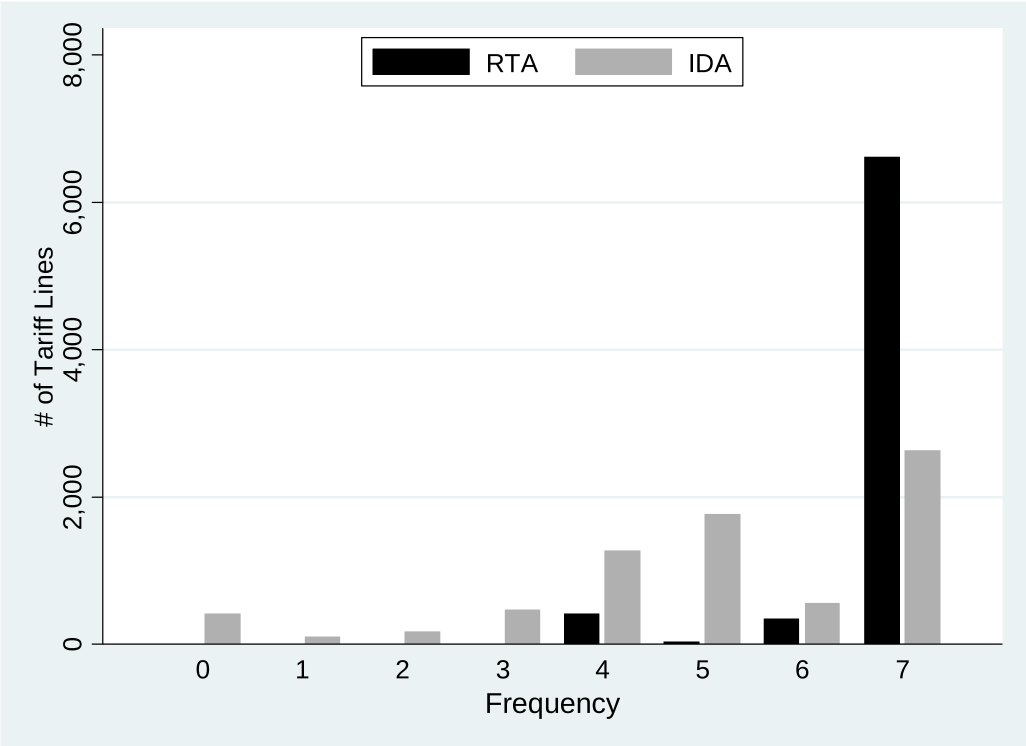
Figure 1: US Tariffs (Simple Means)



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.

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 status (i.e. tariff lines that are liberalized as an RTA enters into force).

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 PTA partners	.45	.23	.51	.005	1.31
Share imports (tariff line) from PTA partners	.21	0	2.63	0	100
Share imports from NAFTA partners	13.15	.73	24.09	0	100
Share exports to FTA partners	.91	.44	.89	.04	2.25

Table 2: LOGIT 'Seven'

	Dependant variable: *SEVEN* (Probability that tariff line *g* is granted IDA to US 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 FTA 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] $p < 0.05$.

Table 3: p-g LOGIT

	Dependant variable: Pr{IDA = 1} (Probability that tariff line g is granted IDA to US market to partner p)					
	(1)	(2)	(3)	(4)	(5)	(6)
Tariff CUT (To. minus Ur.)	1.064[a]	1.099[a]	1.125[a]	1.126[a]	1.115[a]	1.115[a]
	(0.0162)	(0.0197)	(0.0221)	(0.0220)	(0.0212)	(0.0213)
MFN tariff level	0.922[a]	0.931[a]	0.926[a]	0.925[a]	0.930[a]	0.930[a]
	(0.0119)	(0.0125)	(0.0139)	(0.0136)	(0.0134)	(0.0134)
DIFF0 (no U. R. cut)			1.683[a]	1.688[a]	1.623[a]	1.623[a]
			(0.316)	(0.316)	(0.296)	(0.298)
Partner's share of M_g				1.039[a]	1.039[a]	1.041[a]
				(0.0144)	(0.0152)	(0.0128)
Share imports from NAFTA partners					0.996[a]	0.996[a]
					(0.00103)	(0.00103)
SALL: Partner's share of US X+M						0.951
						(0.160)
Sector FE	No	Yes	Yes	Yes	Yes	Yes
Partner FE	No	Yes	Yes	Yes	Yes	No
Obs.	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 at the tariff line) in parentheses. [a] $p < 0.01$, [b] $p < 0.05$.

Table 4: g-Logit on partner-specific sub-sample

	Dependant variable: Pr{IDA = 1} (Probability that tariff line g is granted IDA to US market to partner *p*)						
	(AUS)	(BHR)	(CHL)	(JOR)	(MAR)	(SGP)	(CAFTA)
Tariff CUT (To. minus Ur.)	1.075[b]	1.261[a]	1.120[a]	1.197[a]	1.090[b]	1.175[a]	1.273[a]
	(0.0313)	(0.0411)	(0.0448)	(0.0318)	(0.0369)	(0.0309)	(0.0449)
MFN tariff rate	0.815[a]	0.956[a]	0.895[a]	0.687[a]	0.878[a]	0.640[a]	0.968
	(0.0342)	(0.0142)	(0.0277)	(0.0418)	(0.0282)	(0.0720)	(0.0207)
DIFF0 (no Uruguay Round cut)	2.110[a]	2.440[a]	1.862	2.902[a]	3.097[a]	2.389[b]	2.410[a]
	(0.577)	(0.715)	(0.710)	(0.997)	(1.099)	(0.817)	(0.715)
Share imports from FTA partners	1.017	38.49	0.971[b]	1.083	1.057	0.998	1.019[b]
	(0.0176)	(115.5)	(0.0112)	(0.151)	(0.0351)	(0.00926)	(0.00970)
Share imports from NAFTA partners	0.995[b]	0.995	0.997	0.992[a]	0.997	0.995[b]	0.996
	(0.00210)	(0.00242)	(0.00306)	(0.00222)	(0.00220)	(0.00211)	(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] $p < 0.05$.

Table 5: Non-tariff measures (NTM)

	Dependant variables:			
	SEVEN		Pr{IDA = 1}	
	(1)	(2)	(3)	(4)
Tariff CUT	1.313[a]		1.115[a]	
(To. minus Ur.)	(0.0159)		(0.0212)	
NTM * cutMFN		1.010		0.993
		(0.0375)		(0.00689)
(1-NTM) * cutMFN		1.310[a]		1.140[a]
		(0.0155)		(0.00455)
MFN	0.611[a]	0.603[a]	0.930[a]	0.924[a]
tariff rate	(0.0175)	(0.0173)	(0.0134)	(0.00261)
DIFF0 (no Uruguay	4.253[a]	4.173[a]	1.623[a]	1.700[a]
Round cut)	(0.446)	(0.431)	(0.296)	(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 fixed effects and the controls SM and SNAFTA. 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.

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 fixed effects and the controls SM and SNAFTA. 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 7: LOGIT 'Seven' by type of goods

	Dependant variable: SEVEN (Probability that tariff line g is granted IDA to US market to all 7 partners)					
	Basic-manufacturing	Consumption-goods	Equipment-goods	Intermediate-goods	Mixed-products	Primary
Tariff CUT	1.423[a]	1.181[a]	1.306[a]	1.343[a]	1.404[a]	1.061
(To. minus Ur.)	(0.0433)	(0.0572)	(0.0426)	(0.0404)	(0.0613)	(0.102)
MFN	0.561[a]	0.494[a]	0.838[a]	0.445[a]	0.808[a]	0.201[a]
tariff rate	(0.0301)	(0.0407)	(0.0368)	(0.0344)	(0.0353)	(0.0632)
DIFF0 (no Uruguay	18.62[a]	1.675	3.080[a]	2.493[a]	5.951[a]	2.53e-09
Round cut)	(5.180)	(0.529)	(0.667)	(0.711)	(1.785)	(0.00031)
Share imports	1.018	1.085	1.366	0.676	0.679[b]	1.257[b]
from FTA partners	(0.0716)	(0.103)	(0.260)	(0.229)	(0.125)	(0.121)
Share imports	0.996	0.986[a]	0.976[a]	0.994	0.995	0.990
from NAFTA partners	(0.00352)	(0.00468)	(0.00519)	(0.00434)	(0.00385)	(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] $p < 0.05$.

Table 8: LOGIT 'One'

	Dependant variable: ONE (Probability that tariff line is granted IDA to US 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 FTA 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 9: POISSON regressions

	Dependant variable: *NTL* (Number of times that tariff line *g* is granted IDA to US 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 FTA 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 10 (a): HURDLE regressions

	Panel (a) Logit				
	Dependant variable: 1- *SEVEN* (Probability that tariff line *g* is not granted IDA to US market to all 7 partners)				
	(1)	(2)	(3)	(4)	(5)
Tariff CUT (Tokyo minus Uruguay)	0.877[a] (0.00636)	0.815[a] (0.00727)	0.752[a] (0.00892)	0.751[a] (0.00894)	0.761[a] (0.00924)
MFN tariff rate	1.496[a] (0.0286)	1.522[a] (0.0382)	1.635[a] (0.0466)	1.635[a] (0.0467)	1.637[a] (0.0469)
DIFF0 (no Uruguay Round cut)			0.229[a] (0.0240)	0.228[a] (0.0240)	0.235[a] (0.0247)
Share imports from FTA partners				0.981 (0.0338)	0.990 (0.0334)
Share imports from NAFTA partners					1.008[a] (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] $p < 0.05$.

Table 10 (b): HURDLE regressions (cont.)

Panel (b) Conditional Poisson

	Dependant variable: 7 – *NTL*, conditional on *NTL* < 7 (Number of times that tariff line *g* is *not* granted IDA to US market)				
	(1)	(2)	(3)	(4)	(5)
Tariff CUT (Tokyo minus Uruguay)	0.995	0.982[a]	0.977[a]	0.977[a]	0.977[a]
	(0.00248)	(0.00281)	(0.00311)	(0.00312)	(0.00315)
MFN tariff rate	1.004[a]	1.011[a]	1.012[a]	1.012[a]	1.012[a]
	(0.000331)	(0.00144)	(0.00150)	(0.00150)	(0.00151)
DIFF0 (no Uruguay Round cut)			0.871[a]	0.871[a]	0.873[a]
			(0.0242)	(0.0242)	(0.0244)
Share imports from FTA partners				0.993	0.994
				(0.00762)	(0.00765)
Share imports from NAFTA partners					1.001
					(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 11: 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*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 FTA 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.