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The Clash of Liberalizations: Preferential vs. Multilateral Trade Liberalization in the European Union*

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Abstract

Preferential trade agreements (PTAs) are characterized by liberalization with respect to only a few partners and thus they can potentially clash with, and retard multilateral trade liberalization (MTL). Yet there is almost no systematic evidence on whether the numerous existing PTAs actually affect MTL. We provide a model showing that PTAs hinder MTL unless they entail accession to a customs union with internal transfers. Using product-level tariffs negotiated by the European Union (EU) in the last two multilateral trade rounds we find that several of its PTAs have clashed with its MTL. However, this effect is absent for EU accessions. Moreover, we provide new evidence on the political economy determinants of trade policy in the EU.

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1 Introduction

Over 130 preferential trade agreements (PTAs) were formed between 1995 and 2005—more than in the previous 50 years combined. Nearly all countries are currently members of at least one PTA and at least a third of world trade is carried out under such agreements. Although most economists favor multilateral trade liberalization (MTL), there is no such consensus on the desirability of preferential liberalization. The original concern with PTAs was their ambiguous effect on welfare: positive if the preferential partner is more efficient than the rest of the world but negative otherwise (Viner 1950). During the late 1980s and early 1990s, MTL was stalled while the United States and the European Union pursued PTAs, generating much debate on whether PTAs are a "building block" or a "stumbling block" towards MTL (Bhagwati 1991). This issue was also prominent in the latest multilateral round since several developing countries fear that MTL will erode the preferences provided to them.[1]

An important source of concern with PTAs is that they can hurt non-members. One direct channel by which this occurs is if the PTA members divert their import demand away from the non-members and this effect is large enough to reduce non-members' export prices. There is evidence of trade diversion (Romalis 2007) and also some direct evidence that PTAs do lower export prices for non-members (Chang and Winters 2002). This and other costs to non-members due to discrimination disappear if the preference is fully eroded by MTL. Thus, it is crucial to determine if PTAs hold back MTL and entrench these costs, particularly given that a large share of world trade is not yet covered by any preferences. After much debate there is still no theoretical consensus about and scant empirical evidence of a "clash of liberalizations". We use product-level protection data to estimate the impact of European Union PTAs on its multilateral liberalization in the last multilateral trade round and find evidence of such a clash. The effect is present for all European Union (EU) preferential agreements except those involving full accession where the partner becomes a EU member. Both findings are predictions of our model as we explain below.

There are several compelling reasons for focusing on the EU. First, given that the EU is the world's largest trader, its trade policy surely affects non-members. Second, as we discuss below, the EU has different types of preferential agreements, which allows us to theoretically derive and test a rich set of predictions. Finally, although the EU accounts for at least a fifth of world trade, there is hardly any empirical evidence on the formation of the EU's trade policy in general and none that analyzes how its PTAs affect its MTL.[2]

[1]The latest round was launched in 2001 and according to an article in the leaders section of the *Economist* magazine, a factor that may lead to its collapse is that "Poor countries with preferential access to rich world markets want to make sure that freer trade will not reduce these preferences" ("Talking the Talk", July 17th 2004, p. 14). However, the *possibility* that these preferences would reduce MTL is not new: it was a concern raised when the generalized system of preferences to developing countries was originally proposed (Johnson 1967, p. 166).

[2]Constantopoulos (1974) and Riedel (1977) examine some determinants of protection of individual members before accession to the EU and Tavares (2001) analyzes the determinants of the EU's common external tariff. Those authors use

Most of the early theory focuses on PTAs that are purely motivated by trade related issues and then analyzes their effect on MTL assuming that MTL implies free trade. Therefore, that research centers on how PTAs affect the binary choice between free trade and no MTL, and it effectively asks whether PTAs make a multilateral trade round more or less likely (cf. Krishna 1998 and Levy 1997). Assuming that a round leads to free trade and focusing on the probability of a round simplifies the analysis but generates predictions that are very difficult to test because these rounds are so infrequent. Moreover, countries can choose to conclude a multilateral round with considerable liberalization or one with almost none. Thus, the focus should be on whether PTAs affect the change in multilateral tariffs and not simply the probability of a round.[3]

The model we develop, which builds on Limão (2007), captures key features of the multilateral system and of recent PTAs, and generates several specific predictions that we test. The main one is that multilateral tariffs are higher on goods that a country imports duty-free from preferential partners (PTA goods, henceforth) than on otherwise similar goods (non-PTA goods). The basic intuition for this result is the following. Suppose the EU offers preferential duty-free access in a set of goods to a certain country. The latter benefits from facing lower tariffs than its competitors; the fact that the EU signs the PTA indicates that its member governments value it at given multilateral tariffs. If the EU eliminated its multilateral tariff on that same set of goods, it would effectively eliminate the PTA that it valued. We show that this additional cost of MTL is only present for the subset of PTA goods and affects multilateral tariff levels only when the preferential tariff is already zero since otherwise, the preferential tariff can be reduced to maintain the preferential margin. The model also predicts that there is no stumbling block effect if the PTA involves full accession because the EU can now easily offset any reduction in preferential margins due to MTL through a direct cash transfer to the preferential partner.

The central motive for the EU to offer trade preferences in this theoretical model is to extract concessions from its PTA partners on areas not directly related to trade. This is a key feature of many of the PTAs the EU had in place in 1994, as we describe in Section 2, and so we employ this model's structural equation for the equilibrium trade policy as the basis for our estimation. However, the empirical analysis will include all of the EU's PTAs and we will explain how our instrumental variables approach can address omitted variable issues that could arise when other motives for PTAs are important.

Using detailed tariff data we find that the EU's PTAs generated a stumbling block for its MTL in the last trade round. More specifically, the EU reduced its multilateral tariffs on goods not imported under PTAs by

industry data and we are not aware of any paper that has thus far analyzed product level protection in the EU.

[3]Bagwell and Staiger (1999a) analyze two opposing effects of PTAs on the equilibrium multilateral tariff *level* in a self-enforcing model. They show that PTAs are a stumbling block if countries are very patient and a building block otherwise. Another approach is due to Krugman (1991) who analyzes the welfare path for exogenously expanding trading blocs, which Bond and Syropoulos (1996) also analyze. Winters (1999) surveys this literature.

almost twice as much as on its PTA goods, as predicted by the model. We ensure that the results are robust to reverse causation and other possible sources of endogeneity by employing an IV-GMM estimator and testing for the exogeneity of different variables and the validity of our instruments. The stumbling block effect we estimate is stronger for goods that were exported by all of the EU's PTA partners. Various robustness tests provide further support for the baseline estimates. The effect is not present for goods with a positive preferential tariff or in EU accessions, which are two important auxiliary predictions of our model. In sum, EU accessions in the 1980's and 1995 did not significantly hinder (or promote) its MTL but all of the other PTAs in place in 1994 hindered its MTL. Both types of agreements are important for the EU so it is useful to recall that the latter are the source of the stumbling block effect.[4]

The results are also economically significant. The estimates imply that the average increase on prices received by exporters to the EU due to its multilateral tariff changes was only about half for PTA goods relative to other goods. Moreover, according to the theoretical model, our estimate represents not only the current wedge in the tariffs between PTA and non-PTA goods but also what the actual tariff wedge for the set of PTA goods would be relative to the counterfactual where the EU has no preferences for that *same* set of goods. That wedge is about 1.4 percentage points whereas the current average tariff for PTA goods in our sample is 4.7 percent.

There is a small but growing literature on this important question. Limão (2006) provides detailed evidence of a "clash of liberalizations" for the US in the Uruguay Round. His approach is similar to ours but we test a number of additional predictions that are relevant for the EU and we follow the structural equations more closely. However, it is possible that in other countries PTAs lead to lower protection against non-members. Foroutan (1998) finds lower average non-preferential tariffs for Latin American countries with PTAs after the Uruguay Round. She notes that no causality can be drawn from such a correlation because those countries were moving away from import substitution during the 1990s, which implied considerable unilateral liberalization independently of any effects from PTAs. Two other interesting papers on Latin America go much farther than establishing univariate correlations. Bohara, Gawande and Sanguinetti (2004) estimate that the Argentine unilateral tariffs were lower in industries where the value of imports from MERCOSUR to value added in Argentina was highest. Estevadeordal, Freund and Ornelas (2006) provide a systematic study for ten Latin American countries and find that preferences were associated with lower non-preferential tariffs in the period of 1990 to 2001. None of these three papers model multilateral negotiations so there is no systematic evidence that PTAs lead to more multilateral liberalization. Even if such evidence is found for Latin American and some other countries, it will be difficult to overturn the concern that PTAs slow down MTL because the current evidence supports this conclusion for two of the largest

[4] A narrow measure of their relative importance is total exports to the EU in 1994, which were about twice as high for the non-acceding PTA partners relative to the ones that acceded in the 1980's and 1995. Since 1995 the EU has pursued both types of PTAs.

traders, the EU and the US, which are central to this controversy.

Reciprocity in tariff changes is a key feature of our model and also of the leading economic theory of the GATT (Bagwell and Staiger 1999b). However, some economists question whether it is followed in practice (Finger, Reincke and Castro 2002). We empirically model the multilateral negotiation process and find that reciprocity was followed in the last trade round: the EU's tariff reductions were largest for products exported by countries that provided greater increases in market access. Finally, we model and provide novel evidence of the EU's internal political economy determinants of trade policy. The EU places some, but not much, additional weight on producer than consumer welfare. In this respect our findings are similar to structural estimates of the Grossman-Helpman (1994) model that also find small values for the US.

The paper is organized as follows. We start with a brief description of the EU's trade policy formation that guides the theoretical and empirical modelling. In Section 3, we model the interaction between PTAs and MTL and derive the main theoretical predictions. In Section 4, we first discuss our strategy for empirical identification and then analyze and quantify the estimation results. In the final section we summarize the main results and discuss their implications. All the proofs and details regarding the data are in the appendix.

2 The European Union's Trade Policy

Before its expansion in 2004, the EU's membership was composed of 15 countries that accounted for one third of the world output and more than 20 percent of world trade. The EU succeeded the European Communities that started in the 1950s as a customs union. Currently, its members form a single market with free movement of goods, services, capital, and labor. There is also a very strong element of cooperation between members in non-trade policies, particularly in issues with regional spillovers such as immigration, environment, development of poorer regions, foreign policy, and judicial matters.

The key actors in the formation of the EU trade policy are the European Commission and the Council. The Commission proposes, negotiates and enforces trade policy on behalf of the members. The Council, where each member's government is represented, is the decision maker with the power to approve or reject the Commission's proposals for trade policy negotiations and their eventual outcome. That is, the Council is decisive in approving the common external tariff that the Commission negotiates in multilateral trade rounds as well as any preferential treatment it negotiates with non-EU members.

The EU's expansion of preferential trade treatment has occurred both through increased membership—initially 6 countries and 15 by the time of our analysis (now 27)—and through numerous PTAs with non-members. Table A1 in the appendix provides some details about all of the latter that were in place by 1994 including their abbreviations—here we note only a few key points. First, several of the EU's PTAs do not require the partner to lower their tariffs. Second, many of these PTAs, e.g. preferences to developing countries through GSP and ACP, seek, and at times explicitly require, cooperation in non-trade issues such as labor standards, human rights, migration control, and combat against drugs. The PTAs with the Mediterranean countries are also similar in nature and are aimed at addressing issues with regional externalities, such as immigration.[5] These features are explicitly captured by our model. Several of the countries that benefit from this preferential treatment fear that MTL on the part of the EU will erode these preferences. Thus, they have at times opposed MTL but the EU itself has used the same argument to avoid liberalizing, which is central to our model.[6]

In the estimation we also consider the remaining preferences the EU had in place before the Uruguay Round: to EFTA members that did not eventually join the EU and to some Central and East European economies. These did involve reciprocal trade preferences. However, the East European countries had to comply with several side conditions such as environmental and intellectual property regulations. For the EU, the benefits of these side conditions along with the political integration in Western Europe likely outweighed the preferential treatment provided to the EU exports. A similar argument applies to the accession of Greece, Spain and Portugal to the EU. So our model will focus on this exchange of preferences on the part of the EU for cooperation in non-trade issues, appropriately modified in the cases of accessions where a common external tariff is applied.

3 Theory

In this section, we show how PTAs can induce higher non-preferential (i.e. MFN) tariffs. The model captures key features of the EU's PTAs, as previously described, by extending Limão (2007) along several important dimensions. First, we model a political economy motive for the use of tariffs, which is an important determinant of the cross-sectional tariff structure. Second, we allow for different types of PTAs, both without a common external tariff and with it as well as direct cash transfers across members. In Karacaovali and Limão (2005) we provide additional microfoundations for various parts of the model and show how the results extend in different directions. Here we focus on the simplest model that can provide a number of new predictions about PTAs that are relevant for the EU and derive the structural equations that guide the estimation.

[5] According to Jackson (1997, p. 160) "during the last twenty-five years or so the experience of the GSP in the GATT system has been that ... the industrialized countries often succumb to the temptation to use the preference systems as part of 'bargaining chips' of *diplomacy*." The conditionality of EU's concessions in exchange for cooperation has further been documented for instance in Grilli (1997).

[6] A stark example was the European Commission's argument that a cut in the price support of about 25 percent in EU sugar was not tenable because it would cause an income loss of 250 million euros to ACP countries, some of whom export sugar to the EU under preferential treatment. European Commission (2000), "Commission Proposes Overhaul of Sugar Market," Brussels, October 4$^{\text{th}}$ 2000, IP/00/1109.

3.1 Setup

We first describe the basic economic structure and trade pattern. We then model the government's objective and show how it depends on multilateral and preferential tariffs as well as the supply of regional public goods, which motivate the cooperation in non-trade issues behind many of the EU's PTAs.

Each of the two symmetric regional blocs is composed of two economies, which we denote by $j = L, S$. All variables in the "foreign" bloc are denoted with an "*". Each country produces a numeraire using labor—the only factor—in a constant returns process with productivity normalized to unity. The numeraire is freely traded so the wage is equal to unity in all countries. The supply of non-numeraire goods is fixed and equal to $X_i^j \geq 0$ units of good i in country j.

For expositional purposes we maintain the trade pattern as simple as possible and illustrate the basic results using only two goods (in addition to the numeraire). Country L imports good 1 from both L^* and the regional partner, S. Symmetrically, L^* imports good 2 from L and S^*. The small countries, S and S^*, only export to their respective regional bloc partners and we assume that the prices they receive by exporting are above the maximum threshold price that consumers in S and S^* are willing to pay. Thus, effectively the equilibrium demand for non-numeraire goods in S and S^* is zero, which, along with the fixed supply, implies that their net exports are price inelastic. This allows us to show that a PTA can lead to a stumbling block effect even if it takes place with small countries and leads to no changes in their export volume. This assumption is analytically convenient and also plausible for some but not all EU agreements, as we can see in Table A1. However, the qualitative results are similar if we relax this elasticity assumption, as we show in Karacaovali and Limão (2005).

Country L sets a specific tariff τ_1 on the imports from L^* and a preferential tariff $\pi_1 \leq \tau_1$ on S. The equilibrium domestic price in L for its import, p_1, is then derived from the market clearing condition:

$$M_1(p_1) + M_1^*(p_1 - \tau_1) + M_1^s = 0 \tag{1}$$

where $M_1(.) \equiv D_1(p_1) - X_1$ and $M_1^*(.) \equiv D_1^*(p_1 - \tau_1) - X_1^*$ are the net import demand functions for L and L^*; M_1^s is the respective function for S, which in equilibrium is $-X_1^s$ given that $D_1^s = 0$ as described above. We assume that S^* has no endowment of good 1. The domestic price in L^* for its export is simply the price in L net of the tariff, $p_1 - \tau_1$. A similar market clearing condition holds for good 2, which is imported by L^*. Thus the domestic prices in L and L^* for their respective imports can be written as functions of their own tariffs, i.e. $p_1(\tau_1)$ and $p_2^*(\tau_2^*)$. Implicitly differentiating equation (1) we can show that an increase in τ_1 lowers the exporter price, $p_1 - \tau_1$, and increases p_1. Note that because the net exports of S and S^* are price inelastic, the equilibrium prices are not directly affected by the preferential tariffs.

We now model the government objective and then derive the multilateral tariff in the absence of PTAs. In this part of the exercise we can abstract from the objective of small countries and so defer its presentation until Section 3.2.2, where we analyze the case with PTAs. Thus we also simplify the notation by dropping the country superscript for L's variables. The government in L sets trade policy and chooses the tax rate, e, to finance expenditures on a public good in order to maximize the following political support function

$$\begin{aligned} G(\boldsymbol{\tau}, \boldsymbol{\pi}, \boldsymbol{\tau}^*, e^s, e) \equiv\ & 1 - [\tau_1 M_1^*(p_1(\tau_1) - \tau_1) + \pi_1 M_1^s] + [\upsilon_1(p_1(\tau_1)) + \upsilon_2(p_2(\tau_2^*))] \\ & + [\omega_1 p_1(\tau_1) X_1 + \omega_2 p_2(\tau_2^*) X_2] + [\Psi(e, e^s) - e] \end{aligned} \tag{2}$$

We have normalized the size of the population to one so the first term represents aggregate labor income. The second set of terms in brackets represents tariff revenue. The $\upsilon_i(.)$ terms represent consumer surplus, which depend only on the good's own domestic price.[7] The first set of terms in brackets on the second line represent a weighted value of endowments, where $\omega_i \geq 1$. Note that if $\omega_i = 1$ for all i, the objective reduces to a standard social welfare function. Therefore $\omega_i - 1$ can be interpreted as a political economy weight the government places on suppliers.[8] Since we do not have data on export subsidies, and they are generally not permitted by the WTO, we focus on the case where the large countries do not have a motive to use them in the first place, i.e. $\omega_2 = 1$. This does not affect the stumbling block effect we derive. Note also that we can easily extend the objective in equation (2) to incorporate multiple import and export goods by simply adding their contributions to consumer surplus, endowment value and tariff revenue analogously to goods 1 and 2. Subsequently, when we extend the model in this way, the tariff vectors, $\boldsymbol{\tau}, \boldsymbol{\pi}$ and $\boldsymbol{\tau}^*$ will contain multiple elements.

The key difference between this government objective and the ones commonly found in other trade policy applications is the last set of terms in brackets on the second line. The first term, $\Psi(e, e^s)$, represents the benefit from expenditures on issues with a regional spillover such as the environment, human and labor rights, immigration, etc. We assume that this function is concave and separable in the provision of the local and regional public goods. The direct cost is captured by $-e$, the value of the lump-sum tax (measured in terms of the numeraire) that is used to finance the public good in L.

[7] We assume that the underlying consumer utility is additively separable and throughout we focus on a quadratic form of each good's subutility for L and L^* that gives rise to linear demand curves.

[8] This is a reduced form that can be obtained as a special case from a model where lobbying is given micro-foundations, such as in Grossman and Helpman (1994), provided that in that model the ownership of the specific factors is concentrated. In Karacaovali and Limão (2005) we show that equation (2) can be obtained as the objective for the EU that arises from bargaining between independent EU-member governments—a fair representation of the EU's trade policy formation, as we describe in Section 2.

3.2 Preferential vs. Multilateral Trade Liberalization

3.2.1 MFN Tariffs without Preferences

We first derive the MFN tariffs when PTAs are not allowed, which provides the natural benchmark to determine if PTAs hinder multilateral liberalization. Following Bagwell and Staiger (1999b), we model reciprocal trade liberalization in the WTO as a cooperative outcome between countries that generates a gain from overcoming a terms-of-trade externality.[9] Accordingly, most of the negotiations occur between large countries and follow what is known as the principal supplier rule: if, for a given product, country L is the largest exporter to L^*, then L^* proposes a tariff reduction to L in that product in exchange for L's tariff reduction on L^*'s exports to L. The MFN rule then requires this reduction to be extended to all other WTO exporters of similar goods. Given that both the EU and the rest of the world are several times larger than most of the EU's individual PTA partners we take L and L^* as the principal suppliers. In the estimation section we relax this assumption.

The large countries choose multilateral tariffs to maximize their joint objective. Given the symmetry we can concentrate on maximizing L's objective after imposing the condition that tariffs in their respective import sectors are equal, $\boldsymbol{\tau}^* = \boldsymbol{\tau}$. We abstract from enforcement considerations such as the ones addressed in Limão (2007). The equilibrium multilateral tariffs in the absence of a PTA are given by

$$\boldsymbol{\tau}^m \equiv \arg\max_{\boldsymbol{\tau}} \{G(\boldsymbol{\tau} = \boldsymbol{\tau}^*, \boldsymbol{\pi}, \boldsymbol{\tau}^*, e^s, .) : \boldsymbol{\pi} = \boldsymbol{\tau}\} \tag{3}$$

where the constraint $\boldsymbol{\pi} = \boldsymbol{\tau}$ precludes PTAs and, when L imports only good 1 we have $\boldsymbol{\tau}^m = \tau_1^m$. In the appendix we use the first-order condition (FOC) to derive the advalorem equivalent tariff, $t^m = \tau^m/p$, since it is the focus of our empirical work. For good $i = 1$ this tariff is implicitly given by

$$t_i^m = (\omega_i - 1)\frac{X_i/M_i}{\varepsilon_i} + \frac{M_i^s}{M_i^* + M_i^s}\frac{1}{\varepsilon_i^*} \tag{4}$$

where ε_i denotes L's import demand elasticity and ε_i^* is the foreign export supply elasticity it faces.[10] If good i is not exported by the regional partner, i.e. if $M_i^s = 0$, this expression is similar to several political economy

[9] Bagwell and Staiger (2006) provide evidence for their theory by showing that WTO accession leads to greater tariff reductions in products with higher initial import volumes. Broda, Limão and Weinstein (2006) estimate that several countries have considerable market power in trade in certain goods and use it to set higher tariffs prior to their WTO accession.

[10] Both of these are evaluated at the equilibrium tariff. Their definitions are $\varepsilon_i \equiv -M_i' p_i^*/M_i$ and $\varepsilon_i^* \equiv [\partial(M_i^* + M_i^s)/\partial p_i^*] \times [p_i^*/(M_i^* + M_i^s)]$. Note that import demand elasticities are typically estimated with respect to the domestic price but we define it slightly differently for the purpose of the model discussion and derivations. However, our empirical implementation will take this into account explicitly. See Section A.4 for details.

models (Helpman 1999). The tariff is increasing in the political economy weight, ω_i, and the inverse of the import penetration ratio, X_i/M_i. This term is weighted by the import elasticity for standard Ramsey taxation reasons. The last term represents an MFN externality effect and leads to higher tariffs. It arises if S does not participate in MTL directly because the MFN clause requires L and L^* to lower their tariffs on imports from all partners even if some did not reciprocally lower their own tariffs.

In the empirical work we must address the fact that the EU sets tariffs on multiple goods. The expression in equation (4) applies in such a setting to any good i that is not subject to a preference, i.e. whenever $\tau_i = \pi_i$, and whether PTAs are allowed or not. The reason for this is the additive separability of goods in the government's objective and the symmetry across regional blocs. These two assumptions imply that L can reciprocate any tariff reduction by L^* on the symmetric import sector independently of what occurs in the remaining goods. That is if L also imported an additional good, $i = 3$, symmetric to $L^{*\prime}$s import of $i = 4$, there would be an additional FOC for τ_3 independent of the one for the original import, τ_1. We relax this symmetry assumption in the empirical section. But it is useful to keep in mind that equation (4) captures the benchmark equilibrium rate for the subset of products in which S either does not export or does not receive any preferences even when PTAs are already pursued. Thus, we use these "non-PTA" goods as a control group in the estimation.

3.2.2 MFN Tariffs with Preferences

We first model preferential tariffs and then determine their effect on multilateral tariffs. The PTA between L and S is characterized by a bargaining solution where L grants preferential tariffs, $\boldsymbol{\pi} \leq \boldsymbol{\tau}$, in exchange for an increase in S's provision of the regional public good. To capture the asymmetry in size and bargaining power between the EU and its PTA partners we allow L to make a take-it-or-leave-it offer to S.

Country S also maximizes a political support function, G^s, given by a weighted sum of income net of the cost of providing the public good, $1 - e^s$, and the value of the endowment, which is exported to L.

$$G^s(\boldsymbol{\pi}, \boldsymbol{\tau}, e^s) \equiv 1 - e^s + (p_1(\tau_1) - \pi_1)\omega_1^s X_1^s \tag{5}$$

In the absence of a PTA we have $e^s = 0$ since we simplify by assuming that S places no weight on the regional activity valued by L. When S is a WTO member, the highest credible threat tariff that L can use is to revert to the MFN tariffs, i.e set $\boldsymbol{\pi} = \boldsymbol{\tau}$, as required by the WTO rules. So, for a given MFN tariff, S will accept to participate in the PTA if

$$G^s(\boldsymbol{\pi} = \boldsymbol{\pi}^p, \boldsymbol{\tau}, e^s = e^p) \geq G^s(\boldsymbol{\pi} = \boldsymbol{\tau}, \boldsymbol{\tau}, e^s = 0) \tag{6}$$

Since L extracts all the bargaining surplus from the PTA, the bargaining equilibrium level of e^s, denoted by e^p, is increased until this participation constraint holds with equality. This yields the following equilibrium level of e^s for a given preferential tariff margin of $(\tau_1 - \pi_1)$:

$$e^p = (\tau_1 - \pi_1)\omega_1^s X_1^s \tag{7}$$

Intuitively, L can require S to collect an amount of the lump-sum tax e^p (used to supply the regional public good) that is as high as the value that S places on the revenue transfer from L due to the preference. Higher weight is given to products with larger political influence in S. If S exports multiple goods then the expression on the right-hand side of equation (7) would be a sum over the goods it exports to L.[11]

The governments in L and L^* choose multilateral tariffs as before with a key difference. Now they take into account the effect that these tariffs have on the PTA by changing the preferential margin and consequently the provision of the regional public good in equation (7). Hence, the equilibrium MFN and preferential tariffs are given by

$$\{\boldsymbol{\tau}^{mp}, \boldsymbol{\pi}^p\} \equiv \arg\max_{\boldsymbol{\tau},\boldsymbol{\pi}} \{G(\boldsymbol{\tau} = \boldsymbol{\tau}^*, \boldsymbol{\pi}, \boldsymbol{\tau}^*, e^s, .) : \boldsymbol{\pi} \leq \boldsymbol{\tau}; e^s = e^p\} \tag{8}$$

As we show in the appendix (Section A.1) this yields the following equilibrium advalorem MFN tariff for a good exported by S to its PTA partner, $t_i^{mp} = \tau_i^{mp}/p_i^*$.

$$t_i^{mp} = (G_{e^s}\partial e^p/\partial \tau_i - X_i^s)\frac{1}{M_i p_i'}\frac{1}{\varepsilon_i} + (\omega_i - 1)\frac{X_i/M_i}{\varepsilon_i} + \frac{M_i^s}{M_i^* + M_i^s}\frac{1}{\varepsilon_i^*} \tag{9}$$

The key difference relative to the tariff if PTAs were forbidden, t_i^m, is the first term, which captures the potential for a stumbling block effect due to the PTA, i.e. the potential for $t_i^{mp} > t_i^m$. To interpret and sign this effect, note that $1/M_i p_i' \varepsilon_i$ is positive, so the sign depends only on whether $G_{e^s}\partial e^p/\partial \tau_i > X_i^s$. That is on whether the marginal benefit for L of increasing the preferential margin and obtaining additional e^s, exceeds the marginal cost from lost tariff revenue, X_i^s. This depends on whether the preferential tariff is positive or not. To see this we can use the FOC for the preferential tariff, π_i, that requires

$$-(G_{e^s}\partial e^p/\partial \pi_i + G_{\pi_i}) \geq 0 \tag{10}$$

[11] When L sets the MFN tariffs it assumes that S accepts the PTA and that if S were to renege on it, then L would remove the preference and set its tariff on S equal to the MFN tariff originally agreed upon with L^*. This is perhaps easier to understand in a setup where cooperation is maintained due to repeated interaction, as in our working paper. In the repeated setup case we simply require that if at a future date S stops supplying the regional good, then the large countries do not renegotiate their MFN tariffs. We think this is a plausible assumption in practice given the costs of renegotiating MFN tariffs between rounds. Nonetheless, we can also consider the alternative when the threat point in equation (6) uses the equilibrium MFN tariff without a PTA. This alternative introduces some changes but we obtain similar qualitative results and in particular the condition that we will show is key for obtaining a stumbling block effect—a duty-free preferential tariff—is still necessary in this case (proof available on request).

When the equilibrium preferential tariff is zero, we are at a corner solution and so this condition holds with strict inequality. In this case the marginal benefit to L of *increasing* the preferential margin exceeds the marginal cost but it can not increase the margin by lowering the preferential tariff when it is already at zero, therefore it increases the MFN tariff. So the PTA has a stumbling block effect when the preferential tariff is already zero. If the PTA good is imported under a positive preferential tariff then (10) holds with equality and so the MFN tariff is unchanged, i.e. $t_i^{mp} = t_i^m$. To see this formally we note that $\partial e^p/\partial\tau_i = -\partial e^p/\partial\pi_i$ (equation (7)) and $G_{\pi_i} = X_i^s$ (equation (2)) such that $G_{e^s}\partial e^p/\partial\tau_i - X_i^s = -(G_{e^s}\partial e^p/\partial\pi_i + G_{\pi_i})$. A sufficient condition for a corner solution in the PTA is for L to place enough weight on S's provision of the regional good.

In sum, the intuition for the stumbling block effect is as follows. When the marginal benefit for L of increasing e^s is higher than the cost in terms of the foregone tariff revenue, L prefers to increase the preferential margin given to S and it initially does this by reducing the preferential tariff. However, once the preferential tariff is at zero, the preferential margin can only be increased by raising the MFN tariff and that is the optimal action for L.

A natural definition of the stumbling block effect is the difference of the MFN tariffs with and without PTAs.

$$t_i^{mp} - t_i^m = (G_{e^s}\omega_i^s - 1)\frac{X_i^s}{M_i p_i'}\frac{1}{\varepsilon_i} \geq 0 \tag{11}$$

The focus of the estimation is to identify this term and test if it is positive. Before we describe exactly how we test this and other predictions we briefly note how this effect is also present in more realistic setups.

If the large countries also placed a political economy weight on their exporters then, in the absence of export subsidies, the equilibrium MFN tariff would be lower. This occurs because each country now places a higher weight on price increases for its exporters. It is simple to show that the export "lobby term" that arises in this case is identical with or without PTAs so it does not affect the stumbling block effect.

When we relax the inelastic export supply assumption for S, the difference between the MFN tariffs with and without PTAs includes an additional term due to the trade volume effect. More specifically, the PTA causes an increase in S's exports to L and a decrease of L^*'s exports to L, which causes a decline in the MFN tariff. This channel is similar to the tariff complementarity in Bagwell and Staiger (1999a). The extra trade term attenuates but does not fully offset the stumbling block effect, as we show in our working paper.

We also test two other interesting predictions that follow from the model. First, if the effect is present, then it should be stronger for the PTA goods that are more important exports of S to L since a given margin applies to a higher export volume. Second, when S exports multiple goods, the effect is still present in the set of goods where the preferential tariff is zero but not where it is positive. In the latter case the preferential tariff can be lowered, which is less costly to L than increasing the MFN tariff in *that* good but leads to the same benefit for S.

The last result with multiple goods may not be immediately obvious but we prove it in the appendix (Section

A.2); here we provide the intuition. The EU can and often chooses to provide preferences for only a subset of products and for some of them it sets positive preferential tariffs, e.g. in agreements that do not fall under GATT article XXIV. Why would the EU ever increase the margin in good 1 by distorting its MFN tariff if it could increase its overall transfer to S by lowering the preferential tariff in some other good it imports, say good 3? According to this model the EU would do so when an increase in the preferential margin to good 1 is more valuable to S because its suppliers have more political influence, i.e. $\omega_1^s > \omega_3^s$. So, even though the cost to L of increasing the margin for good 1 by distorting the MFN may be slightly higher, so is the benefit in terms of the amount of regional public good it obtains, as we can see in equation (7). In sum, with multiple goods we must consider the EU's optimal allocation of preferences across goods, which depends on the political weights in S. The model predicts that the stumbling effect is present in the set of goods where the preferential tariff is zero but not in those where it is positive. Thus the model not only provides us with a structural equation to guide the estimation but also a set of predictions that are not obvious from the outset.

3.2.3 Common External Tariff and Direct Transfers

Most results on PTAs depend on whether they have a common external tariff (c.f. Cadot et al. 1999 or Bagwell and Staiger 1999a). Given that between the Tokyo and Uruguay Rounds—the period we analyze in the empirical work—the EU accepted new members that share its common external tariff (CET), we extend the model to analyze the effect of such accessions on MTL.

The use of a CET raises several interesting questions; e.g. which is the optimal country in the union to decide on the CET (Syropoulos 2002) and, more importantly for our purposes, how the tariff revenue is to be distributed over the different countries, e.g. if all goods enter the EU via one port, does that country receive all the revenue? Clearly not; so PTAs with a CET must agree on revenue transfer mechanisms. Therefore, one key difference relative to other PTAs is the existence of a mechanism for transfers, which L can also use to "purchase" the supply of the regional good. We note that the willingness to implement such transfer schemes is often limited, which along with the need to agree on a CET explains why customs unions are rare relative to other PTAs. The implicit costs in the use of direct transfers explain why we ruled them out in deriving the stumbling block effect in the absence of a CET.[12] However, those countries that can agree on a CET clearly do not face prohibitive costs of direct transfers

[12] Using direct transfers may also not be the most efficient way to transfer resources to other countries, as the aid vs. trade literature highlights, or to reward cooperation since the direct transfer may end up in the pockets of a politician without providing the best incentives for cooperation. For example, one of the stated aims of the US in providing preferences to the Andean countries is to raise the relative price of activities other than drug production. Alternatively, the partner country government may place a positive political economy weight on its exporters, $\omega_i^s > 1$, such that it benefits more from a preference that increases the income in good i by \$1 than a lump-sum transfer of \$1 to the government's coffers. Political economy constraints that reduce the effectiveness of direct transfers relative to preferences are present in practice, otherwise

since they use them to redistribute revenue. So the PTA solution must now explicitly allow transfers.[13]

As we prove in the appendix (Section A.1.2), the stumbling block effect disappears under this case. However, despite the ability to use transfers, a preferential rate may still be used because, for given multilateral tariffs, L is indifferent between a transfer *and* preferential tariff reductions. At a given MFN tariff, the cost for L of a reduction in a preferential tariff π is simply the lost tariff revenue, which is no more costly than transferring an equivalent amount in the numeraire good. The difference relative to the PTA without a CET is that now if, at $\pi = 0$, S is still not providing "enough" e^s, the optimal solution is for L to increase the transfer rather than the MFN tariff, since the latter distorts the prices. Thus, in equilibrium the MFN tariff is not affected by this type of a PTA.

When S exports multiple goods there is another feature of a PTA with a CET that makes a stumbling block effect less likely even in the absence of transfers. Customs unions are usually formed between partners in the same region, so their trade values are often large across many goods. Often in these agreements, and certainly for the EU accessions that we consider, the preferential tariffs cover all goods and are set to zero. This feature may be due to different constraints, e.g. compliance with GATT article XXIV (under which EU accessions fall), some cost of maintaining border controls to levy positive preferential tariffs on a subset of goods when all other goods and factors are freely mobile, etc. The constraints that lead L to set duty-free preferences on most of S's exports make it less likely that L would also need to set a higher MFN tariff (or make transfers) to obtain the regional public good relative to a situation where L is unconstrained in the choice of preferential rates for individual goods. In sum, the broad application of duty-free rates in EU accessions lowers the need to distort the MFN tariff to increase the preference margin in specific goods.

4 Estimation

4.1 Predictions and Identification

We now derive the model's estimating equation, point out its main predictions and analyze how it is identified.

The general expression for the MFN tariff on a product i at time t nests the case where the good receives a preference (equation (9)) or not (equation (4)). Simplifying the notation by using x_i for $\frac{X_i/M_i}{\varepsilon_i}$, m_i for $\frac{M_i^s}{M_i^* + M_i^s}\frac{1}{\varepsilon_i^*}$ and ϕ_i for the stumbling block term, $(G_{e^s}\omega_i^s - 1)\frac{X_i^s}{M_i p_i'}\frac{1}{\varepsilon_i}$ in equation (11), we have :

$$t_{it} = \phi_{it} I_{it} + (\omega_{it} - 1)x_{it} + m_{it} \tag{12}$$

we would not observe several of the current preference schemes.

[13] In fact, the transfer is the only difference relative to the PTA without a CET since, in the simple case we consider here, S does not consume any non-numeraire goods so the CET is not binding for S. In the working paper we show that the result is identical if S consumes some of its export and the CET binds for it.

where, according to the model, I_{it} is an indicator for whether good i is imported from a preferential partner and receives a zero preferential rate at time t. The econometric model in error form is then

$$t_{it} = \phi_t I_{it} + \beta_t x_{it} + m_{it} + u_{it} \tag{13}$$

where $E(u_{it}|x_{it}, m_{it}, I_{it}) = 0$ and β_t measures the average extra weight the government places on importers. The key parameter is ϕ_t, which measures the *average* stumbling block, i.e. the average difference of MFN tariff levels with and without PTAs as defined in equation (11); the central question is whether it is positive. We also augment the econometric model to test three other predictions. Namely that this effect is: (i) stronger for goods with higher PTA exports; (ii) not present for goods with a positive preferential tariff and (iii) not present for goods exported by countries that recently joined the EU and have access to transfers.

The theoretical model captures key features of trade policy determination. However, it is parsimonious and possibly not fully specified, e.g. tariffs may be affected by foreign lobbies and tend to be highly persistent, which may be due to various unobserved product effects. Such effects may also influence whether a good receives a preference and thus generate an omitted variable bias. We address this by estimating the model in differences and employing instrumental variables. Since the model focuses on MFN tariffs, which the EU changes very infrequently, we take the difference between the MFN tariffs negotiated in the Uruguay Round (UR) and those in place before it. The latter were largely set during the Tokyo Round since the WTO shows no record of renegotiation for the EU between these rounds.

Estimation in differences can still provide an estimate of the *level* of the stumbling block effect given the timing and properties of certain EU PTAs. Some PTAs were not in place during the Tokyo Round, so $I_{it-1} = 0$ for certain goods, and so the EU's multilateral tariff in the goods exported by those countries should not have been affected. This is also true for several products of PTAs that were in place at $t-1$ because the EU did not offer duty-free access to several goods. As we describe in Table A1 there have been important extensions in the coverage of preferences between rounds.[14] Moreover, the range of products that each PTA can expect to export has increased substantially as trade costs fell. In sum, a large range of products changed status from $I_{it-1} = 0$ to $I_{it} = 1$ and when we difference equation (12) for those products and write it in error form we obtain

$$\Delta t_i = \phi_t I_{it} + \beta \Delta x_i + \Delta m_i + \Delta u_i \tag{14}$$

[14] For instance, the preferences have been expanded through a number of revisions for MED, ACP, GSP, EFTX, which are the agreements that initially took effect before the Tokyo Round. These revisions also included new requirements on cooperation relating to human rights and democracy from all recipients (Brown 2002, Raya 1999). In addition to that, the European Economic Area with EFTA in 1992 provided a much deeper economic integration between partners.

where we assume the weights, ω_i, are time-invariant—an assumption that we test and find to be reasonable in our working paper. Therefore, in this case β and ϕ_t have the same interpretation as in the level equation (13). Figure 1 illustrates why this is the case. We plot the tariffs for two goods that are similar except that one becomes a PTA good between the Tokyo and Uruguay rounds. The tariff increase up to the dashed line, i.e. $t^{mp} - t^m$, indicates the stumbling block effect predicted by the model if the preference is duty-free. The EU may choose not to change the bound multilateral tariff immediately because this would impose a renegotiation cost (as well as the costs from higher tariffs on EU products that other countries would be allowed to set) but, when the new round occurs, the difference in the reduction in the two products reflects the stumbling block effect. That is $(t^m - t^{m'}) - (t^m - t^{mp'}) = t^{mp'} - t^{m'} = \phi_t$. The same prediction applies to goods that, at the time the UR was negotiated, were expected to have a duty-free preference.

[FIGURE 1 HERE]

Some products were already PTA goods during the Tokyo Round and therefore did not change status. In this case an OLS estimate of the parameter on I_{it} in equation (14), call it $\tilde{\phi}$, reflects a weighted average of the level effect for new PTA goods and the change in the stumbling effect for the existing PTA goods. We show this in Section A.3 of the appendix where we also argue that $\tilde{\phi}$ represents a downward biased estimate of the level effect, ϕ_t. The basic reason is the attenuation bias due to measurement error. For continuing PTAs we do not have information on which subset of products were already PTA goods at $t-1$. Thus an instrument that is correlated with the new PTA goods in I_{it} but not the existing ones reduces or eliminates this bias. A comparison of OLS and alternative IV estimates supports this view, as we will show.

One potentially important determinant that is not explicitly reflected in equation (14) is reciprocity—the extent to which the EU lowered its tariffs in response to other countries' reductions in the UR. Reciprocity is an important principle in WTO negotiations and a basic feature of our theoretical model; it is not reflected in equation (14) because we assumed symmetry across the regional blocs and then solved for the equilibrium tariff. By relaxing this assumption and controlling explicitly for reciprocity we minimize the possibility for omitted variable bias.

We follow Limão (2006) who constructs a measure of market access concessions that is consistent with the practice in multilateral tariff negotiations. The variable is defined at the product level as
$R_i = \sum_k s_{it}^k [\sum_j w_j^k \Delta t_j^k / t_{jt}^k]$ where $\Delta t_j^k / t_{jt}^k$ is the percentage tariff reduction by a non-EU country k in good j and w_{jt}^k is the import share of good j in total imports of k. Therefore the term in brackets captures country k's average concession, which is multiplied by the export share of a principal supplier k in good i to the EU, s_{it}^k. An exporter is a principal supplier of good i if it is one of the top 5 exporters of good i to the EU. The prediction is that if k offers relatively larger concessions, then the EU reciprocates through larger MFN tariff reductions in the products it imports from k. We also test and find that the results are similar if we follow the structural model more literally and do not include this variable.

The final issue in deriving the estimating equation is data availability for the MFN-externality variable, m. We do not have a bilateral record of which countries negotiated with the EU on each 8-digit product during the trade rounds and therefore we can not construct the exact variable, Δm. But we proxy for it by using information on the share of small exporters by product, i.e. the share of those countries that are not one of the top-5 exporters in product i to the EU. Increases in this share between the rounds imply that the probability of an MFN-externality increases, since the EU would have to negotiate with more exporters each of whom now has a higher incentive to free-ride. We consider the change in this share between 1994 and 1989, the earliest year when this 8-digit data is reported. If the change in this period is sufficiently large, then it will be positively correlated with the change over the full period between rounds. We capture this by constructing an indicator variable, P, for whether the change in this share for good i is above the median change.[15] The model then predicts a positive coefficient on this variable.

Introducing the proxy for Δm and augmenting equation (14) to explicitly account for reciprocity yields our basic estimating equation:

$$\Delta t_i = c + \phi_t I_{it} + \beta \Delta x_i + \rho R_i + \mu P_i + v_i \tag{15}$$

where we include a constant term, c, and modify the error term to explicitly allow for the measurement error in classifying new PTA goods (and possibly in the proxy and reciprocity variables). Since we are interested in establishing causality, we now discuss how we address the potential endogeneity issues in estimating (15).

The most important endogeneity concern is reverse causation since the preference may depend on MFN tariff changes, e.g. if a PTA partner expects a small MFN reduction in a product, it is more likely to request a preference in it than in a product where it expects a large MFN reduction. To tackle this we employ instrumental variables. The central variable is the PTA good indicator, $I_i^{pta} = PR_i^{pta} D_i^{pta}$, which is equal to one if good i faces a zero preferential tariff, i.e. $PR_i^{pta} = 1$, and that PTA partner exports it to the EU, $D_i^{pta} = 1$. Since the main reverse causation concern arises due to the preference component, the main instrument we employ for I_i^{pta} is D_i^{pta} for the year 1994 regardless of whether it receives a preference or not. This instrument is clearly correlated with I_i^{pta} but we expect it to be uncorrelated with the error term in (15) since the changes in the MFN tariff we use as a dependent variable are implemented starting only in 1995. When we consider goods exported duty-free under any or every PTA we adjust the instrument accordingly so it reflects whether good i was exported by any or every PTA.

To predict whether a good was likely to receive a preference around the time of the UR, we also include as an instrument whether the EU set non-tariff barriers on the good in 1993. A country is more likely to request a preference in a good if it expects that otherwise it would be subject to an NTB. This effect would be magnified if

[15] The estimation results do not change when we employ 75^{th} or 90^{th} percentiles instead.

the country already exported this product and the NTB applied to all countries (the variable D^{ntball}), hence we interact these two variables as well. These instruments are motivated by two related effects. First, an NTB can reduce the price received by exporters to the EU and a good that receives a preference is less likely to face an NTB. Second, NTBs tend to raise domestic prices in the EU, which is the price that the preferential exporter receives if it faces a zero preferential tariff and no NTBs. Since setting an NTB on all partners can also indicate some other good characteristic that may affect MFN tariff changes, we also show that the results are robust to dropping the instruments that include D^{ntball} and its interactions.

The IV-GMM approach we employ allows us to carefully test the exogeneity of these instruments via alternative over-identifying restriction tests. The set of over-identifying restrictions that permits us to perform these tests arise from including other instruments, such as "world" price changes between 1992 and 1994, which are unlikely to depend on the changes in MFN tariffs that only take place in the subsequent years. There are two related channels by which price changes help to predict the identity of a PTA good, I_i^{pta}. First, a PTA partner is more likely to export a good in 1994 if between 1992 and 1994 there were large price increases that now allow the exporter to overcome any fixed trading costs in i. Second, the benefit of receiving a preference on some fixed level of exports X_i^s rather than facing an advalorem MFN tariff t_i is equal to $p_i^w t_i X_i^s$, so it is increasing in the world price. Thus increases in p_i^w raise the incentive for the PTA partner to negotiate a preference in this good. Since these effects are likely to vary with the price changes, we allow for non-linearities by also including higher order terms.

Another benefit of the IV-GMM approach is that it allows us to address potential omitted variables bias. This could arise for a number of reasons including the possibility that some of the EU's PTAs reflect additional motives not modeled in the theory. So, the IV approach allows us to include all of the EU's preferential programs in 1994 in our empirical analysis and obtain estimates for the stumbling block effect, ϕ, even when other motives for PTAs are present. We cannot rule out the theoretical possibility that under alternative motives for PTAs, our exclusion restrictions for the instruments are invalid. But ultimately this is an empirical question, which we address with the over-identification restriction tests just described and by testing the robustness of the results to excluding different sets of instruments.

The variable that captures the political economy effect, Δx, is likely to depend on the MFN tariffs since it involves the production/import ratio weighted by the import demand elasticity, all of which are functions of the EU's domestic prices and hence its MFN tariffs. Therefore, we employ the levels of these variables before the MFN tariff is implemented, e.g. 1978 for x_{t-1} and 1992 for x_t. If these variables exhibit persistence over time, using the lag values will not address the endogeneity so we also instrument Δx. We employ the change in a measure of scale economies (Value added/number of firms) and its interaction with the average world price change in the *industry* between 1992 and 1994. All else equal, industries with large fixed costs of entry and thus high economies of scale are likely to have higher X/M.[16] World price changes between 1992-94, on the other hand, directly affect domestic prices, which are important determinants of all the components of x, i.e. X, M, and ε but do not depend on the tariff changes on the left hand side, which are implemented only after 1994.

Finally, the reciprocity variable is another potential source of endogeneity due to reverse causation since the total tariff reduction by other WTO members in the UR partially depends on EU reductions. Our instrument is the unilateral portion of the total tariff reductions that were eventually offered at the UR. More specifically, several countries undertook unilateral trade liberalization between 1986 and 1992. Their liberalization was unilateral because it was undertaken outside of GATT negotiations without an expectation that it would be reciprocated since, the very completion of the round was in doubt until 1992. However, when the final multilateral cuts were negotiated, between 1992 and 1994, the unilateral reductions undertaken from 1986 to 1992 were explicitly reciprocated because they had taken place after the official start date of the round (Finger et al. 2002). Therefore, we employ the unilateral liberalization by WTO members between 1986 and 1992 as an instrument for what was eventually used as a basis for their reciprocal liberalization—the total liberalization between 1986 and 1995.

4.2 Data

In Section A.4 we provide detailed definitions of the variables, their construction, and sources; Table A2 contains summary statistics. Here we note some of its salient characteristics. We employ the advalorem MFN tariffs from the WTO schedules of concessions, and the preferential tariffs from UNCTAD, both at the 8-digit Harmonized Standard (HS) level.[17] To construct the reciprocity variable, we employ the data in Finger et al. (2002). They use the available tariff reductions for each WTO member during the UR, and aggregate it from the product level into country-averages. We take these average country concessions and construct a product specific measure of reciprocity by using those countries' export shares to the EU by 8-digit product (from EUROSTAT).

We use data on production and other industry-level variables for constructing x and its instruments. This data

[16] According to the model, and similar to Grossman and Helpman (1994), once we have accounted for the size effect, x, in the protection equation, other variables, such as scale, should have no direct effect and can thus be excluded. This is the rationale in Goldberg and Maggi (1999) to exclude a similar scale variable from the main trade protection equation and use it as an instrument for X/M and political organization.

[17] The model prediction applies to goods that have or are expected to have a duty-free preference when the UR was negotiated. Hence we employ the preferential rates reported for 1996 for all PTAs except where unavailable (EFTA, for which we use 1993). Moreover, we exclude products with a zero MFN tariff before the UR for two reasons. First, when the MFN tariff is zero there is often more noise in the data about whether a preference exists or not, since it is in effect irrelevant. Second, all the tariffs in the sample that were initially zero remained unchanged and are likely to share an unobserved common characteristic. Thus, including those observations would bias the estimates if the proportion of zero tariffs is different for PTA goods relative to the rest of the goods.

is available for individual EU members and we aggregate it as suggested by the theoretical model.[18] UNIDO's industrial database provides the most comprehensive source covering all EU members and dating back to 1978. It is collected at the industry level and hence more aggregated than the trade and tariff data. We use clustering at the industry level to correct the standard errors for this fact.[19] Since UNIDO does not provide production data for agriculture, we exclude those products, but processed agricultural products are included. Given the prevalence of non-tariff barriers and EU subsidies in agriculture we don't believe this is a drawback since an analysis that focuses on tariffs without taking these other forms of protection into account could be inappropriate for agriculture.

In Table 1, we present the tariff levels and their changes for our sample. Although our analysis is conducted at the product level, we provide some statistics aggregated by industry. The highest tariff rates before and after the UR appear in the tobacco sector: an average of 42 and 25 percent respectively. The lowest pre-UR tariffs are in the miscellaneous petroleum and coal products sector with 3.9 percent, whereas the iron and steel industry became the least protected in terms of tariffs after the UR, 0.4 percent. The footwear sector experienced the least liberalization, 0.8 percentage points, and tobacco the highest, 17. Note also that there is a considerable amount of variation in tariff changes both within industries, with coefficients of variation between 0.28 and 1.5, and across industries, with a coefficient of variation of 0.44.

[TABLE 1 HERE]

4.3 Estimation Results

The unconditional mean reduction in MFN tariffs by the EU during the UR was 4.4 percentage points for non-PTA goods but only 2.9 for goods exported duty-free by any PTA partner. A simple t-test confirms that the difference of 1.5 percentage points, with a standard error of 0.1, is statistically significant. This difference may be

[18] In Karacaovali and Limão (2005) we show that equation (2) can be obtained as the objective for the EU that arises from bargaining between independent EU-member governments, which suggests that we should simply add production for each industry over the EU members to obtain X_i and divide it by aggregate imports in i. The weight, ω_i, can then be interpreted as an average of the individual member weights, $(\omega_i - 1) = \sum_c (\omega_i^c - 1)\xi_i^c$ where ω_i^c is individual member c's weight for a given producer and ξ_i^c is the production share in the EU.

[19] In calculating the variable $(X_i/M_i)/\varepsilon_i$ the remaining variables that we employ are at the same level of aggregation as the production data. The fact that this data is more aggregated could potentially introduce some measurement error. Although we can not rule out this possibility, we note that it may not be such an important concern for the following reason. If the EU negotiators use the production data at the most disaggregated level available for most of its members, as we do, then our measure is actually the relevant one. The interpretation of β is now as the average EU-wide extra weight taken over the different industries rather than products. We are comfortable with this interpretation, since producers tend to organize at the industry level to lobby for protection. This is particularly true in the EU, where there is more variation in protection across industries than within them. Therefore, to the extent that the extra weight reflects a political economy motive, the best way to identify it is at the industry level.

due to other factors that are correlated with the PTA variable. Therefore, in Table 2 we present the estimates of the parameters in equation (15). In order to address the endogeneity issues discussed above, we use the two-step efficient generalized method of moments estimator (IV-GMM), which is robust to heteroskedasticity with an undetermined form and we cluster at the industry level.

4.3.1 Stumbling Block Estimates

The indicator variable I^{any0} in Table 2 takes the value one if the EU imports the good from *any* partner at a duty-free preferential rate. It excludes countries that acceded as full EU members after the Tokyo Round, which we estimate separately as suggested by the theory. The coefficient for I^{any0} provides an estimate of ϕ and we find that it is positive and significant at the 1 percent level under all specifications. This provides evidence of a smaller reduction in the EU's MFN tariffs for its PTA goods (with a zero preferential tariff) relative to its non-PTA goods as predicted by the model. The effect is similar if we use I^{any}, which is equal to one if the good has any preference. But we will focus on a baseline that uses I^{any0} as the model requires and test if the effect is different for goods with positive preferences. Before quantifying the importance of this stumbling block effect, we test other predictions.[20]

[TABLE 2 HERE]

The model also predicts a stronger stumbling block effect for important exports of a PTA partner. We test this by introducing an additional variable, $I^{hi\,\exp}$—I^{any0} interacted with $D^{hi\,\exp}$, where $D^{hi\,\exp}$ is one if the share of a PTA partner's exports in good i relative to its total exports to the EU is above a certain threshold. In column 3 of Table 2 we estimate that such an *extra* effect is present and significant. The results are qualitatively similar if we use different thresholds such as the median or 75^{th} percentile instead of the 25^{th}.

Although we did not explicitly model simultaneous PTAs we expect that in such an extension, if a product is exported by several preferential partners, then an increase in the margin of preference that benefits every partner generates a stronger stumbling block effect. In effect, it is equivalent to an increase in preferential exports from a given partner. We test this in column 2 of Table 2 by including an additional variable, I^{evy0}, which is an indicator for whether the EU imports the product at a tariff of zero from *every* preferential partner. We estimate a significant *additional* effect of 0.7 percentage points for this subset of goods.

Standard Hausman tests reject the null of consistency of the OLS estimates for some but not all specifications (the p-value ranges from 0.04 to 0.52 across different specifications). Moreover, there are good reasons to expect the

[20] A potential cost of using data that is finely disaggregated is that it is more likely to suffer from product misclassification when a shipment is recorded. We try to minimize this problem by classifying a good as being exported by a PTA to the EU only if the value registered in that year is above a certain low threshold. In our estimations, we employ the 5^{th} percentile of the value of a given PTAs' exports in that year as the threshold. The baseline results using a different threshold, the 10^{th} percentile, are similar. The results without imposing any threshold are identical and reported in our working paper.

OLS estimates to be biased, both because of reverse causality and measurement error, as we noted in Section 4.1. Therefore we focus on the IV estimates, which may be inefficient but consistent over all the regressions. We also calculated the OLS counterpart for each specification and found that the stumbling block results are qualitatively similar to the IV estimates. In column 4 of Table 2 we confirm this for the baseline specification; the coefficient on I^{any0} is 0.013 and it is significant at 1 percent. It is smaller than the IV estimate indicating some attenuation bias, which we will further address in the robustness section.

According to the model, MFN tariff changes for products imported from countries that joined the EU between the last two trade rounds should be identical to those of other products. We find this to be true in the data. In column 1 of Table 3 the variables I^{afs} and I^{spg} are indicators for products exported by Austria, Finland, and Sweden and Portugal, Spain, and Greece respectively, which are statistically insignificant. The stumbling block effect generated by the PTAs that do not share a common external tariff with the EU remains unchanged both in magnitude and significance.

[**TABLE 3 HERE**]

According to the model the stumbling block effect is only present for products with a zero preferential tariff. We test this in column 2 of Table 3 where the variable I^{any} takes the value one for the goods imported by the EU at a preferential tariff rate—either zero or positive—whereas I^{pos} is one for the subset of goods with a positive preferential tariff. The sum of these coefficients measures the total effect of a good with a positive preferential tariff and a formal test shows that we cannot reject the hypothesis that the tariff reduction for such goods is identical to the non-PTA goods. This prediction is fairly specific to the model and thus its confirmation provides it strong support.

The estimates presented so far refer to an average effect of all of the EU's PTAs. It is also interesting to test whether the effect is driven by any given PTA in particular. Although there is a positive correlation among the variables for different programs, we do identify a stumbling block effect originating from each in column 3 of Table 3. All individual effects are significant with the exception of the one for ACP, which is nonetheless significant when tested jointly with GSP, a program with preferences highly correlated with those in ACP. It is also worth noting that the effect for PTAs signed after the Tokyo Round, such as CEC where we know for sure that the classification as a new PTA good is not subject to measurement error, is positive and significant.

In Table 4, we present the first stage regressions for some of the main specifications, which indicate that the instruments are jointly significant in all of our specifications with F-statistics higher than 10. Moreover, the rows at the bottom of Tables 2 and 3 labeled "Hansen's J" show that we cannot reject the null hypothesis that the excluded instruments are uncorrelated with the second stage error term, and correctly excluded from the estimated equation at any conventional level of significance. When the set of instruments is large, this test may have low power. Therefore, we also test the subset of instruments that are a priori more likely to be endogenous, such as the export dummy and NTB variables. The results are found in the row "C-Stat" and indicate that we cannot reject the orthogonality of the smaller subsets either.[21]

[TABLE 4 HERE]

In sum, the evidence supports not only the model's main predictions but also its auxiliary ones, which rely on variation between PTA and non-PTA goods as well as variation within PTA-goods (e.g. every PTA, high exports and positive preferential tariffs). This considerably increases our confidence that the results are not spurious. We will consider a number of additional robustness tests including alternative instruments, sample, and regressors. Since the main message will remain unchanged, we defer this to Section 4.5 and first quantify the effects and discuss the remaining determinants of EU multilateral tariffs.

4.3.2 Reciprocity and Political Economy Determinants of EU Tariffs

Tariff changes are notoriously hard to predict and in fact most empirical studies that employ a structural approach focus on explaining the cross-section. Nonetheless, given how sparse the evidence is for the EU's trade policy determination, we think that it is interesting to ask whether the remaining variables of our parsimonious model have any explanatory power.

The coefficient on Δx provides an estimate of $\omega - 1$, which can be interpreted as the production weighted average of the extra importance attached to producer surplus relative to social welfare in the EU. We find it to be positive with the IV estimates ranging from 0.003 to 0.004. To our knowledge these are the first such estimates for the EU. Goldberg and Maggi (1999) estimate this extra weight to be approximately 0.014 for the US, whereas Gawande and Bandyopadhyay (2000) estimate a much lower value of 0.0003. Thus, our estimate for the EU lies in between these for the US.[22]

As we point out in the Introduction, reciprocity is a key variable in the theory behind MTL but there is disagreement about its use in practice. We find that the EU had larger reduction in tariffs on goods exported by trading partners that reduced their own tariffs by a greater amount. Reciprocity may magnify the stumbling block

[21] An additional robustness test is to exclude the export dummy and its interactions from the instrument list altogether. Naturally the first stage R-square for the PTA variable falls but the F-statistic of the excluded instruments is still significant at the 1 percent level since the price variables help to predict PTA goods. The second stage estimate for ϕ^{any0} in this case is also positive and significant at the 1 percent level and slightly higher than the baseline in column 2 of Table 2.

[22] Goldberg and Maggi (1999) report $1/\omega = 0.986$ (p. 1145) whereas Gawande and Bandyopadhyay (2000) report $1/(\omega - 1)$ (p. 147). In Karacaovali and Limão (2005), we relax the constraint of constant weights across sectors and show that those with higher share of employment and regional concentration receive higher tariff protection, but we also find that the restriction of constant weights (over industries and time) is reasonable for the EU.

effect, because smaller reductions in the EU will be reciprocated by smaller reductions in the trading partners. Since Limão (2006) also finds reciprocity to be a significant factor in the US multilateral tariff reductions during the UR, we expect that the stumbling block effect of the EU and the US had an indirect effect at least in the reciprocal tariff reductions between the two.

Our proxy for the change in MFN externality term, Δm, is generally positive as predicted but always insignificant. One explanation for this is that the reciprocity variable already accounts for this effect. Since those countries that free-ride will have small average tariff reductions, the EU will "reciprocate" with smaller tariff reductions of its own as well.

4.4 Quantification and Interpretation

The simplest interpretation of the coefficient on the PTA variable is that it represents how much the MFN tariff for PTA goods increased relative to the non-PTA goods. Its value is 1.5 percentage points for goods exported under any PTA and about 2.2 for every PTA. Since the reduction for non-PTA goods was 3.4 percentage points, the magnitude of the stumbling block effect is not trivial.

We can quantify the tariff effect in terms of price changes to assess its economic importance. This type of quantification is important because one key concern with PTAs is that they have an impact on other countries by affecting the prices received by excluded countries, for example by causing higher MFN tariffs, as our model shows. Moreover, there is considerable evidence of imperfect pass-through both from exchange rate changes (c.f. Goldberg and Knetter 1997) and from tariff changes (c.f. Finger 1976 and Feenstra 1989).

In the context of the average price effects generated by tariff changes during the UR, the stumbling block effect is not negligible. This is clearest from employing the ratio of the relative (domestic) price growth effects

$$\Delta \ln p^d_{pta} / \Delta \ln p^d_{mfn} = \Delta \ln(1 + t^{mp}) / \Delta \ln(1 + t^m) \approx \Delta t^{mp} / \Delta t^m = 1 + \phi/c$$

where c is the estimated average tariff change for non-PTA goods. This equality applies to a "benchmark" good with no changes in either the market access or the elasticity adjusted production/import ratio. If the stumbling block effect completely offsets the average price effect, then $1 + \phi/c = 0$ and if the price effect for the PTA goods were identical to the non-PTA goods, the statistic would be equal to 1. As we show in the working paper this statistic also measures the relative world price effects for goods where there is imperfect pass-through, that is $\Delta \ln p^w_{pta} / \Delta \ln p^w_{mfn} \approx 1 + \phi/c$, provided that the pass-through rate for PTA goods is similar to other goods.

The closer $1 + \phi/c$ is to zero the stronger the stumbling block effect. For example, a value of 0.5 indicates that a non-PTA country received only half of the export price increase from EU's MFN tariff changes in the UR by exporting a PTA good relative to a similar non-PTA good. An alternative interpretation in light of the theoretical model is that the export price for the PTA goods was half of what it would have been in the absence of EU's PTAs. At the bottom of tables 2 and 3, the row labeled $1+\phi/c$ provides the estimates of the price effects as well as their confidence intervals. The effect of any PTA is about 0.55 (column 1, Table 2) and it is not very sensitive to controlling for the exports of AFS or SGP (column 1, Table 3) or a positive preferential tariff (column 2, Table 3). The effect for goods exported by every PTA is stronger: 0.39 (column 2, Table 2).

An interesting question is whether our estimates carry any information about the unobserved counterfactual of what the average EU tariff would have been in the absence of any PTAs. A strict interpretation of our estimates according to the theoretical model is that MFN tariffs are 1.5 percentage points higher for PTA goods relative to a counterfactual without PTAs and, since PTA goods represent a large share of our sample, the average over *all goods* is 1.4 percentage points. However, there has also been debate on whether the PTAs pursued by the US and the EU increased or decreased the completion probability of the Uruguay Round (UR). Since multilateral trade rounds are too infrequent, whether PTAs increase or decrease the probability of a round cannot be answered econometrically but we can provide some bounds for our results. For instance, suppose that in the absence of PTAs the UR would not have been completed, and the mere existence of PTAs assured its completion. In this case the *total* stumbling block effect, which incorporates possible changes in the probability of a round, can be shown to be equal to 1.2 percentage points (Karacaovali and Limão 2005). Thus, even under this extreme assumption, these EU PTAs are a stumbling block and would have been so unless the average reduction in tariffs by other countries had been almost six times larger than what we actually observed. If, on the other hand, the probability of completing the round without PTAs were nearly one and with PTAs it became nearly zero, then the total stumbling block effect would be at least 1.4 percentage points.

Finally, examining the importance of the PTA variable differently, the explained amount of variation in the tariffs across goods which can be attributed to it is significantly higher than the rest of the variables.

4.5 Robustness Analysis

We now test whether the stumbling block results given in Table 2 are robust along various dimensions such as the choice of controls, instruments, measurement error in the PTA variable, and the sample. We summarize the results in Table 5, where the first column repeats the basic information from Table 2 for comparison. The first two rows correspond to robustness tests of specification 1 in Table 2, that is ϕ^{any0} and $1+\phi^{any0}/c$—the quantification discussed in the previous section—respectively. The third and fourth rows present similar statistics when we account for the effect of every PTA, analogously to column 2 of Table 2.

We did not derive the reciprocity variable directly from our theoretical model. Thus, we ask whether the main results are affected if we drop it and follow the structural equation more closely. The second column of Table 5 shows that the results are unchanged.

[TABLE 5 HERE]

As we describe in Section 4.1 and show in the appendix (Section A.3), there is a potential for misclassification of new PTA goods that may generate attenuation bias. An instrument that is correlated with the new PTA goods in I_{it} but not the existing ones reduces or eliminates this bias and in fact we showed that the OLS estimates were slightly lower than the IV. However, the IV specifications we analyzed thus far are unlikely to fully solve this problem. A better alternative may be to use as instruments those goods that we know are "sure switches" to PTA goods. This is the case for the goods exported by the former communist countries in the CEC group, because this PTA was completely unexpected at the time of the Tokyo Round and in place by the UR. When we use the CEC exports as an instrument for I^{any0}, the stumbling block estimate is 0.036, as shown in column 3 of Table 5. This value is significantly higher than the OLS estimate of 0.013 and also the baseline IV estimate of 0.015, which supports our earlier argument that the misclassification of new PTA goods will if anything lead to an underestimate of the true stumbling block effect. The price effect estimate is now 0.36, which is about 30 percentage points stronger than the OLS estimate.

We do not expect our estimates to be biased due to omitted variables because, even if they are correlated with the included regressors, we instrument, test, and confirm the orthogonality of the excluded instruments relative to the error term. Nevertheless, we want to explicitly address the effect of including the initial tariff rate in the estimation. The average initial MFN tariff for PTA goods in our sample is 7.6 percent, whereas it is 12.8 percent for the non-PTA goods. Although in the UR no explicit formula was followed such that higher tariffs would be cut by more than the lower ones, it is certainly a possible outcome and may lead us to find bigger cuts in the non-PTA goods. When we add the initial tariff level as an endogenous regressor, we find that its coefficient is typically negative, so products with higher initial tariffs had slightly bigger cuts, but it is not always statistically significant. Moreover, the initial tariff does not affect the sign, magnitude or significance of the basic stumbling block effect. As shown in column 4 of Table 5, the relative growth effect evaluated at the average initial tariff is at least as large as the ones found in Table 2. Since the main results are not sensitive to the inclusion of the initial tariff and, according to the Schwarz criterion, the specification without it is preferred, we choose to focus on the latter, which also follows our theoretical model more closely.

There is considerable variation across PTAs in the set of goods each exports duty-free to the EU. This fraction in our sample ranges from 0.2 for GSPL to 0.9 for countries that acceded the EU and the median for the PTAs with non-acceding countries, which are our main focus, is 0.58. The fraction of PTA goods exported by *any* PTA is 0.94 and two-thirds of the ISIC three-digit industries contain both PTA and non-PTA goods. The fraction of such goods exported by every PTA is 0.13 spread over three quarters of the industries.

Despite the variation just described, the food products industry, ISIC 311, contains approximately half of all the products in our sample that do not enter the EU duty-free through *any* preferential agreements. Although this category does not include primary agricultural products (it includes processing of food related products), it does share one important characteristic with agriculture: high protection. To the extent that this feature is time-invariant, then it is immediately addressed by the fact that we estimate the equations in differences. Moreover, the initial tariff, which is on average higher for ISIC 311 than for other industries, does not seem to be biasing the results, because, as we have just shown, its inclusion does not affect the results significantly. To investigate whether the stumbling block effect is merely driven by a cross-industry difference in the average tariff cut between ISIC 311 and other industries, we re-estimate the model by dropping the observations in 311. The estimates in column 5 of Table 5 are qualitatively similar to those in Table 2 in terms of the signs and significance of the coefficients.

When we exclude products in ISIC 311 the point estimate for ϕ^{any0} is smaller than the baseline. One reason for the lower estimate for ϕ^{any0} when ISIC 311 is excluded is that the attenuation bias from the measurement error is severely aggravated. The reason is simple, goods that were not PTA goods in the latter period, were also not PTA goods earlier (since preferences and the range of goods exported have increased) so they were correctly classified as no switches. By dropping a large number of these non-PTA goods we increase the fraction of PTA goods and therefore the proportion of potentially misclassified goods. We confirm this in column 6 where we apply the alternative instruments of sure switches described above to the subsample without ISIC 311. The point estimate for ϕ^{any0} is now 0.028, which is statistically indistinguishable from the earlier one in column 3 of 0.036, as is the price effect, 0.39.

For the products in the industries mentioned above, we were able to identify potential problems and test the robustness of the results to excluding them. However, there may exist other unobserved industry characteristics driving tariff changes and they may occur in industries that are defined differently, e.g. by the Harmonized Standard (HS) tariff code. Controlling for unobserved industry characteristics is particularly important when one employs a different empirical approach, which does not follow a structural equation as closely as we do. This is the case in Limão (2006) who controls for "industries" defined by the HS classification when studying the US. Nonetheless, one may ask whether our baseline results are also robust to controlling for such unobserved industry characteristics. Thus, in columns 7 and 8 of Table 5 we include industry effects—broadly defined here by the HS-1 digit classifications. We still obtain a positive and significant stumbling block effect. Since including these HS dummies implies that there is less variation available to identify the effect, it is important to address the measurement error in the PTA variable in order to minimize any attenuation bias. Column 8 does this by employing the sure switch instruments. When we do so, we find that the results are both qualitatively and quantitatively close to the analogous ones in column 3.[23]

[23] Limão (2006) controls for industries defined at the finer HS-2 digit level. The EU tariff changes do not permit us to control for such disaggregate industries because the share of the between variation at the HS 2 level is about two times higher in the EU than in the US tariff changes. This also implies that for the EU, the HS-1 effects account for a reasonable share of the tariff change variation relative to the other variables we include, as evidenced by the statistical significance of

It is possible that goods in which the EU has an NTB on all partners share other properties that affect its MFN tariff change. As we previously reported, alternative over-identifying restrictions tests reject this hypothesis. But we can go further and test if this affects our results by dropping all instruments that include the NTB variable, D^{ntball}. The results, shown in the last column, are the same as in the baseline: ϕ^{any0} is 0.014 instead of 0.015 and ϕ^{evy0} is 0.009 instead of 0.007.

The Hansen-Sargan test of over-identifying restrictions cannot reject the null hypothesis that the excluded instruments are uncorrelated with the second stage error, and correctly excluded in any of the specifications, with a p-value of 0.5 or higher in columns 1 through 6 and a p-value of 0.22 or higher in columns 7 through 9. The same is true for the difference in Sargan test for subsets of instruments.

The identity of a PTA good may be correlated with other characteristics that affect tariff changes and this was a key motive for instrumenting the PTA variable and extensively testing the validity of the instruments. However, we go further and now summarize various tests directly related to the identity of the PTA goods. In addition to the one described for ISIC 311, the baseline result is also robust to excluding the export variable from the instrument list and using price changes to predict which goods are exported (see footnote 21). The large share of goods exported by at least one PTA, I^{any0}, is explained by several factors, including the very large and diverse number of PTA partners (over 150) and for some, the low trade costs due to proximity to the EU (e.g. EFTA, MED, CEC). As we see in Table A1 , the fraction of HS-8 lines with a duty-free preference varies across PTAs from 70% (GSP) to 94% (EFTA). EFTA exports in 90% of the HS-8 lines and receives preferences in most, so if we recalculate I^{any0} excluding EFTA, its mean falls to 0.78. Given this, we can ask if the baseline result is driven by EFTA by using the redefined I^{any0} variable and including the EFTA variable separately in that regression (adjusting the instruments accordingly). The results are very similar to the baseline, with $\phi^{any0} = 0.016$ and a coefficient on the EFTA variable positive and significant, similar to the one found for the individual PTA results in column 3 of Table 3. Therefore, the various robustness tests provide additional evidence that the identity of the PTA goods does not significantly bias the IV results.

5 Conclusion

We analyze the effects of PTAs on multilateral trade liberalization—a controversial issue where the evidence has been scarce. The model we develop captures key features of the current trading system and provides a rich set of predictions regarding the impact of PTAs on MTL. We derive structural equations of protection that guide the estimation. Using detailed tariff data for the EU during the last two multilateral rounds, we find evidence that its PTAs slowed down MTL. As the model predicts, this occurred only in products with a zero preferential tariff

these HS-1 effects in the regression and the improvement in the Schwarz criterion.

and was not present in agreements with a common external tariff and transfers, i.e. accessions to the EU in the 1980's and 1995. Our model also incorporates domestic political economy motives for tariffs and we find a negative relation between import penetration and tariff levels working through the extra weight that governments place on producer surplus. We also find evidence of multilateral reciprocity.

In the absence of its PTAs, the EU would have lowered its MFN tariff on PTA products by an additional 1.5 percentage points. Since the average reduction for non-PTA products is almost twice as high, the average price effect due to the EU's multilateral tariff changes is 50-60 percent for PTA goods relative to other goods. We also discuss how this wedge between PTA and non-PTA products provides an estimate of the effect of PTAs on the expected average reductions in *all* products relative to a situation where the EU has no PTAs and show that the effect is at least 1.2 percentage points.

We find no evidence of a stumbling block effect for countries that acceded the EU in the 1980's and 1995. But we do find evidence for all other preferential agreements of the EU in place at the time of the Uruguay Round. Since the EU continues to expand both types of preferences, we think that both results are important. Moreover, the lack of evidence on preferences promoting multilateral liberalization for the EU (or the US) suggests that we should be concerned about a "clash of liberalizations". Similar work that carefully establishes causation is required for other countries. Even if the EU and the US turn out to be the exceptions, this concern has to be still addressed because their share of world trade implies that their PTAs have a potentially large impact on non-members. These estimates suggest that the stumbling block effect may be even worse for the Doha round, which is currently under negotiation. The motive is simple; after the UR, preferences for existing and new PTAs were greatly expanded, partly as a way to counter the preference erosion generated by the lower MFN tariffs.

The inevitable final question is what, if anything, can be done to minimize this clash. The current enthusiasm for PTAs means that prohibiting them is not feasible and we have not shown that doing so would necessarily be optimal either. However, there may be ways to grant preferential treatment that do not slow down MTL. Recall that, according to the model, the effect of PTAs on MTL only occurs when the preferential tariff is zero and cannot be further lowered. From this perspective, the answer is simple: remove the non-negativity constraint on preferential tariffs and allow import subsidies. Limão and Olarreaga (2006) estimate that the additional MTL thus permitted generates a Pareto improvement for the three groups of countries: non-members, and preference granting and receiving countries. This or other proposals that target the source of the problem and take into account the effects on these three groups of countries are the most likely to be accepted by them and minimize any further clash of liberalizations.

A Appendix

A.1 Tariff Expressions

A.1.1 No Common External Tariff

Equation (4): Consider good $i = 1$, which is imported by L from L^* and S, and its symmetric counterpart $i = 2$, exported to L^*. Using equation (2) we simplify the first-order-condition (FOC) for an interior solution to equation (3) and obtain

$$\begin{aligned} G_{\tau_1} + G_{\tau_2^*} + G_{\pi_1} &= -[M_1^s + M_1^{*} + \tau_1(p_1' - 1)M_1^{*\prime}] - p_1' M_1 - (p_2'^{*} - 1)M_2 + (\omega_1 - 1)p_1' X_1 \\ &= -[\tau_1(p_1' - 1)M_1^{*\prime}] - (p_1' - 1)M_1 - (p_2'^{*} - 1)M_2 + (\omega_1 - 1)p_1' X_1 \qquad (16) \end{aligned}$$

where $p' \equiv \partial p/\partial \tau$, $M^{*\prime} \equiv \partial M^*/\partial p^*$ and we use the market clearing condition in equation (1) for $i = 1$. Equating to zero and solving for τ_1, we obtain

$$\tau_1^m = \frac{(\omega_1 - 1)p_1' X_1 - (p_1' - 1)M_1 - (p_2'^{*} - 1)M_2}{(p_1' - 1)M_1^{*\prime}} \qquad (17)$$

To derive equation (4) in the text, we do the following. The symmetry implies that $p_1' = p_2'^{*}$, $M_1 = M_2^*$, and $M_1^s = M_2^{s^*}$. Market clearing implies that $M_2^* + M_2^{s^*} = -M_2$, where we replace the symmetry conditions to obtain $M_2 = -(M_1 + M_1^s)$. Finally, we divide equation (17) by p_1^*, use $p' = M^{*\prime}/[M^{*\prime} + M']$ (from implicit differentiation of equation (1) and $M_1^{\prime s} = 0$) and employ the following elasticity definitions: $\varepsilon \equiv -M' p^*/M$ and $\varepsilon^* \equiv [\partial(M^* + M^s)/\partial p^*] \times [p^*/(M^* + M^s)] = [\partial M^*/\partial p^*] \times [p^*/(M^* + M^s)]$.

Equation (9): Using equations (1) and (2) and the fact that net exports of S are X_1^s in equilibrium, the FOC for τ^{mp} for an interior solution to equation (8) is

$$\begin{aligned} G_{e^s}\frac{\partial e^p}{\partial \tau_1} + G_{\tau_1} + G_{\tau_2^*} &= G_{e^s}\frac{\partial e^p}{\partial \tau_1} - [M_1^{*} + M_1 + \tau_1(p_1' - 1)M_1^{*\prime}] - (p_1' - 1)M_1 - (p_2'^{*} - 1)M_2 + (\omega_1 - 1)p_1' X_1 \\ &= G_{e^s}\frac{\partial e^p}{\partial \tau_1} - X_1^s - \tau_1(p_1' - 1)M_1^{*\prime} - (p_1' - 1)M_1 - (p_2'^{*} - 1)M_2 + (\omega_1 - 1)p_1' X_1 \qquad (18) \end{aligned}$$

Equating to zero and solving for τ_1 we obtain

$$\tau_1^{mp} = \frac{(\omega_1 - 1)p_1' X_1 - (p_1' - 1)M_1 - (p_2'^{*} - 1)M_2 + [G_{e^s}\frac{\partial e^p}{\partial \tau_1} - X_1^s]}{(p_1' - 1)M_1^{*\prime}} \qquad (19)$$

To obtain equation (9), we apply the symmetry conditions described in the previous derivation, divide equation (19) by p_1^*, and employ the same elasticity definitions.

A.1.2 Common External Tariff and Direct Transfers

We show that a stumbling block effect is absent in the presence of direct transfers by focusing on the simpler case when S exports only good 1 and $\omega_1^s = 1$. Allowing a transfer T from L to S under a PTA implies that the participation constraint for S is now $G^s(\boldsymbol{\pi}^p, \boldsymbol{\tau}, e^p) + T \geq G^s(\boldsymbol{\pi}, \boldsymbol{\tau}, e^s = 0)$. Thus the equilibrium level of e^s for a given preference, $\tau_1 - \pi_1$, and transfer, T is

$$e^{pcu} = T + (\tau_1 - \pi_1)(-M_1^s) \tag{20}$$

The large countries maximize their joint objective net of transfers made to the regional partner.

$$\{\tau^{mcu}, \pi^{pcu}, T^{cu}\} \equiv \arg\max_{\tau,\pi,T}\{G(\boldsymbol{\tau} = \boldsymbol{\tau}^*, \boldsymbol{\pi}, \boldsymbol{\tau}^*, e^s = 0, .) - T : \pi \leq \tau; e^s = e^{pcu}\} \tag{21}$$

The FOCs for a good obtaining a preference under this program of CET with transfers are

$$G_\tau + G_{\tau^*} + G_{e^s}\frac{\partial e^{pcu}}{\partial \tau} \leq 0; \qquad G_\pi + G_{e^s}\frac{\partial e^{pcu}}{\partial \pi} \leq 0 \; ; \qquad G_T + G_{e^s}\frac{\partial e^{pcu}}{\partial T} \leq 0 \tag{22}$$

where we recall that $G(.)$ is defined by equation (2). Evaluating the FOC of the MFN tariff in the absence of preferences, i.e. equation (16), at the level of the MFN tariff under the CU in equation (22), we obtain

$$\left[G_{\tau_1} + G_{\tau_2^*} + G_{\pi_1}\right]_{G_\tau + G_{\tau^*} + G_{e^s}\frac{\partial e^{pcu}}{\partial \tau} = 0} = G_{\pi_1} - G_{e^s}\frac{\partial e^{pcu}}{\partial \tau_1} = G_{\pi_1} + G_{e^s}\frac{\partial e^{pcu}}{\partial \pi_1} = 0 \tag{23}$$

where the second equality uses equation (20). In order to establish the last equality we must show that the FOC for π in equation (22) is zero. Suppose it is not, such that $G_\pi + G_{e^s}\frac{\partial e^{pcu}}{\partial \pi} = -M^s + M^s G_{e^s} = (-1 + G_{e^s})M^s < 0$ when evaluated at $\tau^{mcu} = \tau^m$. This implies that L would gain by further lowering π but it can only do so until it is zero. However, $G_T + G_{e^s}\frac{\partial e^{pcu}}{\partial T} = -1 + G_{e^s}$, which is positive if $G_\pi + G_{e^s}\frac{\partial e^{pcu}}{\partial \pi} < 0$. Thus $\tau^{mcu} = \tau^m$, $T = 0$ and $\pi = 0$ is not a solution. T must be increased until $-1 + G_{e^s} = 0$, which implies that $G_\pi + G_{e^s}\frac{\partial e^{pcu}}{\partial \pi} = 0$. Therefore we obtain $\tau^{mcu} = \tau^m$, $T^{cu} > 0$ and $\pi = \pi^{pcu} < \tau^{mcu}$.

A.2 Stumbling Block Effects with Multiple PTA Goods

In Section 3.2.2 we claim that when S exports multiple goods, the stumbling effect is present in the set of goods where the preferential tariff is zero but not in those where it is positive and that this depends on the existence of different political weights in S. We show this for the case when L imports goods 1 and 3 from L^* and S, and the symmetric imports for L^* are $i = 2, 4$. The equilibrium level of e^p in equation (7) is now

$$e^p = \sum_{i=1,3} (\tau_i - \pi_i)\omega_i^s X_i^s \tag{24}$$

The FOC for the two MFN and two preferential tariffs arising from the program in equation (8) are

$$\begin{aligned} G_{e^s}\frac{\partial e^p}{\partial \tau_1} + G_{\tau_1} + G_{\tau_2^*} &\leq 0 \;\; ; \;\; G_{e^s}\frac{\partial e^p}{\partial \tau_3} + G_{\tau_3} + G_{\tau_4^*} \leq 0 \\ G_{e^s}\frac{\partial e^p}{\partial \pi_1} + G_{\pi_1} &\leq 0 \;\; ; \;\; G_{e^s}\frac{\partial e^p}{\partial \pi_3} + G_{\pi_3} \leq 0 \end{aligned}$$

The objective is to show that in an equilibrium where both goods receive a preference but there is room for increasing the preferential margin in good 3, i.e. $\pi_3 \in (0, \tau_3)$, then if $\omega_1^s > \omega_3^s$, we have $\pi_1 = 0$, $\tau_1^{mp} > \tau_1^m$, and $\tau_3^{mp} = \tau_3^m$. By construction, the solution for π_3 is an interior one in this equilibrium and therefore its FOC holds with equality and $G_{e^s}\frac{\partial e^p}{\partial \pi_3} = -G_{\pi_3}$. Substituting this in the FOC for π_1 we obtain

$$\left(\frac{-G_{\pi_3}}{\partial e^p/\partial \pi_3}\right)\frac{\partial e^p}{\partial \pi_1} + G_{\pi_1} = -X_3^s\frac{\omega_1^s X_1^s}{\omega_3^s X_3^s} + X_1^s = \left(1 - \frac{\omega_1^s}{\omega_3^s}\right)X_1^s < 0$$

where the first equality uses the definition of G and equation (24), and the final inequality is due to $\omega_1^s > \omega_3^s$. The expression for the MFN tariff in good i is still given by equation (9) and, as we show in the text, the stumbling block effect is present in that good iff $-(G_{e^s}\partial e^p/\partial \pi_i + G_{\pi i}) > 0$. The LHS terms are simply the FOC for π_i and in this equilibrium we know they are equal to zero for good 3 since $\pi_3 > 0$ and so $\tau_3^{mp} = \tau_3^m$. But this inequality does hold for good 1 and so $\tau_1^{mp} > \tau_1^m$.

If the optimal preferential tariff implied by the FOC is $\pi_3 > \tau_3^m$, then L must follow WTO rules and set $\pi_3 = \tau_3^m$. In this case we are effectively back to the situation with one PTA good so it is not sufficient for $\omega_1^s > \omega_3^s$ in order to obtain $\pi_1 = 0$. But provided that in equilibrium we do have $\pi_1 = 0$ (e.g. because ω_1^s is sufficiently large relative to ω_3^s) then we continue to have the same prediction as before. That is, in the duty-free PTA goods, there is a stumbling block, $\tau_1^{mp} > \tau_1^m$, and in the goods that do not receive a preference, the tariff expression is the same as in the case with no PTAs, so $\tau_3^{mp} = \tau_3^m$.

A.3 Interpretation of Estimates for Continuing PTAs

By differencing equation (12) and assuming time-invariant political weights, we obtain the general econometric model in error form

$$\Delta t_i = \phi_{it}(I_{it} - I_{it-1}) + I_{it-1}\Delta\phi_{it} + \beta\Delta x_i + \Delta m_i + \Delta u_i \tag{25}$$

where $E(\Delta u_{it}|I_{it}, I_{it-1}, \Delta x_{it}, \Delta m_{it}) = 0$. For observations that are not PTA goods at $t-1$ we obtain equation (14). EU preferences in PTAs that existed by the Tokyo Round expanded between the two rounds. Moreover, the sets of products that any given country can export have generally increased and if a product was exported in the recent past, the PTA government is likely to expect to export it in the immediate future. Both facts strongly suggest that $I_{it-1} = 1$ implies $I_{it} = 1$ and since both are binary variables, we have $I_{it-1} = I_{it-1}I_{it}$. This is obvious from the matrix below that shows all the possible combinations for a given product's status and clarifies its practical implication: ruling out the possibility in the last row where a good's status is initially 1 and then changes to 0.

I_{it}	I_{it-1}	$I_{it}I_{it-1}$	$I_{it} - I_{it-1}$
0	0	0	0
1	0	0	1
1	1	1	0
0	*1*	*0*	*−1*

Replacing $I_{it-1} = I_{it-1}I_{it}$ in equation (25) and factoring out I_{it} we obtain

$$\Delta t_i = ((1 - I_{it-1})\phi_{it} + I_{it-1}\Delta\phi_{it})I_{it} + \beta\Delta x_i + \Delta m_i + \Delta u_i \tag{26}$$

So an OLS estimate of the parameter on I_{it} for continuing agreements, call it $\tilde{\phi}$, can be interpreted as the average over all goods of $(1 - I_{it-1})\phi_{it} + I_{it-1}\Delta\phi_{it}$. This is a weighted average of the current stumbling block for goods with new preferences and the change in the stumbling block of the goods with continuing preferences, as we note in the text. Thus $\tilde{\phi}$ is the average *change* in the stumbling block between rounds, i.e. how much smaller tariff reductions were.

For PTAs signed between rounds, $\tilde{\phi}$ reflects exactly the stumbling block effect *level* since in that case $I_{it-1} = 0$. For those PTAs that previously had some subset of goods exported under preferences, $\tilde{\phi}$ generally provides a downward biased estimate of the level effect, ϕ_t. To see this, we simplify (26) to obtain

$$\Delta t_i = (\phi_{it} - I_{it-1}\phi_{it-1})I_{it} + \beta\Delta x_i + \Delta m_i + \Delta u_i \tag{27}$$

Recall from the model that $\phi_{it-1} \geq 0$, so generally the OLS coefficient on I_{it} will be a downward biased estimate of ϕ_t. This occurs because we should use only true switches to identify ϕ_{it}. But for some agreements we can not observe which subset of I_t corresponds to true switches (no product data on I_{it-1} for agreements in place before Tokyo Round) so we misclassify some goods with $I_t = 1$ as switches even though they had no change in status.

Equation (27) also indicates how the bias can be ameliorated with IV. When the instrument is correlated with I_t but not with I_{t-1} the term $I_{it-1}\phi_{it-1}$ can be left in the error term and we obtain a consistent estimate of ϕ_t. The programs with new preferences and/or exports, e.g. CEC, have $I_{t-1} = 0$ and so are good candidates for instruments.

A.4 Data

[TABLES A1, A2 and A3 HERE]

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TABLE 1. EU Tariff Levels and Changes in the Uruguay Round: Distribution by Industry

ISIC		Before UR		After UR		Change		
Code	Sector Name	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std.Dev.	Coef.Variation
311	Food products	0.16	0.09	0.11	0.07	0.05	0.03	0.55
313	Beverages	0.11	0.04	0.07	0.02	0.04	0.02	0.51
314	Tobacco	0.42	0.20	0.25	0.12	0.17	0.09	0.52
321	Textiles	0.10	0.03	0.07	0.02	0.03	0.02	0.77
322	Wearing apparel except footwear	0.13	0.03	0.11	0.03	0.02	0.01	0.47
323	Leather products	0.05	0.02	0.03	0.02	0.02	0.01	0.56
324	Footwear except rubber or plastic	0.10	0.05	0.09	0.04	0.01	0.01	1.50
331	Wood products except furniture	0.06	0.02	0.02	0.03	0.04	0.01	0.31
332	Furniture except metal	0.06	0.01	0.01	0.02	0.05	0.01	0.28
341	Paper and products	0.09	0.02	0.04	0.02	0.04	0.02	0.39
342	Printing and publishing	0.09	0.03	0.05	0.02	0.05	0.02	0.36
351	Industrial chemicals	0.08	0.03	0.06	0.02	0.03	0.03	1.08
352	Other chemicals	0.07	0.02	0.03	0.03	0.04	0.03	0.81
353	Petroleum refineries	0.05	0.02	0.03	0.02	0.02	0.01	0.53
354	Miscellaneous petroleum and coal products	0.04	0.02	0.03	0.03	0.01	0.01	0.85
355	Rubber products	0.05	0.02	0.03	0.02	0.02	0.01	0.53
356	Plastic products	0.11	0.05	0.08	0.05	0.03	0.02	0.63
361	Pottery china earthenware	0.08	0.03	0.06	0.03	0.02	0.01	0.58
362	Glass and products	0.07	0.03	0.05	0.03	0.03	0.01	0.46
369	Other non-metallic mineral products	0.05	0.02	0.02	0.02	0.02	0.01	0.38
371	Iron and steel	0.06	0.02	0.00	0.01	0.05	0.02	0.39
372	Non-ferrous metals	0.06	0.02	0.04	0.03	0.02	0.01	0.67
381	Fabricated metal products	0.06	0.02	0.03	0.02	0.03	0.01	0.50
382	Machinery except electrical	0.05	0.01	0.02	0.01	0.03	0.01	0.48
383	Machinery electric	0.06	0.03	0.03	0.02	0.03	0.02	0.55
384	Transport equipment	0.08	0.05	0.05	0.05	0.02	0.02	0.75
385	Professional and scientific equipment	0.06	0.01	0.03	0.02	0.03	0.01	0.41
390	Other manufactured products	0.06	0.02	0.03	0.02	0.03	0.02	0.50
	Total	0.08	0.05	0.05	0.04	0.03	0.02	0.73

Advalorem tariff rates are reported. The total number of observations in our sample is equal to 6294. Products with zero initial tariff rates are excluded from the sample as explained in the text.

TABLE 2. Impact of EU Preferences on MFN Tariffs: Main Predictions

	IV-GMM			OLS
	(1)	(2)	(3)	(4)
$I^{any0\dagger}$	0.015***	0.015***	0.011***	0.013***
(ϕ^{any0}>0)	(0.003)	(0.003)	(0.002)	(0.001)
$I^{evy0\dagger}$		0.007***		
(ϕ^{evy0}>0)		(0.002)		
$I^{hiexp\dagger}$			0.005***	
(ϕ^{hiexp}>0)			(0.002)	
$R^{\dagger}$	0.006*	0.005*	0.006**	0.010***
(ρ>0)	(0.003)	(0.003)	(0.003)	(0.002)
$\Delta x^{\dagger}$	0.004*	0.003**	0.003**	0.000
(β>0)	(0.002)	(0.001)	(0.001)	(0.000)
P	0.001	0.000	0.001	0.001*
(μ>0)	(0.001)	(0.000)	(0.001)	(0.000)
Constant	-0.034***	-0.034***	-0.036***	-0.038***
(c)	(0.005)	(0.004)	(0.003)	(0.002)
Observations	6294	6294	6294	6294
Hansen's J p-val [a]	0.506	0.567	0.569	n/a
C-stat p-val [b]	0.729	0.562	0.553[e]	n/a
Endogeneity p-val [c]	0.517	0.248	0.270	n/a
Heterosked. p-val [d]	0.000	0.000	0.000	0.000
$1+\phi/c$ [e]	0.55	0.58	0.69	0.66
	(.41,.68)	(.47,.69)	(.56,.81)	(.60, .71)
$1+(\phi+\phi^{\text{evy, hi}})/c$ [e]	n/a	0.39 (.21,.56)	0.56 [f] (.44, .68)	n/a
Schwarz info.	-7.58	-7.60	-7.61	-7.68

Robust standard errors clustered at the 3-digit ISIC level in parentheses. * significant at 10%; ** significant at 5%; *** significant at 1%. The model's parameter and sign predictions are indicated in parentheses below each variable. IV-GMM obtained from an instrumental variable efficient generalized method of moments estimator. OLS obtained from Cragg's heteroskedastic ordinary least squares estimator. See Table 4 with the first stage regressions for the list of instruments.

(a) Hansen-Sargan test of over-identifying restrictions. Probability value for the null hypothesis that the excluded instruments are uncorrelated with the second stage error term, and correctly excluded from the estimated equation.

(b) Difference in Sargan (C) statistic. Probability value for the rejection of the null hypothesis of exogeneity of a subset of f instruments marked with "‡" in Table 4. The first stage regression results for specification (3) are similar to those of (2) in Table 4 but employ the export indicator for high exports, D^{hiexp} , instead of D^{evy}. The statistic used is C = J^r- J^u with a chi-square distribution of f degrees of freedom where J^r represents the minimized value of the GMM objective function for the restricted and efficient regression (with all over-identifying restrictions) and J^u the value for the unrestricted, inefficient but consistent regression without the questionable instruments.

(c) Endogeneity test of the variables marked with a "†" based on the C statistic. Probability value at which we reject the consistency and efficiency of OLS.

(d) Pagan-Hall heteroskedasticity test. Probability value for the null hypothesis that the disturbance is homoskedastic.

(e) Confidence intervals calculated using the delta method.

(f) The value for the combined effect of I^{any0} and I^{hiexp}.

TABLE 3. Impact of EU Preferences on MFN Tariffs: Auxiliary Predictions

	IV-GMM		
	(1)	(2)	(3)
$I^{any0\dagger}$	0.015***		
(ϕ^{any0}>0)	(0.003)		
I^{afs}	0.002		
(ϕ^{afs}=0)	(0.002)		
I^{spg}	-0.001		
(ϕ^{spg}=0)	(0.001)		
$I^{any\dagger}$		0.013***	
(ϕ^{any}>0)		(0.003)	
$I^{pos\dagger}$		-0.026	
($\phi^{pos}+\phi^{any}$=0)		(0.032)	
$I^{gsp\ \dagger\text{f}}$			0.003***
(ϕ^{gsp}>0)			(0.001)
$I^{gspl\ \dagger\text{f}}$			0.003***
(ϕ^{gspl}>0)			(0.001)
$I^{acp\dagger\text{f}}$			0.000
(ϕ^{acp}>0)			(0.001)
$I^{eftx\ \dagger\text{f}}$			0.002***
(ϕ^{eftx}>0)			(0.001)
$I^{med\dagger\text{f}}$			0.002***
(ϕ^{med}>0)			(0.001)
$I^{cec\ \dagger\text{f}}$			0.002***
(ϕ^{cec}>0)			(0.001)
$\Delta x^{\dagger}$	0.003**	0.004*	0.003**
(β >0)	(0.002)	(0.002)	(0.001)
$R^{\dagger}$	0.006*	0.008*	0.008***
(ρ>0)	(0.004)	(0.004)	(0.003)
P	0.001	0.001	-0.000
(μ>0)	(0.001)	(0.001)	(0.000)
Constant	-0.035***	-0.030***	-0.028***
(c)	(0.004)	(0.005)	(0.003)
Observations	6294	6294	6294
Hansen's J p-val [a]	0.495	0.559	0.159
C-stat p-val [b]	0.722	0.446	0.114
Endogeneity p-val [c]	0.525	0.399	0.210
Heterosked. p-val [d]	0.000	0.000	0.000
$1+\phi/c$ [e]	0.57 (.47,.66)	0.57 (.42,.72)	n/a
$1+\sum\phi^{pta}/c$ [e]	n/a	n/a	0.53 [f] (.37, .69)
$\phi^{pos}+\phi^{any}$=0 p-val	n/a	0.68 (accept)	n/a
Schwarz info.	-7.60	-7.56	-7.64

All the notes to Table 2 apply. For the specification in the last column, we employ export indicators for each PTA, D^{pta}, instead of D^{any}. The following coefficient restrictions are employed in the last column, $\phi^{\text{gsp}}=\phi^{\text{gspl}}$, $\phi^{\text{eftx}}=\phi^{\text{cec}}$, based on a test failing to reject their equality. (a)-(e): See notes to Table 2. (f) Calculated for a product exported under every program. The other values, with confidence intervals in parenthesis, are GSP and GSPL: 0.88 (.82, .94), ACP: 0.99 (.93, 1.06), MED: 0.92 (.87, .97), EFTX and CEC: 0.93 (.88, .97).

TABLE 4. First Stage Regressions for Main Predictions

	Table 2- Column 1			Table 2-Column 2			
	I^{any0}	Δx	R	I^{any0}	I^{evy0}	Δx	R
D^{anyexp} ‡	0.936***	-0.866***	0.021***	0.933***	-0.003	-0.774***	0.020***
	(0.023)	(0.217)	(0.007)	(0.023)	(0.020)	(0.215)	(0.007)
D^{evyexp} ‡				0.035***	0.801***	-0.755***	0.013***
				(0.007)	(0.006)	(0.063)	(0.002)
D^{ntball} ‡	0.008	-0.888**	-0.010	0.013	-0.014	-0.974**	-0.009
	(0.041)	(0.388)	(0.012)	(0.041)	(0.036)	(0.383)	(0.012)
D^{ntball}xD^{anyexp} ‡	-0.229***	0.506	0.010	-0.251***	0.000	0.319	0.009
	(0.042)	(0.392)	(0.012)	(0.041)	(0.036)	(0.389)	(0.012)
D^{ntball}xD^{evyexp} ‡				0.190***	-0.109***	1.497***	0.012*
				(0.026)	(0.022)	(0.240)	(0.007)
Δp_{9294}	-0.003	0.051	-0.004**	-0.002	-0.001	0.035	-0.004**
	(0.006)	(0.059)	(0.002)	(0.006)	(0.005)	(0.059)	(0.002)
$(\Delta p_{9294})^2$	-0.009***	-0.005	0.000	-0.007***	0.000	-0.024	0.001
	(0.003)	(0.028)	(0.001)	(0.003)	(0.003)	(0.027)	(0.001)
$(\Delta p_{9294})^3$	0.002*	-0.007	0.000	0.002*	0.000	-0.004	0.000
	(0.001)	(0.009)	(0.000)	(0.001)	(0.001)	(0.009)	(0.000)
D^{ntb} ‡	0.051***	1.014***	0.025***	0.049***	0.020***	1.072***	0.024***
	(0.006)	(0.060)	(0.002)	(0.006)	(0.006)	(0.059)	(0.002)
$\Delta scale$	0.189***	-0.497***	-0.032***	0.190***	0.003	-0.550***	-0.031***
	(0.009)	(0.082)	(0.002)	(0.009)	(0.008)	(0.081)	(0.002)
$(\Delta p_{9294})^{avg}$x$\Delta scale$ ‡	-0.362***	-0.917	0.102***	-0.285***	-0.017	-1.465*	0.119***
	(0.085)	(0.803)	(0.024)	(0.085)	(0.074)	(0.797)	(0.024)
R^{uni}	0.041**	-0.075	0.754***	0.042**	0.024	-0.058	0.754***
	(0.017)	(0.163)	(0.005)	(0.017)	(0.015)	(0.161)	(0.005)
P	0.039***	-0.274***	-0.004***	0.036***	0.003	-0.252***	-0.005***
	(0.005)	(0.046)	(0.001)	(0.005)	(0.004)	(0.045)	(0.001)
Constant	-0.039*	-1.129***	-0.275***	-0.042*	0.003	-1.084***	-0.275***
	(0.024)	(0.222)	(0.007)	(0.023)	(0.020)	(0.219)	(0.007)
Observations	6294	6294	6294	6294	6294	6294	6294
Adj. R^2	0.378	0.078	0.792	0.389	0.767	0.100	0.794
Shea's partial R^2	0.323	0.066	0.780	0.342	0.670	0.082	0.783
F-test excl. p-val [a]	0.000	0.000	0.000	0.000	0.000	0.000	0.000

Robust standard errors clustered at the 3-digit ISIC level in parentheses. * significant at 10%; ** significant at 5%; *** significant at 1%. The first stage regression results for specification (3) in Table 2 are similar to those of (2) shown above but employ the export indicator for high exports, D^{hiexp}, instead of D^{evy}.

‡: the subset of instruments further tested for exogeneity. The probability value for the difference in Sargan (C) statistics for these instruments are reported on the row labeled C-stat p-val. in Table 2.

(a) Probability value for the F–test of H_0: The excluded instruments are jointly insignificant. All F-statistics exceed 10.

TABLE 5. Robustness and Specification Analysis

Specification	Robustness procedure: / Test	IV-GMM (1)	IV-GMM (excl. reciproc.) (2)	IV-GMM (Sure switch) (3)	IV-GMM (Initial tariff) (4)	IV-GMM (Excl. 311) (5)	IV-GMM (Sure switch & excl. 311) (6)	IV-GMM ("Industry" effects) (7)	IV-GMM (Sure switch & "industry" effects) (8)	IV-GMM (Exclude ntb_all iv) (9)
Table 2 (1)	ϕ^{any0}	0.015	0.016	0.036	0.018	0.005	0.028	0.007	0.038	0.014
	(s.e.)	(0.003)	(0.003)	(0.006)[b]	(0.004)[c]	(0.002)	(0.008)	(0.003)	(0.007)	(0.006)
Table 2 (1)	$1+\phi/c$ [a]	0.55	0.56	0.35	0.36	0.79	0.39	0.77	0.36	0.48
	(95% CI)	(.41,.68)	(.41,.71)	(.23,.48)	(-.03, .70)[d]	(.66,.92)	(.25,.53)	(.61,.93)	(.24,.49)	(.13,.83)
Table 2 (2)	ϕ^{evy0}	0.007	0.007	0.006	0.009	0.007	0.006	0.004	0.002	0.009
	(s.e.)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.001)	(0.001)	(0.003)
Table 2 (2)	$1+(\phi^{evy0}+\phi^{any0})/c$	0.39	0.41	0.28	-0.11	0.52	0.35	0.58	0.35	0.24
	(95% CI)	(.21,.56)	(.22,.59)	(.15,.41)	(-.85,.63)	(.28,.78)	(.16,.54)	(.45,.72)	(.25,.45)	(-.26,.73)
Table 2 (1)	Schwarz info.	-7.58	-7.57	-7.59	-7.45	-7.68	-7.65	-7.72	-7.64	-7.49
	Observations	6294	6294	6294	6294	5875	5875	6294	6294	6294

IV-GMM obtained from an instrumental variable efficient generalized method of moments estimator.
The Hansen-Sargan test of over-identifying restrictions cannot reject the null hypothesis that the excluded instruments are uncorrelated with the second stage error term, and correctly excluded from the estimated equation in any of the specifications, with a p-value of 0.5 or higher in columns 1 through 6 and a p-value of 0.22 or higher in columns 7 through 9. The same is true for the difference in Sargan test for subsets of instruments.
Column (1) replicates the results from columns 1 and 2 of Table 2.
Column (2) drops the reciprocity variable and its instrument from the specifications in columns 1 and 2 of Table 2.
Column (3) employs instruments for the subset of products that are known to have switched status, i.e. uses D^{cec} instead of D^{anyexp} for I^{any0}.
Column (4) includes the initial tariff as an additional endogenous regressor.
Column (5) excludes all products in ISIC 311.
Column (6) excludes all products in ISIC 311 and employs instruments for subset of products that are known to have switched status, as in column 3.
Column (7) includes fixed effects for each of the Harmonized Standard 1-digit classifications.
Column (8) includes fixed effects for each of the Harmonized Standard 1-digit classifications and employs instruments for the subset of products that are known to have switched status, as in column 3..
Column (9) excludes all instruments that are constructed using the D^{ntball} variable.
(a) Measures the relative growth of world prices due to PTAs when the pass-through of tariffs to domestic prices is imperfect and similar for PTA and non-PTA goods, i.e. $\zeta^{PTA} \approx \zeta < 1$. Confidence intervals calculated using the delta method.
(b) We can reject the null that the coefficient is equal to 0.013, the OLS value in Table 2, with a p-value of =0.00
(c) The coefficient on the initial tariff variable (t_{t-1}) and the associated standard errors for specifications 1 and 2 are respectively: -0.195 (0.111) and -0.143 (0.068).
(d) Refers to the relative growth at the mean initial tariff i.e. $1+\phi/(c+0.0789^*\phi^{ini})$. When we account for the different average initial tariffs which are 0.0757 at I^{any0}=1 and 0.128 otherwise, and then calculate $(c+0.0757^*\phi^{ini}+\phi)/(c+0.128^*\phi^{ini})$, we obtain 0.26 (-.09, .61).

TABLE A1. Details on EU PTAs

Program	Recipients	Start Date	Type of Preference	Import share 1994	Share in '94 PTA imports	% of duty-free HS-8 lines	Notes / Non-trade Issues / Extensions	Source
ACP (African, Caribbean, and Pacific)	Over 70 countries, mostly former colonies of EU members	1976	Unilateral PTA	0.5%	1.9%	84%	Colonial ties major motivation. Financial and political cooperation; human rights play a role. (its earliest predecessor Yaoundé in 1963) In 1990 became the first development agreement to incorporate a human rights clause as a 'fundamental' part of cooperation (article 5).	Brown (2002), EU Commission 2004 [4]
CEC[1] (Central and East European)	Slovak Republic, Czech Republic, Poland, and Hungary	1992	Bilateral FTA	1.4%	6%	80%	Serves as a transition to full membership. Recipients committed to pass laws such as in intellectual property rights to conform with EU.	Busse (2000)
EFTX	Switzerland, Norway, Iceland, Liechtenstein.	1973-1974	Bilateral FTA	5.2%	22.3%	94%	EFTA members excluding AFS. Mainly industrial goods, excludes most agricultural products. Provides further for common competition rules, rules for state aid and government procurement, as well as harmonization of rules and standards for goods and services	EFTA [5] Secretariat 2004
GSP[2] (Generalized System of Preferences)	More than 100 developing countries	1971	Unilateral PTA	13%	57%	70%	Widest program; includes non-duty-free rates. Preferential rates vary according to competitiveness of the recipient countries. In 1981 the tariff quotas that were initially (in 1971) applied on 'sensitive products' were largely reduced.	Atkinson (1999), EU Commission 2004 [4]
GSPL (GSP for least developed)	About 50 of the poorest nations in the world	1971	Unilateral PTA	0.3%	1.2%	77%	Objective: Improving access to global markets for agricultural and industrial goods and services. It's a special arrangement for LDCs under the GSP program.	European Commission 2004
MED[3] (Mediterra-nean countries)	Algeria, Israel Morocco, Tunisia, Egypt, Jordan, Syria	1975-1977	Unilateral PTA	2.7%	11.3%	81%	Cooperation in social affairs, migration, human rights, and democracy. Preferences on industrial goods only, with strict rules of origin. A redirected Mediterranean Policy (RMP) was initiated in 1992 and substantially increased the amount of development aid and extended EU trade preferences.	EU Commission 2004 [4], Raya (1999)
AFS and SPG	EU members joined between Tokyo and Uruguay rounds	See note	Bilateral CU	n/a	n/a	100%	AFS: Austria (1995), Finland (1995), Sweden (1995). SPG: Spain (1986), Portugal (1986), Greece (1981). Accession years in parenthesis.	EU Commission 2004 [4]

Notes: (1) Romania and Bulgaria signed FTA agreements with the EU in 1993, hence should be part of the CEC category but they are not included in our analysis due to lack of data. (2) There exist special arrangements supporting measures to combat drugs under the GSP program. The recipients are the Andean group (Colombia, Venezuela, Ecuador, Peru, and Bolivia) and the Central American Common Market group (Costa Rica, El Salvador, Guatemala, Honduras, and Nicaragua). The extra concessions data for these groups is not available and hence they are considered only as part of the general GSP program. (3) Lebanon made a unilateral PTA arrangement with the EU in 1977 but it is not included in our estimations due to again lack of data. (4) http://europa.eu/pol/comm/index_en.htm accessed June 2006. (5) http://secretariat.efta.int/ accessed June 2006.

TABLE A2. Summary Statistics

Variable	Mean	Std. Dev.	Min	Max
Δt	-0.030	0.022	-0.268	0.000
I^{any0}	0.939	0.239	0	1
I^{evy0}	0.133	0.339	0	1
I^{any}	0.954	0.210	0	1
I^{pos}	0.015	0.121	0	1
I^{afs}	0.902	0.297	0	1
I^{spg}	0.875	0.330	0	1
I^{hiexp}	0.891	0.311	0	1
I^{gsp}	0.646	0.478	0	1
I^{gspl}	0.190	0.392	0	1
I^{acp}	0.291	0.454	0	1
I^{eftx}	0.870	0.337	0	1
I^{med}	0.508	0.500	0	1
I^{cec}	0.671	0.470	0	1
Δx	-2.004	1.853	-13.884	5.466
R	-0.460	0.118	-0.960	0.000
D^{anyexp}	0.984	0.126	0	1
D^{evyexp}	0.167	0.373	0	1
D^{gsp}	0.899	0.302	0	1
D^{gspl}	0.224	0.417	0	1
D^{acp}	0.333	0.471	0	1
D^{eftx}	0.904	0.294	0	1
D^{med}	0.594	0.491	0	1
D^{cec}	0.810	0.355	0	1
D^{hiexp}	0.948	0.221	0	1
R^{uni}	-0.267	0.139	-0.922	0
D^{ntb}	0.288	0.453	0	1
D^{ntball}	0.096	0.295	0	1
D^{ntball}xD^{anyexp}	0.091	0.288	0	1
D^{ntball}xD^{evyexp}	0.010	0.101	0	1
D^{ntball}xD^{gsp}	0.069	0.253	0	1
Δp_{9294}	0.001	0.461	-3.912	4.874
$\Delta scale$	0.307	0.322	-0.351	1.047
$(\Delta p_{9294})^{avg}$x$\Delta scale$	0.005	0.032	-0.053	0.126
P	0.502	0.500	0	1
t_{t-1}	0.079	0.046	0.005	0.65

The number of observations in our sample (n) is 6294. There were 8688 non-missing values for Δt (the dependent variable), which reduce to 7784 when we omit the lines with zero initial tariffs. Missing import and market access variables to construct the reciprocity variable reduce the sample to 6837, and missing price data to 6721. Production related data accounts for the remaining missing values and leaves us with n=6294.

TABLE A3. Data Sources and Definitions

Variable	Definition	Sources
Δt_i	The change in the bound advalorem MFN tariffs between the pre-Uruguay Round (UR) and post-UR periods for the HS 8-digit product *i*.	WTO
$I_i^{any0}\ [I_i^{evy0}]$	Equals 1 if HS 8 product *i* is imported in 1994 and receives a zero preferential rate by EU under any [all for I_i^{evy0}] of its PTAs. Excludes PTAs with CET	Comext-Trains
$I_i^{any}\ [I_i^{pos}]$	Equals 1, if HS 8 product *i* is imported in 1994 and receives either a duty-free or positive preferential tariff rate [a positive preferential rate only for I_i^{pos}].	Comext-Trains
$I_i^{afs}\ [I_i^{spg}]$	Equals 1 if HS 8 product *i* is imported by the EU in 1994 from the "recent" members Austria, Finland, or Sweden [Spain, Portugal, or Greece for I_i^{spg}].	Comext
$I_i^{hi\exp}$	An indicator equal to 1 if the exports of any of the PTA partners (excluding the ones with CET) to the EU in HS 8 product *i* is greater than the 25th percentile of its exports to the EU and is exported under the respective preferential program, i.e. $I_i^{any0} \times D_i^{hi\exp}$ (see below)	Comext-Trains
R_i	$\equiv \sum_k s_i^k \left(\sum_j w_j^k \Delta t_j^k / t_{jt}^k\right)$ change in market access provided by the major exporters of HS 8 product *i* to EU during UR; where $\Delta t_j^k / t_{jt}^k$ is the percentage tariff reduction by country *k* in good *j* between 1986 and 1995, w_j^k is the 1992 import share of good *j* in total imports of *k*, and s_i^k is the exports of a principal supplier (a top 5 exporter) *k* to EU in HS 8 product *i* as a share of total exports of *i* from EU's all principal suppliers.	Finger et al. (2002) and authors' calculations from Comext
R_i^{uni}	Reciprocity variable computed only for the unilateral liberalization (between 1986 and 1992) by the major exporters to the EU. The computation is otherwise similar to that of *R*.	same as R_i
Δx_I	The EU-wide change in the elasticity adjusted inverse import penetration ratio between 1978 (pre-Tokyo Round) and 1992 (pre-UR), for each 3-digit ISIC industry *I*. Computed using the 1978 members of the EU. Production values measured in domestic prices, $p_{IT}X_{IT}$, whereas import values measured in world prices, $p^*_{IT}M_{IT}$. Our elasticity measure is also evaluated at the domestic prices (and is for the EU as a whole), hence we calculate the measure required by the model as $x_{It} = X_{It} / M_{It}(M'_{It} / M_{It}) p^*_{It} = p_{It} X_{It} / p^*_{It} M_{It} (M'_{It} p_{It} / M_{It})$.	Kee et al. (2004) Unido, Comtrade and Penn World Tables
P_i	Proxy for the change in the MFN externality effect, Δm. Computed as a dummy equal to one if the change in the share of the non-top 5 exporters in total exports to the EU between 1989 and 1994 is above the median for an HS 8-digit product *i*.	Comext
$I_i^{pta\_name}$	Equals 1 if EU imports HS 8 product *i* at a zero preferential rate under the "pta_name" program, which includes GSP, GSPL, ACP, MED, CEC, EFTX.	Comext-Trains
$D_i^{any\exp[evy\exp]}$	Equals 1 if HS 8 product *i* is imported by EU from any [all for $D_i^{evy\exp}$] of its PTA partners in 1994 (regardless of a preference or not).	Comext
$D_i^{hi\exp}$	Equals 1, if HS 8 product *i* is an important export for a PTA partner (greater than the 25th percentile of its exports to the EU).	Comext
$D_i^{ntball[ntb]}$	Equals 1 if HS 8 product *i* is subject to a non-tariff barrier that applies to all exporters [at least one exporter for D_i^{ntb}] of *i* to the EU.	Trains
Δp_{i9294}	Change in world price of HS 8 product *i* between 1992 and 94. Computed using unit values for the EU and averaged over all of its exporters.	Comext
$\Delta scale_I$	The change in the EU-wide value added/number of firms by ISIC 3 sector *I* between 1978 and 1992.	Unido
$(\Delta p_{i9294})^{avg} \times \Delta scale_I$	Interaction of Δp_{i9294} averaged over industry *I* and $\Delta scale_I$.	Unido-Comtrade

FIGURE 1. Identification of the Stumbling Block Effect through MFN Tariff Changes