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Nuno Limão

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Nuno Limão, University of Maryland and CEPR

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Centre for Economic Policy Research
90–98 Goswell Rd, London EC1V 7RR, UK
Tel: (44 20) 7878 2900, Fax: (44 20) 7878 2999
Email: cepr@cepr.org, Website: www.cepr.org

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ABSTRACT

Preferential Trade Agreements as Stumbling Blocks for Multilateral Trade Liberalization: Evidence for the US*

Most countries are members of preferential trade agreements (PTAs). The effect of these agreements has attracted much interest and raised the question of whether PTAs promote or slow down multilateral trade liberalization, i.e. whether they are a 'building block' or a 'stumbling block' to multilateral liberalization. Despite this long-standing concern with PTAs and the lack of theoretical consensus there is no systematic evidence on whether they are actually a stumbling block to multilateral liberalization. We use detailed data on US tariff reductions during the most recent multilateral trade round to provide the first systematic evidence that the US's PTAs were a stumbling block to its multilateral liberalization. We also provide evidence of reciprocity in multilateral tariff reductions that amplify the stumbling block effect.

JEL Classification: D78, F13, F14 and F15
Keywords: MFN tariff concessions, multilateral trade negotiations, preferential trade agreements and reciprocity

Nuno Limão
Department of Economics
University of Maryland
3105 Tydings Hall
College Park, MD 20742
USA
Tel: (1 301) 405 7842
Fax: (1 301) 405 3542
Email: limao@econ.umd.edu

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1 Introduction

Nearly all countries are currently part of preferential trade agreements (PTAs). The importance of PTAs is difficult to overstate and thus understanding their effects has become a central concern for both politicians and economists. During the late 80's and early 90's multilateral trade negotiations were stalled whereas the U.S. and the E.U. pursued PTAs. This raised the question of whether PTAs are a "building block" or a "stumbling block" to multilateral trade liberalization (MTL), as Bhagwati framed it in 1991. That is whether PTAs promote or slow down MTL. After a decade of theoretical work on that question no consensus has been reached. What is more surprising is the absence of empirical evidence on this issue. Given the continued prominent role of PTAs and the current negotiations for a new multilateral trade round, answering this question is as pressing as ever.[1] In this paper we use detailed data on U.S. tariff reductions in the last multilateral trade round to provide the first systematic evidence that its PTAs did constitute a stumbling block to MTL.

PTAs can affect MTL through various channels. They can divert scarce negotiation resources, alter the number of negotiating parties and their bargaining power and affect a country's optimal multilateral tariff. Most of the theoretical literature has focused on this last channel; with an interesting strand modeling the decision of signing a PTA as endogenous and studying how it affects a country's incentives for MTL. Levy (1997) asks if the possibility of entering a preferential agreement causes the median voter to subsequently reject multilateral free trade. He shows that the median voter may reject multilateral free trade after voting for a PTA even though he would have accepted it if no PTA had been available. Grossman and Helpman (1995) show that when producers form lobbies the PTAs most likely to arise are those where partners have higher costs of production than the rest of the world, thus causing trade diversion. Krishna (1998) arrives at a similar conclusion using a different setup. More importantly, he argues that these PTAs can reduce the incentive for MTL. This occurs because the export rents generated by such PTAs disappear when countries liberalize multilaterally, and so the producers that benefit from those rents will oppose MTL.

Other authors have shown that PTAs can be a building block to MTL. Given the inherent complexity in modeling PTAs it is not surprising that both outcomes are possible. It is therefore perhaps more interesting to show when each outcome arises within the same model as in Bagwell and Staiger (1998). They use a repeated game to analyze how PTAs affect the incentives to set a self-enforcing multilateral tariff between two countries, A and B, which gain from a PTA because they are relatively more patient than a third one, C. If no PTAs are allowed then, in order to cooperate with each other, A and B must extend the same tariff to C. However, if PTAs are allowed then A and B can cooperate without lowering the tariff on C. This discriminatory effect leads to a higher multilateral tariff that is increasing in A and B's patience. This can be offset by a complementarity effect between the preferential and multilateral tariff that arises because the PTA leads A and B to trade more with each other and less with C. Because the last effect is independent of how patient A and B are Bagwell and Staiger show that PTAs are a stumbling block if countries are very patient and a building block otherwise.[2]

[1] From 1948-94 there were 124 notifications to the WTO of distinct PTAs by its members. Since 1995 there have been at least an additional 130. The WTO estimates that 300 PTAs will be in force by 2005. <http://www.wto.org/ >.

[2] There are other important contributions, for a survey see Winters (1999). A different strand of the literature started by Krugman (1991) analyzes the welfare path for exogenously expanding trading blocs; Frankel, Stein and Wei (1996) analyze important extensions in this literature.

The majority of the models are driven by the effects of the PTAs on trade volumes and prices. However, a large number of recent PTAs have taken place between large and small countries, what we will refer to as LSPTAs. Small countries are unlikely to have much effect on prices or the trade volumes of large countries with the rest of the world. Therefore, most of the existing models predict that the recent LSPTAs, including some U.S. agreements, do not affect the incentives of large countries for MTL.

One important feature of LSPTAs is that large developed economies lower their trade barriers in exchange for cooperation by small developing countries in non-trade issues such as labor and environmental standards, intellectual property protection and drug enforcement.[3] Limão (2002) models the interaction between LSPTAs and MTL. He shows that LSPTAs generate a strategic motive for large countries to maintain some of their multilateral tariffs relatively higher. Thus even agreements with countries that are small, in the sense of not affecting trade volumes or prices for the large country, can generate a stumbling block effect. More specifically, this model predicts that the large country maintains higher *multilateral* tariffs on products imported from the preferential partner relative to similar products that it imports only from the rest of the world. In the next section we show how the theory yields this prediction and argue that the basic strategic effect also applies to more general types of PTAs. From there we will search for a stumbling or building block effect in the data. Given the absence of empirical evidence we think this is a more reasonable first approach than testing specific models against each other.

The predictions generated by most previous models focused on the probability of multilateral trade round occurring. These are nearly impossible to test in a systematic way. This is why it is important to derive a prediction that emphasizes the effects of PTAs on the incentives to set tariffs on different products. The latter can be tested using tariff data at the product level for a single country in a given trade round, provided there is sufficient variation in tariff *changes*. This is the case for the U.S. both across and even within industries, which allows us to identify the stumbling block effect precisely. Moreover, focusing on the impact of PTAs on tariff changes captures an important concern of the WTO, particularly when it applies to a large country such as the U.S.[4]

By using detailed data on product level tariff concessions by the U.S. we can control for unobserved product characteristics that determine tariff levels. Moreover, our data allows us to control for all determinants of tariff *changes* that occur at the industry level. These include whether an industry is a large employer, is organized into a lobby and how powerful the lobby is relative to its foreign counterparts. These and other industry characteristics are emphasized by various theoretical models and explain a non-negligible part of the cross-sectional industry variation of trade protection; thus our ability to control for them is an important feature of this work.[5]

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[3] Current examples of LSPTAs include the U.S.'s agreements with Jordan, the Andean countries and other Latin American and Caribbean countries; the Eastern European and Mediterranean agreements signed by the E.U. and the preferential treatment that the E.U. and the U.S. extend to most developing countries through the Generalized System of Preferences.

[4] According to a 1999 WTO decision "Any preferential tariff treatment implemented pursuant to this Waiver [of the MFN rule for the purpose of preferential tariff treatment for LDCs …] shall not constitute an impediment to the reduction or elimination of tariffs on a most-favoured-nation basis." In WT/L/304, p. 2, accessed April 2003 on <http://www.wto.org/english/docs_e/legal_e/ waiver1999_e.pdf>.

[5] Trefler (1993), Goldberg and Maggi (1999). Most studies do not study the variation over time; Baldwin (1985) is one exception. He considers the changes in the average U.S. MFN tariff for 292 industries during the Tokyo Round and finds that lower skill, high employment industries are correlated with smaller cuts.

We find that a stumbling block effect is present. More specifically, we estimate that the U.S. cuts in *multilateral* tariffs were smaller for products that it also imports under its various PTAs relative to similar products that did not receive preferential treatment. The effect is stronger for products that are exported under all PTAs or constitute relatively larger shares of a given PTA's exports to the U.S. Moreover, the effect is robust to the possibility of reverse causation, which can arise if the choice of goods that receive preferences depends on MFN tariffs. We address this by using several instruments that we justify and test in some detail such as whether the product was exported by a PTA partner before the multilateral tariff changes, transport costs and world price variables.

Using the results on multilateral tariff reductions to derive the effect on domestic U.S. prices and world prices we estimate that the price effects for goods exported under any PTA were on average only 52% of those in similar non-PTA goods. If the good was exported by every PTA then the effect was even smaller, 23%. Furthermore, the effects are not simply driven by NAFTA. We find a similar effect when we separate NAFTA from all other PTAs and even individual agreements with small countries had a significant effect.

We also model tariff reciprocity empirically and provide novel estimates of reciprocity in multilateral trade agreements. Reciprocity in tariff concessions is a key concept in the GATT and forms the cornerstone of the main economic models of multilateral trade liberalization.[6] However, thus far, there has been no systematic evidence that countries respond with larger tariff cuts when offered larger tariff reductions by their partners. We test and find that partner countries' reciprocal tariff reductions are endogenous and argue that the timing and mode of the last multilateral trade negotiations provide a good instrument. After instrumenting we find reciprocity for U.S. products that were not subject to non-tariff barriers. The reciprocity finding is interesting in its own right. Moreover, it can amplify the impact of the stumbling block effect as other countries reciprocate the smaller cuts in U.S. tariffs with fewer reductions of their own.

The paper is organized as follows. In the next section we outline the theory that generates the testable prediction and present preliminary evidence. In section 3 we discuss relevant information about the GATT/WTO and the PTAs used for the test. In section 4 we provide an empirical model. In section 5 we estimate and quantify the importance of the stumbling block effect. We summarize the results and discuss some implications in the last section.

2 Theoretical prediction and preliminary evidence

Several models predict that PTAs can affect a country's multilateral tariff through its effect on trade volumes and prices, which can generate trade diversion or creation. In this section we outline a model where those effects are not necessary for PTAs to affect the U.S. multilateral tariffs. The objective is to provide a motivation to study the agreements the U.S. has even if the partners are too small affect prices and quantities in the U.S. significantly. We describe how the specific prediction that we test is generated in the context of a specific model but also argue that the basic mechanism that generates this prediction is fairly general and extends to other models. Finally we present some anecdotal and preliminary evidence.

[6] See for example Bagwell and Staiger (1999).

Limão (2002) models two symmetric regional blocs, each containing two countries, *Large* and *Small*. There exist two externalities within each bloc. First, both governments in a bloc can provide a public good with regional spillovers. *Small* places no weight on the public good, which results in an underprovision from *Large*'s perspective. Second, there is a terms-of-trade externality; *Large* can use a tariff to depress the price of *Small*'s exports. The countries can internalize these effects via a PTA where *Large* lowers its tariffs on *Small*'s exports in exchange for an increase in the latter's provision of the public good. Importantly, the PTA is modeled such that it does not directly affect prices or trade volumes of any good subject to tariffs in *Large*; its only direct trade effect is to increase the price that *Small* receives for the fixed amount of its exports. So the impact of this PTA on multilateral tariffs is present even in the absence of trade creation or diversion.

The large countries in the two blocs also face a terms-of-trade externality between them and internalize it via multilateral trade negotiations. We assume these negotiations occur in a stage prior to the PTAs and maximize the large countries' joint welfare.[7] According to the Most-Favorite-Nation rule (MFN) in the WTO a multilateral tariff reduction between two members must be extended to all members. However, WTO rules allow exceptions in the form of "lower than MFN" tariffs. That is, if *Large* negotiates an "MFN" tariff with the other large country it cannot set a higher tariff on the small partner if the latter is also a WTO member. However, *Large* can offer a lower preferential tariff to *Small*.

The stumbling block effect should now be clear. If *Large* negotiates a multilateral tariff of zero in the goods it imports from *Small* it cannot offer a preferential tariff reduction. By raising the multilateral tariff slightly *Large* incurs no first order cost and it gains since it can offer a preference to *Small* and induce it to supply some of the regional public good.[8] This extra cost of reducing the multilateral tariff is not present in similar goods that are not imported from *Small*.

The mechanism outlined above suggests the specific prediction we test: whether the *multilateral* tariff reductions that the U.S. made during the Uruguay Round were smaller in goods that the U.S. imports both from its PTA *and* from its non-PTA partners. Thus the control group is formed by products in the same industry that the U.S. does *not* import under preferential agreements.

Some PTAs fit the specific assumptions of this model closely, as we describe in the next section, and therefore the prediction we test applies directly to them. However, the basic intuition may also apply to PTAs such as NAFTA, where the gains for both partners are more directly related to trade. Lowering multilateral tariffs reduces the preferential margins that can be negotiated by any two WTO members. If these preferences have value for the members then there is an extra cost to reducing multilateral tariffs.[9]

The conjecture that this strategic incentive to maintain relatively higher multilateral tariffs in some products is also present in models that address mainly trade issues is supported by the results in Bagwell and Staiger (2003). Their paper analyzes accession to the WTO using a sequential bargaining mechanism similar to Limão (2002) in the context of a pure trade model and show that it can cause WTO members to hold back tariff reductions on goods exported by countries expected to accede to the WTO.

[7] Introducing political economy motives for tariffs here does not affect the results qualitatively; it simply rescales the level of the cooperative multilateral tariff, as shown in Bagwell and Staiger (1999).

[8] We assume that Large cannot simply make a transfer payment to obtain the regional public good from Small. If Large faced no such constraint then we should not observe any LSPTAs.

[9] In Krishna (1998) the PTA is valued because it generates profits for producers exporting to the PTA partner. Therefore the PTA reduces the incentive to reduce multilateral tariffs on those goods. If Krishna's model were to include similar products not traded between the PTA partners we should obtain a similar prediction to the one we test.

The strategic effect we have described is recognized by governments and taken into account during negotiations. Both UNCTAD and beneficiary countries of the U.S.'s GSP program regularly expressed their concern during the Uruguay Round about how MTL in certain products caused an "erosion of preferences" for GSP beneficiaries exporting to the U.S. More importantly policy makers in the U.S. were clearly aware of the impact of an "erosion of preferences" on the willingness of PTA partners to cooperate on non-trade issues. According to the International Trade director of the General Accounting Office:

> "Because GSP benefits are limited, and would decline if the GATT Uruguay Round agreement is enacted, the program provides only a modest degree of leverage to encourage beneficiary country governments to change their country practices [for enforcing intellectual property rights and labor standards] (…) It should be noted that tariff reductions negotiated in GATT, if implemented, will reduce the value of the GSP's tariff preference by 40 percent and, therefore, the incentive for beneficiary countries to participate in the GSP program."[10]

There is also anecdotal evidence that the U.S. responds to concerns by PTA partners of "erosion of preferences" by maintaining higher multilateral tariffs on PTA products even though it lowers the tariffs for similar products not exported by PTA partners. The following quote provides a good example and it shows exactly how we uncover the stumbling block effect in the data. One of the largest exports from several Caribbean countries to the U.S. is low-valued bottled and bulk rum that enters the U.S. duty-free under the Caribbean Basin Initiative.

> "*In WTO tariff negotiations in 1996, U.S. and E.U. negotiators had initially agreed to phase out all tariffs on rum and other 'white spirits' by 2000*. This unexpected development was met with alarm by Caribbean governments, Administration officials and Members of Congress. They emphasized to the trade negotiators that such a drastic change in the tariff structure for rum would deal a severe blow to the economies of the U.S.VI, Puerto Rico, and the Caribbean. *In response to this outcry, U.S. and E.U. negotiators, as well as Caribbean governments and producers, revisited rum tariffs (…) Under this compromise, the United States agreed to substantially liberalize duties on expensive rum. However, to protect the interests of the U.S.VI and other Caribbean island producers, the United States also agreed to maintain existing MFN duties on low-value bottled and bulk rum.*"[11]

The anecdote above is supported by the preliminary evidence in Figure 1, which suggests that the PTAs signed by the U.S. were a stumbling block in the Uruguay Round. The figure plots the average cuts in the U.S. MFN tariff factor in a given sector. The horizontal axis measures the cuts for goods exported by each PTA partner and the vertical axis applies to all other products in the same sector. Most of the points lie above the 45^o line indicating that the U.S. did not cut its MFN tariffs on PTA goods by as much as on similar goods not exported by its PTA partners. We will show that the result suggested by figure 1 is confirmed after we address several potential econometric issues.

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[10] Testimony before the Subcommittee on International Trade. In "International Trade: Issues Concerning the Generalized System of Preferences", GAO/T-GGD-94-174, June 20th , 1994.

[11] Emphasis added. Testimony before the House Committee on Ways and Means, May 8th, 2001. Accessed June 2003 on <http://waysandmeans.house.gov/legacy/trade/107cong/5-8-01/5-8chri.htm>.

3 The U.S.'s preferential trade agreements and the Uruguay Round

3.1 The U.S.'s preferential trade agreements

By 1994 the U.S. provided preferential trade treatment to Canada and Mexico under NAFTA (effective in 1994), Israel (1985), several developing countries under the Generalized System of Preferences (GSP, 1976), Caribbean countries (CBI, 1984) and Andean countries (ATPA, 1992). Nearly all products with a positive MFN tariff receive preferential treatment under NAFTA and imports of these products from NAFTA accounted for approximately 12% of total U.S. imports in 1994. The import shares for products that enter the U.S. under preferential treatment for the other agreements are smaller: approximately 3% for the GSP and less than 0.4% for Israel, as well as for the ATPA and the CBI.[12]

The aggregate import shares suggest that except for NAFTA, and possibly the GSP, the PTAs in effect by 1994 were unlikely to have a very large direct effect on the aggregate volume of U.S. trade with the rest of the world. However, the remaining PTAs may still have a significant effect on the U.S. MFN tariff for two reasons. First, there are a few specific products in which the GSP, CBI, ATPA and Israel have large shares. More importantly, as pointed out previously, the PTA partner need not have a large share in any given good. Provided that the U.S. places sufficient weight on the concessions that the partner provides it will have a bargaining incentive to maintain MFN tariffs relatively higher on the set of products it imports under the PTAs.

When the bargaining effect is present the number of goods exported under the PTA provides information about the scope of the potential stumbling block effect. Even at the most disaggregated tariff line (8-digit Harmonized Standard) the smaller PTAs export a significant proportion of different types of goods (over 15% for any of the PTAs in our sample) and in nearly all of those lines the U.S. also imports from non-PTA partners. Therefore the small PTAs have the potential to affect a large number of non-PTA countries and consequently a non-negligible amount of U.S. trade.

We estimate the average stumbling block effect for all PTAs and also for each individually to capture whether small countries have any impact. We can interpret the estimates for the CBI and ATPA as capturing the strategic stumbling block effect identified in Limão (2002) because these agreements fit the key assumptions of that model. In particular, the largest trade partner for both is the U.S., which provides duty-free access to their exports. In exchange the ATPA and CBI do not provide preferential trade access to U.S. exports but must cooperate on labor standards, intellectual property enforcement and, in the case of the Andean countries, the war on drugs. The ATPA and CBI countries are small both in terms of the U.S. import share and also because they have no *direct* effect on multilateral trade negotiations. Moreover, the non-trade issues on which the ATPA and CBI cooperate are important to the U.S., as is clear from the $1.6bn of aid approved by the U.S. Congress to eradicate plantations of coca in the Andean countries.

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[12] <http://www.usitc.gov/ave.pdf> accessed June 2003.

3.2 Uruguay Round negotiations

The question of whether preferential agreements promote or hinder MTL is ultimately about changes in multilateral tariffs. By focusing on the U.S. behavior during the UR we implicitly assume that those changes will mainly occur after multilateral trade negotiations, which is in fact what we observe in the U.S. data in the 90's. However, if a country's incentives to raise its multilateral tariffs increase due to PTAs then why should it wait for a multilateral trade round to implement those changes? The reason is simple; the U.S.'s MFN tariffs are bound by previous multilateral trade agreements. This implies a potential additional cost to the U.S. from raising the MFN tariffs in the PTA products—higher tariffs on U.S. exports as other WTO members retaliate. So, raising bound tariffs is a costly option. Moreover, given the timing of the ATPA and NAFTA at least, the U.S. could simply wait and not reduce the MFN tariffs on PTA goods by as much as on other goods to avoid any penalty from retaliation.

The cost of focusing on the changes that take place during a multilateral trade round is that we must model it empirically. Despite great interest and recent theoretical work providing an economic theory for GATT rules and multilateral tariff negotiations there is surprisingly little empirical evidence on the presence and importance of some of the basic features that characterize those models, such as reciprocity.[13] Thus modeling the negotiations also provides novel estimates that are interesting in their own right.

We require a model that provides reasonable estimates of the tariff changes that occur during multilateral trade negotiations to consistently estimate the stumbling block effect. Moreover, if reciprocity is present and also applies to other WTO members it amplifies the importance of PTAs as stumbling blocks. As we will see the U.S. offered smaller tariff reductions in PTA goods. Suppose country k does not have a PTA with the U.S. but exports only PTA goods to the U.S. Reciprocity implies that k will offer fewer reductions on the goods it imports from the U.S. The U.S. internalizes this effect on itself. However, countries that export goods similar to the U.S. to k will also face k's higher MFN tariff and potentially lower export prices. Therefore the combination of the stumbling block effect, reciprocity and MFN can affect third countries even if they do not trade directly with the U.S.

As we will describe the timing of the negotiations is crucial in solving a potential endogeneity problem associated with estimating tariff reciprocity. For now we note that the negotiations started in 1986 but were stalled until 1992. It was only in July of 1993 that the U.S., E.U., Japan and Canada announced significant progress. Some details were still unresolved in early 1994 and the agreement's implementation started in 1995.[14]

4 Empirical model

Our empirical model of the U.S. MFN tariff is given by the following linear approximation.

(E1) $$\tau_{it} = \phi G_i z_T + \phi_2 z_T + \phi_3 G_i + \alpha_i + \alpha_t + \alpha_{It} + \beta(b_t - b^k_t)\, 1^k_i + (-\rho m a^k_t)\, 1^k_i + \varepsilon_{it} \qquad i=1,\dots,N;\ t=1,2$$

[13] Finger et al (1999) is an important exception. They calculate market access concessions for various countries during the Uruguay Round, measured as weighted tariff changes, and find that the tariff concessions given were not balanced by those received for most countries. Given that both other trade instruments and concessions in non-trade issues were object of negotiation this finding is not surprising. Their finding does not preclude the notion of product reciprocity that we estimate, which requires a positive correlation between tariff concessions given and received by pairs of countries but not necessarily a balance.

[14] In <http://www.wto.org/english/thewto_e/whatis_e/tif_e/fact5_e.htm>, accessed April 2003.

where we initially consider the case where the U.S. is negotiating with a single country, k, we generalize this to multiple countries below. The dependent variable τ_{it} is a measure of the U.S. MFN bound ad valorem tariff rate negotiated in period t on the 8-digit product i. The indicator variable G_i denotes whether the good is exported to the U.S. under a preferential agreement. The indicator z_T is one after the PTA and zero before it. Thus ϕ is the coefficient of interest, predicted to be positive if the PTA is a stumbling block. Both G_i and z_T also enter separately to ensure that their interaction correctly estimates the stumbling block effect and not an average product or time effect. Some products have consistently high or low average tariffs; all the determinants of the historical average value of product i's tariff are captured by α_i, a full set of product dummies. The common determinants of the average tariff in period t such as the need for tariff revenue (obviously low in the case of the U.S.), are captured by the time effects, α_t. Not all parameters in (E1) are separately identifiable but the ones of interest are, as we will show.

The tariff determinants for the group of goods i in each industry I during period t are captured by α_{It}. This set of time-varying industry dummies captures factors that are emphasized by different political economy theories such as: the effects of an industry's import penetration ratio, labor intensity and lobbying strength of domestic industries relative to their foreign counterparts. Because the industry effects are time-varying they account for different *changes* in tariffs across industries, which helps to ensure that the stumbling block effect is not biased due to omitted industry level determinants.[15]

4.1 Reciprocity and bargaining in trade negotiations

The variable 1^k_i is an indicator for whether country k exports good i to the U.S. Thus the last two variables vary by product. They capture two important determinants of tariff changes during multilateral negotiations: the U.S.'s bargaining power relative to country k and a measure of product specific reciprocity. To describe and model these effects it is useful to difference (E1) across the last two multilateral trade rounds.[16]

$$\text{(E2)} \quad \Delta\tau_{it} = \phi G_i + \phi_2 + \Delta\alpha_t + \Delta\alpha_{It} + \beta\Delta(b_t - b^k_t)1^k_i + (-\rho\Delta ma^k_t)\, 1^k_i + \Delta\varepsilon_{it} \qquad i=1,\ldots,N$$

[15] Marvel and Ray (1983) use the average U.S. tariff for 261 industries to analyze its determinants in 1970 and find that the 1965 tariff is an important determinant. The coefficient they find is lower than unity suggesting that higher initial tariffs lead to smaller reductions. This result is likely driven by omitted industry variables--organized industries are likely to have higher initial tariffs and, given they are already organized, are able to lobby for fewer reductions. By using time-varying industry dummies we control for such omitted variables.

[16] In (E1) the timing t=1 corresponds to the period of the final negotiation of the previous trade round, Tokyo (1977-78) and t =2corresponds to the final stage of the negotiation of the UR (1993-94). For the ATPA T=1991-92. Similarly NAFTA, Israel and the CBI were initiated after the Tokyo Round so differencing (E1) between t and t-1 leads exactly to (E2) for those four agreements. Some preferences under the GSP were already included in the Tokyo Round and may therefore already be reflected in the MFN tariff at t-1. However, the U.S. tightened the eligibility criteria for GSP between the two rounds. So the incentive to maintain a higher MFN tariff for GSP products and therefore the stumbling block effect may also have increased and we should then interpret the estimate of ϕ in (E2) for GSP as such an increase.

Some theoretical models of GATT negotiations recognize the potential importance of bargaining power but do not model its underlying determinants.[17] Therefore we will take a broad and simple view of bargaining power and argue that country *k* exporting good *i*, is expected to obtain a relatively larger cut in the U.S. tariff if any of the following conditions holds. First, if *k* is relatively better prepared for the negotiation, perhaps because it has more resources to dedicate to it. Second, if it imports more of the U.S. goods such that it has relatively high market power. Third, if *k* attracts relatively more U.S. foreign investment, allowing it to offer (perhaps implicitly) better terms during trade negotiations for issues such as taxation of profits of U.S. multinationals. We use relative GDP to proxy for these three motives. Thus the prediction is that an increase in country *k*'s bargaining power relative to the U.S., i.e. a decrease in the log difference of GDP in the U.S. and country *k*, $(b_t - b^k_t)$, causes the U.S. to lower its tariff in the goods exported by *k*, i.e. $\beta>0$.

The other important determinant we model is reciprocity—the extent to which the U.S. reduces its tariffs by more when its partners offer larger tariff reductions on U.S. products. Reciprocity in multilateral negotiations has typically been sought in first-differences (Bhagwati 1991). This means that negotiators focus on measures of *change* in protection offered in a particular negotiation as opposed to the level; these are often called market access concessions. If product reciprocity is followed then the larger the *aggregate* market access concession *k* offers to the U.S., i.e. the larger Δma^k_t is, the larger the U.S. tariff reduction, $-\Delta\tau_{it}$, is on the products exported by *k*, thus if product reciprocity occurred then $\rho>0$.

We expect to find product reciprocity because during the UR negotiators used an item-by-item approach.[18] That is, instead of agreeing to reduce tariffs across the board according to some formula, as occurred during the previous round, governments negotiated different tariff reductions for different products

Since each country imports several goods there is a need to aggregate its proposed percentage tariff changes over all its imports to determine their proposed aggregate change in market access. GATT negotiators use different methods to measure proposed changes in market access. Two common methods that differ in the weights used for this aggregation are the following. First, a simple import weighted average tariff cut uses *k*'s import share of good *j* in all of *k*'s imports as the weight. Second, the "trade coverage" measure simply uses the value of *k*'s imports in *j* as the weight.[19]

Using either of the measures above we can express the *aggregate* change in market access proposed by country *k* as $\Delta ma^k_t = \Sigma_j(-\Delta\tau^k_{jt})w^k_{jT}$, where $-\Delta\tau^k_{jt}$ is the percentage tariff *reduction* by *k* in each good *j* and w^k_{jT} is the weight that depends on which measure is used. For a given weight, larger tariff reductions by *k* provide more market access to the U.S. and therefore, if the negotiators follow reciprocity, lead to a lower U.S. tariff in each of the goods *i* that country *k* exports to the U.S. So, again if reciprocity is present, we should find that $\rho>0$.

Using the market access definition derived and writing (E2) in terms of estimable coefficients we have:

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[17] Bagwell and Staiger (1999) for example model reciprocal trade negotiations and argue that the initial stage is precisely the determination of a specific world price vector, which is the outcome of bargaining.

[18] See Hoekman-Kostecki (2001) p.133 and Jackson (1997) p. 146.

[19] Jackson (1997), p.147 and Hoekman-Kostecki (2001), p.124. The two measures are identical up to a scalar (the total value of *k*'s imports) if *k* imports all goods from a single country but they differ if *k* imports from different countries. The "trade coverage" measure for *k*'s concessions to the U.S. may reflect the gains to the U.S. more accurately since it uses the value of U.S. exports to *k*, whereas the import weighted measure uses the value of *k*'s imports for each good *j* both from the U.S. and other countries as a share of its total imports. However, the import weighted measure for each country *k* is more easily available during negotiations because it is not partner specific. Deciding which is a better measure of reciprocity during negotiations clearly depends on which the negotiators actually use for each country. However, provided that both measures are positively correlated we expect that either can capture if reciprocity in tariff concessions did occur.

$$\text{(E3)} \quad \Delta\tau_{it} = \phi G_i + a + a_I + \beta\Delta(b_t - b^k_t)1^k_i + \rho(\Sigma_j \Delta\tau^k_{jt}\, w^k_{jT})1^k_i + u_i \qquad i=1,\ldots,N$$

The coefficient a denotes an intercept that estimates the average MFN tariff change for the excluded industry (miscellaneous manufacturing); a_I represents the set of included industry dummies. Note that although the reciprocity variable captures the aggregate concession by k to the U.S. it continues to vary by product because no country k exports all goods to the U.S. Moreover, for any given good i there will typically be several exporters to the U.S. According to the principal supplier rule in the GATT countries negotiate only with the top exporters. Thus we aggregate the observations for each good i over its principal suppliers to the U.S. using their individual export value as a share of their total value exported to the U.S., s_{iT}^k.

$$\text{(E4)} \quad \Delta\tau_{it} = \phi G_i + a + a_I + \beta\, \Sigma_k\, s_{iT}^k \Delta(b_t - b^k_t) + \rho\, \Sigma_k\, s_{iT}^k (\Sigma_j\, \Delta\tau^k_{jt}\, w^k_{jT}) + u_i \qquad i=1,\ldots,N$$

In sum, we estimate (E4) to test the following basic predictions. If PTAs are a stumbling block then ϕ is positive. If bargaining power and reciprocity are important determinants of multilateral tariff setting then we expect β and ρ to be positive.

4.2 Endogeneity

To obtain consistent estimates for the parameters in (E4) we must address the following potential endogeneity problems associated with the reciprocity and PTA variables.

The potential for endogeneity of the reciprocity variable is due to reverse causation. If reciprocity was followed then a fraction of country k's tariff reductions in the UR would be due to reductions in U.S. tariffs on k's products. We use two related facts about the timing and mode of negotiations in the UR that allow us to construct an instrument for reciprocity.

Between the start of the negotiations, 1986, and the final market access negotiations, 1992-93, a number of countries undertook unilateral trade liberalization.[20] When each of these countries decided to liberalize unilaterally the decision was not conditional on other countries lowering their tariffs, i.e. it was not reciprocal. Nonetheless, based on interviews with the negotiating delegations Finger et al (1999) report that "according to delegations the informal practice was […that] countries which had, after 1986, unilaterally reduced their tariffs would be given 'credit' at the round to the extent that they bound these cuts at the round."(p.9) Therefore the unilateral liberalization by U.S. partners between 1986-92 was predetermined relative to U.S. tariff reductions negotiated between 1992-94 but relevant in affecting the extent of those U.S. cuts, because the U.S. responded to that unilateral liberalization ex-post in the same way it responded to tariff reductions, negotiated in 1992-94, that were explicitly reciprocal. Thus we can use a country's unilateral liberalization (1986-92) as the instrument for its total liberalization (1986-95).

[20] In the data all countries show some unilateral liberalization with the least occurring among E.U. members.

The PTA variable may also be endogeneous to tariff changes. For NAFTA all the products in our sample receive preferences, whether or not they are exported by Canada and Mexico. However, for the other PTAs that is not true. This raises the question of which goods exported by a PTA partner receive preferences and whether this decision was determined by *changes* in the U.S. MFN tariffs. The first step we take in addressing this problem is to use preferences that were determined before the changes in the U.S. MFN tariff were implemented. However, using the "pre-determined" preferences may still be insufficient because, if some of those tariff changes were expected, there may still be a correlation between the preferences and an omitted variable in the error term of (E4). One potentially important source of those expected changes are industry level characteristics unobservable to us but known by the U.S. and its PTA partners. The solution is to control for these characteristics, as we do, by using industry dummies in (E4).

There may also be product level determinants that influence both MFN tariffs and preferences. The most obvious is an unobserved fixed product characteristic such that the current and expected future MFN tariff level is either high or low. Differencing should address this. But there may be product characteristics that indicate if the tariff *change* will be high or low. For example a product which is protected by a non-tariff barrier (NTB) on any country can indicate how a tariff will subsequently change. Therefore we control for the existence of NTBs in estimating (E4).

We test and find that using predetermined preferences and adding industry and product controls does not fully solve this endogeneity problem. Therefore we instrument the variable G_i. The list of instruments we use includes whether the PTA partners exported the good to the U.S. before the UR *independently* of whether the preference is received or not. This is a reasonable candidate for an instrument because it is correlated with the set of goods that are exported *and* receive preferences. Moreover, the decision to export a good prior to the UR may be independent of the subsequent MFN tariff changes so that the instrument is uncorrelated with the error term in (E4).[21] To test the exogeneity of exports as an instrument we also include other instruments that can capture whether a particular good is likely to be exported. In particular we use transport costs and world price changes prior to the UR, which are a priori more likely to be exogenous relative to tariff changes in the UR. Therefore we use the extra orthogonality conditions implied by this last set of instruments to verify the exogeneity of exports as an instrument.

5 Estimation

5.1 Data

In appendix B we provide summary statistics and detailed definitions for each variable. We construct the PTA variable, G_i, and the relative bargaining power variable, as described above. The tariff based variables are constructed as follows. First, the dependent variable we use measures the growth of the MFN ad valorem tariff *factor*, $\Delta\ln(1+\tau_{it})$. This measure can be translated into price effects as we will show; therefore it is more informative about the economic effects of the stumbling block effect than simply considering changes in the tariff level.

[21] This is possible because the decision to export may depend mainly on a country's relative resource and factor abundance and the world price and trading costs associated with that good.

Second, for the reciprocity variable for a given country k, $\Sigma_j \Delta\tau^k_{jt}\, w^k_{jT}$, we use the import weighted average tariff reduction calculated by Finger et al (1999). They use the available tariff reductions for each WTO member, k, in the UR for each product, j, weighted by that products' import share for each country k. Finger et al (1999) calculate two variants of this variable for each country: all tariff reductions from 1986-95 and reductions from 1992-95. From these we construct our measure of unilateral liberalization from 1986-92, which we use as an instrument.

5.2 Estimates

Before presenting the regression results we briefly describe how the dependent variable varies according to PTA vs. non-PTA goods without any controls. First, consider the basic difference-in-difference estimate. The mean change in the MFN tariff *factor* for PTA goods is -2.75%. For the other goods it is -4.03%. Thus the reduction for non-PTA goods is larger and the difference between the two is significant at the 1% significance level according to a simple t-test of difference in means.

Consider now the entire distribution of the changes in the MFN tariff factor for PTA and non-PTA goods. Figure 2 presents kernel density estimates for these two sub-samples, i.e. the "smoothed" histograms. The estimated density for the MFN tariff changes for the PTA goods is the dashed line, which is shifted towards zero. More specifically, few PTA goods had reductions of more than 6%. That is not the case for non-PTA goods, as is clear from the "fatter" dashed tail.

We note also that the difference in tariff reductions between PTA and non-PTA products, which we use to identify the stumbling block effect, ϕ, econometrically, is present for many industries. Thus the result will not be driven by any single industry. This is already obvious to some extent from Figure 1 that shows the variation at the more aggregated sector level. Moreover, within 69 out of the 93 2-digit harmonized standard industries there exist products in both the control and treatment group, which identify the stumbling block effect, ϕ. These 69 industries span a varied set of products including agriculture, prepared foods, chemicals, metals, plastics, leather, textiles and apparel, footwear and machinery and vehicles. In the 2-digit industries with non-PTA products their proportion can be as high as 60% and on average it is 12%.

We now confirm that the stumbling block effect suggested by the preceding two results and by figure 1 is robust to introducing the relevant controls and addressing various econometric issues; in section 5.5 we quantify its economic importance.

5.2.1 Stumbling block effects

Table 1 contains the results from estimating (E4). The first two columns contain OLS estimates for comparison purposes. The remaining columns contain different specifications in which we apply a GMM estimator and instrument for the endogenous variables, denoted by the symbol "†". All the specifications include industry dummies, which we do not report here. [22]

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[22] We only include products with positive tariffs before the UR for two reasons. First, when the MFN tariff is zero there is often no information about whether a preference exists, 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. In addition we include a dummy variable, DS, which accounts for 8-digit tariff lines that were subdivided in the WTO data and which we averaged over; eliminating this dummy does not affect the results.

The first point to note is that the estimate for the coefficient on ANYPTA, ϕ in (E4), is positive and significant over the different specifications, which is consistent with a stumbling block effect. The OLS estimate for ϕ in column 1 is 0.98, therefore it is biased downward relative to the IV estimate in column 3, 1.28. As we discuss below we can't reject endogeneity and thus we should focus on the IV estimates.[23]

We may expect different magnitudes of the stumbling block effect over different goods. In column 4 we include EVRPTA which is equal to one for the subset of goods that are exported under every PTA and find that the stumbling block effect is 60% larger for those products. The rum anecdote for the CBI in section 2 suggests that the stumbling block effect should be stronger in products that are important exports for a PTA partner. We test this in column 5 by including the interaction of the ANYPTA variable with an indicator equal to one if the export value of the product exported by a given PTA is above the mean of its total exports to the U.S. The effect is 48% higher for important exports. A similar qualitative result is obtained if we use the median as the threshold.

If the estimates just described are merely capturing the effect of NAFTA then PTAs with small countries should not affect the U.S. multilateral tariffs, as predicted by most standard models. To test this in column 6 we separate NAFTA from the remaining agreements and find that the agreements with smaller countries also have a significant effect, which has a magnitude statistically identical to NAFTA. Moreover, the number of goods affected under NAFTA is approximately the same as under all other PTAs combined.

In column 7 we identify the stumbling block effect for each of the individual PTAs, with the exception of Israel, so even PTAs with small countries such as the ATPA and CBI affect a large country's MFN tariffs. Moreover, the estimates for the ATPA and the CBI are identical, which suggests a similar mechanism is at work for those agreements. Since those two PTAs closely fit the assumptions of the model in Limão (2002) the similarity of the coefficients may indicate further support for that model.[24]

5.2.2 Multilateral negotiation effects

The bargaining power elasticity, corresponding to the coefficient β in (E4), is positive, as predicted, and significant. It indicates that a decrease in the U.S.'s bargaining power relative to a given country of 31% (the average in the sample), causes a decrease of 0.28% in the U.S. MFN tariff factor in *each* of the products for which that country is the main exporter to the U.S.

There are two important factors in identifying the reciprocity effect—captured by ρ in (E4)—controlling for the existence of NTBs and addressing its endogeneity. The reciprocity measure constructed by Finger et. al (1999) does not reflect any tariff equivalent changes due to NTB changes. Therefore, the tariff reciprocity variable should work as expected for the products where no NTBs are present. We test and confirm this by interacting the NTB indicator with the reciprocity variable.[25]

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[23] The estimate for ϕ is also positive and significant when we use the percentage change in tariffs instead of the log growth as a dependent variable.

[24] Note that the individual PTA variables are defined to be zero if the good is not exported under that PTA but some of those goods may still be exported by another PTA, so the control group when using the individual PTA variables is slightly changed relative to ANYPTA. To address this we re-estimated specifications (6) and (7) in table 1 including also the ANYPTA variable. The ANYPTA variable is then statistically insignificant and the coefficients for ATPA, CBI, GSP and Israel remain nearly unchanged. However, ANYPTA is highly correlated with the NAFTA variable, 0.76, since it is the PTA under which the largest number of goods is imported. Therefore the NAFTA coefficient becomes insignificant if ANYPTA is also included but the two coefficients are jointly significant (at 1%). Moreover, the coefficients of NAFTA and ANYPTA are identical so when we impose that constrain we again find a positive and significant coefficient for both NAFTA and ANYPTA. In the latter case the coefficients for ATPA, CBI, GSP and ISR are also similar to those presented in columns 6 and 7 of table 1.

[25] In products with NTBs an exporter is less likely to try to get the U.S. to reduce its tariffs because it will still face the protection from the NTB.

The IV estimates of reciprocity are positive and significant, ranging from 0.014 to 0.018. So a decrease in the tariff of a U.S. partner that exports good *i* leads to a decrease in the U.S. tariff of that product. As previously discussed the presence of reciprocity potentially amplifies the stumbling block effect. Finally, comparing the IV estimates with the insignificant OLS estimates we see the importance of instrumenting for this endogenous variable.

5.3 Specification tests and instrument choice

The bottom rows of table 1 present test statistics for the IV estimation. First, we perform Hausman endogeneity tests of the PTA and reciprocity variables, denoted with the symbol "†", and reject the consistency of the OLS estimates. Second, we reject the null hypothesis of homoskedasticity for all specifications, which suggests that GMM is more efficient than standard IV. Third, the equations are overidentified and we can't reject the validity of the overidentifying restrictions. Finally, because tests of overidentifying restrictions may have low power, we also test a subset of the instruments used and can't reject their exogeneity.

The first three columns in table 2a present the first stage regressions for the potentially endogenous variables in the basic specification in table 1 (column 3): ANYPTA, reciprocity and its interaction with NTB. The instruments in all three first stage regressions explain a significant part of the variation, with a partial R-square ranging from 0.49 to 0.72, which is a large proportion of the overall R-square. Moreover, we can reject the null that the instruments are jointly zero at a significance level always less than 1%. The first stage results for the specifications that include EVRPTA and ANYPTxHI_EX respectively are similar to those in table 2a so we do not report them here. The last three columns in table 2a present the first stage results for column 6 table 1. Table 2b presents the first stage for the specification with the individual agreements. The explanatory power of the instruments is also significant in this case for each of the regressions.

As previously noted the PTA variable may be endogenous with respect to MFN tariff changes because countries signing a PTA with the U.S. may not seek preferences in goods for which they expect the U.S. to make large multilateral concessions. In the case of NAFTA this is not an issue because in our sample all goods receive preferences. However, if other PTA partners selected products on which they receive preferences based on the expected MFN tariff change this it will bias any OLS estimates.

The first stage regressions for the PTA variables can be interpreted as linear probability models of whether a good is exported to the U.S. *under* a (or any) preferential program—so some of the results may be interesting in their own right. We use several instruments for the PTA variable(s). First, whether a country exported the good to the U.S. at all (with or without a preference) before the UR tariff changes were implemented. Second, whether the good was subject to a U.S. NTB that applied to *all* countries.[26] The rationale for this instrument is that a country is more likely to try to negotiate a preference in a good if it expects that otherwise it will certainly face an NTB. This applies more strongly if the country exports the good so we also include the interaction of these two variables. The export variable has the expected positive and significant effect in predicting the PTA variable. This is true not only for ANYPTA but also for the different specifications where the PTA variable is defined for the individual agreements in table 2b. For the individual agreements we also find that the NTB_ALL variable has a positive effect and is generally significant.

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[26] The variable NTB_ALL captures a subset of the goods for which NTB=1. The test of overidentifying restrictions does not reject excluding NTB_ALL from any specification.

Whether a PTA partner exported a particular good to the U.S. before the UR may depend on the U.S. MFN tariff *level* but is unlikely to depend on the *change* in the tariff that occurred afterwards. Nonetheless we test the export variables and respective interactions with the NTB_ALL variable and fail to reject their exogeneity. In the specification with specific agreements, column 6 table 1, this test also applies to the NAFTA export variable. The symbol "‡" next to a variable denotes that the instrument was tested for exogeneity and the probability at which we reject the null hypothesis of exogeneity is listed in the row in table 1 labeled "c-stat p".[27]

The main instrument we use for the reciprocity variable (and its interaction with the NTB variable) is unilateral reduction (and its interaction with NTB), as explained previously, and it has a positive and significant effect on total reduction (and its interaction with NTB).

Below we confirm that the stumbling block effect is robust to other factors. Before doing so it is useful to consider how to interpret and quantify the importance of the effects estimated.

5.4 Interpretation of estimates as price effects

We first show how to quantify the results in terms of domestic price changes in the U.S. and then in terms of potential world price changes. The domestic price for a traded product subject to an ad valorem tariff τ can be written as

$$(1) \qquad \ln p_t^d = \ln(1+\tau_t) + \ln p_t^w$$

Now define the pass-through rate from tariffs to domestic prices as

$$(2) \qquad \pi \equiv \Delta\ln p_t^d/\Delta\ln(1+\tau_t) = 1+\Delta\ln p_t^w/\Delta\ln(1+\tau_t)$$

If there is full pass-through to domestic prices, i.e. if $\pi=1$, then the stumbling block parameter, ϕ, represents the growth in the U.S. price for a PTA type good relative to a similar non-PTA good. To see this use (1), (2) and recall that we use $\Delta\ln(1+\tau_t)$ as the dependent variable in (E4) to obtain

$$(3) \qquad \Delta\ln p^d_{PTAt} - \Delta\ln p_t^d = \Delta\ln(1+\tau_{PTAt}) - \Delta\ln(1+\tau_t) = \phi$$

where x_{PTAt} denotes that we are considering the value of variable x when the PTA variable, G, is one. Alternatively we can calculate the ratio of the domestic price growth of a PTA good to a similar non-PTA good using the following

$$(4) \qquad \Delta\ln p^d_{PTAt} /\Delta\ln p_t^d = \Delta\ln(1+\tau_{PTAt})/\Delta\ln(1+\tau_t) = 1+\phi/\tilde{a}$$

where $\tilde{a}$ is the estimated average tariff growth for non-PTA products so the last equality applies to a "benchmark" good with no NTBs, no changes in bargaining power or market access.

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[27] See Hayashi (2000). Testing for the exogeneity of a subset of instruments requires an alternative set, which are considered exogenous a priori. This explains why we also use the average U.S. international transport cost for each good as an excluded instrument. This is a potential predictor for whether a good is exported under the PTA because, *conditional* on a PTA partner exporting a good to the U.S. it will gain more from a preference in that good if there are high average transport costs since this implies that it will face less competition from the rest of the world in the U.S. market. In the first stage the polynomial of the transport cost variable is jointly significant in explaining the PTA variables across different specifications. For the specification with individual PTAs we also use changes in the world price between 1990 and 1992 as an instrument as well as its interaction with transport costs.

We can also provide an interpretation of the results when there is incomplete pass-through, i.e. $\pi<1$. This is important since the motivation for countries to seek reciprocal trade liberalization in a number of models is precisely that a reduction in the tariffs of the importing country increases the price received by the exporter (or maintains the terms-of-trade unchanged as the importer is also reducing its own tariffs).

Kreinin (1961) estimates that less than 1/3 of tariff reductions that the U.S. enacted during the Kennedy trade Round were passed on to U.S. consumers in the form of lower import prices, and thus 2/3 were passed on in the form of higher prices for the exporters.[28] We do not know of comparable estimates of pass-through for the U.S. during the UR. However, we can provide an estimate of the effect on the relative export prices of PTA goods vs. non-PTA goods exported by non-PTA countries. Using (2) to write the world price as a function of the pass-through and tariff changes we obtain

(5) $$\Delta \ln p_t^w = (\pi-1)\Delta \ln(1+\tau_t)$$

Using (5) we can write the ratio of the growth in the world price of a "benchmark" PTA vs. a non-PTA good as

(6) $$\Delta \ln p^w_{PTAt} / \Delta \ln p_t^w = \{\Delta \ln(1+\tau_{PTAt}) / \Delta \ln(1+\tau_t)\} * (\pi_{PTA} - 1)/(\pi - 1)$$

$$\approx 1+\phi/\tilde{a} \quad \text{if } \pi_{PTA} \approx \pi < 1$$

Therefore, we obtain an expression similar to the relative domestic price growth in (4). The expression in (4) applies only if $\pi=1$. The expression in (6) applies if there is imperfect pass-through and we are agnostic about its relative size for two goods in a given industry—that differ only in whether one of them is exported only by the rest of the world. Obviously $1+\phi/\tilde{a}$ underestimates the negative impact on prices for countries that export PTA type goods if the U.S. has more market power in those goods, i.e. if $\pi_{PTA}<\pi$.

5.5 Quantification

In table 3 we quantify the importance of our estimates of the stumbling block both in terms of price effects and in relation to the multilateral negotiation variables. Each row provides an alternative measure and the columns correspond to the IV specifications in table 1.

The first column of table 3 uses the estimates in column 3 of table 1. If we assume $\pi=1$ then the U.S. domestic price of PTA goods increased by 1.28% relative to similar non-PTA goods. The average U.S. domestic price growth for a "benchmark" non-PTA good was approximately -2.67% and therefore the reduction of the domestic price for the PTA good in the U.S. was only 52% of that experienced by the average non-PTA good. Assuming imperfect and similar pass-through we can use (6) to interpret the relative price growth measures in row 2 in the following way. The increase in the export price for a country that does not have a PTA with the U.S. but exports any of the PTA-type goods is only 52% of the price increase for a similar non-PTA good.

In the second column we consider the effect for the goods that are exported by every PTA. We note in particular that the relative price growth is only 23% and that this estimate has a standard error of 13 so we can't reject the hypothesis that the average relative price growth for the goods exported by every PTA

[28] Further evidence of market power is provided by Chang and Winters (2002) who find that the elimination of internal tariffs between Argentina and Brazil caused other countries to reduce their prices on goods exported to Brazil. There is also ample evidence that there is imperfect pass-through from exchange rate changes to domestic prices, about 1/2 according to Goldberg and Knetter's 1997 review, and that the effect is symmetric to that of tariff changes (Feenstra [1989]).

was as low as zero. For important exports for any PTA the relative price growth is a mere 31% as shown in the third column.

The last set of columns in table 3 uses the significant estimates for individual PTAs in column 7 of table 1. The strongest effect is for NAFTA with a 68% relative price growth, followed by GSP, 74%. The estimates for the ATPA and CBI are identical, 84%. Moreover, the upper bound of the 95% confidence interval for these four PTAs is below 100%. So, we can reject the hypothesis that the price effects for PTA and non-PTA goods are similar even for the case of the ATPA and CBI.[29]

To compare the importance of the PTAs relative to the reciprocity effect we consider a country that does not have a PTA with the U.S. but exports the same good as a PTA partner. We then ask how much *more* this country must lower its average tariff in order to obtain the same average tariff reduction by the U.S. in that good as the one received by a country that exports a similar non-PTA good. For the basic specification with any PTA the answer is 91% if that country has an export share close to one. This is a large effect if we consider that the median decrease in the tariff for the exporters used in our sample is 44%, which is about the extra reduction required to overcome the stumbling block effect for NAFTA alone. Countries with lower export shares require proportionally larger decreases in their tariffs to overcome the stumbling block effect.[30]

One of the leading explanations for the relatively high tariffs that small developing countries face in developed countries is that developing countries have not been required to reciprocate and have consequently not received large cuts on products they export. Accounting for reciprocity as we do above implies that this is not the channel the PTA variables are capturing. Moreover, the quantification above suggests that the payoff for small developing countries in lowering their own tariffs may be relatively small, in terms of tariff concessions by the U.S., when compared to the elimination of particular WTO rules that allow PTAs.[31]

We compute a statistic for the bargaining power variable similar to the one for reciprocity. A country's GDP would have had to grow by 91% relative to the U.S. between the Tokyo and Uruguay rounds to overcome the effect from NAFTA, and over 40% to overcome the effect from either the ATPA or CBI. To place this figure in context, the countries in our sample with the highest increase in bargaining power relative to the U.S. were Singapore (51%), E.U. (43%) and Thailand (35%).[32] But of course quite a few of the countries in the sample lost bargaining power relative to the U.S. between the Tokyo and Uruguay Round.

The estimates above provide evidence that the PTAs signed by the U.S. constituted a stumbling block towards its own multilateral trade liberalization and support the basic findings in figures 1 and 2 and. These PTAs may also have constituted a stumbling block for the MTL of other countries via reciprocity.

5.6 Robustness

The main results reported thus far already include a number of specification and robustness tests. Here we report two additional ones: robustness to finer industry controls and an alternative test of our hypothesis that attempts to capture whether there is an important missing variable bias in the estimation.

One possible concern with the results presented thus far is that the industry time-varying dummies are not capturing all the time-varying political economy effects or other industry characteristics that affected the tariff reductions. Therefore we re-estimate the main specifications in table 1 using a richer set

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[29] The reason for the individual PTA estimates of ϕ to be lower than the estimate for any PTA is that several of the goods exported by the smaller PTAs are also exported under NAFTA.

[30] The foreign tariff changes are defined as percentage changes of the tariff not the tariff factor, thus the magnitudes are larger than those for the dependent variable.

[31] Obviously small countries may still gain from unilateral trade liberalization for the standard reasons.

[32] The E.U. refers to the group of countries that belonged to that agreement during both the Tokyo and UR.

of industry time-varying effects that span over 800 4-digit industries in our sample. Some of the interesting variation may be lost by considering only changes within such finely defined industries and in fact we must drop some observations where there are only two or fewer observations per 4-digit industry. However, it is a useful robustness test and it confirms the main results in table 1.

Table 4 contains the results of re-estimating the IV specifications in table 1 with 4-digit industry dummies. The sign and significance of the PTA, bargaining and reciprocity variables are similar to those in table 1, which were obtained using 2-digit industry dummies. This provides further evidence that the coefficients in table 1 are not likely to be biased by political economy effects or other unobserved industry time-varying effects. In terms of model selection the estimates with 4-digit dummies explain a larger fraction of the variation but at the cost of introducing many parameters and, according to the Schwarz criterion, the more parsimonious models with the 2-digit dummies are preferable.[33]

5.6.1 An alternative test

We now provide an alternative robustness test. In the absence of a stumbling block effect we should expect that, within any given industry and conditional on reciprocity and bargaining effects, there should be no *systematic* difference in the change in MFN tariffs for products exported by a given PTA and any *given* combination of one or more non-PTA countries.

To test this hypothesis we can include an additional control group, a country or combination of countries c that do not have a PTA with the U.S. Obviously we can't introduce all countries simultaneously in c since then it would not be possible to identify any effect. Thus we generate various alternative groups of countries and estimate the following equation for each combination c:

$$\text{(E5)} \quad \Delta\tau_{it} = \delta_c G_i + \eta_c G_{ic} + a_c + a_{cI} + \beta_c \Sigma_k s_{iT}^k \Delta(b_t - b^k_t) + \rho_c \Sigma_k s_{iT}^k \Sigma_j \Delta\tau^k_{jt} w^k_{jT} + u_{ic} \qquad i=1,\ldots,N$$

where G_i is the PTA variable and G_{ic} is 1 if any of the countries in the group export to the U.S., 0 otherwise. The null hypothesis is then that $\delta_c = \eta_c$ in each regression. Since we do not know which preferences would be offered if the countries in c had a PTA with the U.S. we assume that it would offer preferences in all goods. This also implies that the only PTA that we can test against each of the combinations c is NAFTA since it is the only one in which all products in our sample receive preferential treatment. However, we can and do control for the other PTAs.

In principle one could use any one or any combination of countries that are not PTA partners. However, we think this test is more powerful if the control countries are on some important dimension not completely different from the PTA partners we test against, Canada and Mexico. For example we would not want to generate multiple estimates reflecting combinations of microstates that export a handful of products to claim that in a majority of cases we confirm or reject that $\delta_c = \eta_c$. Therefore, to construct a reasonable control group, we restrict the countries to satisfy at least one of two criteria. Either they must export at least one product in each of the 20 sectors considered, as Canada and Mexico do, or of the over 5000 tariff lines in the sample they must export at least 266 products, which corresponds to approximately 10% of the 2658 lines that Mexico exports. The criteria are sensible since the identification depends on differences in the products exported within a given industry. The results are not very sensitive to these criteria.

The criteria for the control group leaves us with 23 countries with which the U.S. has no PTA and includes all its major trading partners (see appendix B). Recall that many of the countries are not included because they benefit from preferences under the GSP. To compare with NAFTA we construct all possible combinations of one and 2 countries, thus we obtain 276 combinations and estimate (E5) for each using GMM and instrumenting as in column 6 of table 1.

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[33] This is true even after we adjust the sample in table 1 to be identical to the one in table 4.

We find that the coefficient for NAFTA *exceeds* that for the non-PTA partners in 255 out of the 276 cases. When the point estimate of the difference is negative it is not statistically significant. The best way to describe the difference for all the estimates and whether it is significant is to plot the z-statistic for δ_c-η_c.[34] Figure 3 plots the "smoothed" distribution. If there were no systematic differences then the distribution should be centered around zero. Clearly the majority of the estimates are positive and significant. Therefore this test provides further support that the U.S. did not reduce its MFN tariffs on products imported under NAFTA by as much as on other products not imported from its PTA partners.[35]

The level of the NAFTA coefficient, δ_c, is also informative. It is always positive and significant and its median value for all the different specifications is 0.74, which is within less than a standard error of the value estimated in table 1, column 6, 0.799. These estimates for the level of the PTA coefficients suggest that the estimates in table 1 are not simply capturing an omitted variable that is positively correlated with both the PTA variables and the export vector of most countries exporting to the U.S.[36]

6 Conclusion

Are preferential agreements a stumbling block for multilateral liberalization? This question has been of central concern to economists and politicians in the last 10 years. We provide the first estimates to show that this concern is well founded for the U.S. The U.S. multilateral tariff reductions in PTA goods were smaller than those in similar goods not imported from PTA partners. On average an exporter to the U.S. that did not belong to one of its PTAs received about 52% the benefit (in terms of price increases) if it exported any PTA good instead of a similar non-PTA good. This effect is even stronger if the good was exported by all PTAs or was an important export for a PTA partner.

The stumbling block effect is present even for agreements with small countries, which is important for two reasons. First, there are many more PTAs that the U.S. can still sign with small countries. Second, the PTAs with small countries affect products exported by other small developing countries to the U.S. These are countries relative to which the U.S. is likely to have market power and thus smaller reductions in the U.S. tariffs affects their export price adversely.

In estimating the stumbling block effect we also provide new estimates of tariff reciprocity in multilateral negotiations solving the potential endogeneity problem associated with it. We confirm that reciprocity is indeed present for products that were not subject to non-tariff barriers. That is, the U.S. tariff reductions are relatively higher in products for which the main exporters also offered large tariff reductions. If other countries followed reciprocity then the PTAs signed by the U.S. had an added effect. Smaller U.S. tariff reductions were reciprocated by smaller reductions on U.S. products by certain partners.

The prediction we estimate is motivated by a specific theoretical model, the assumptions of which are closely matched by the ATPA and CBI. However, we argue that a similar model, which includes in addition trade diversion and creation effects is also likely to lead to testable predictions similarly based on product level variation. Ideally future work will focus on testing such a model against alternative explanations. However, given the absence of evidence thus far we believe that our estimates provide an important starting point. We are currently analyzing if there are similar effects for the E.U.

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[34] We estimate (E5) in a modified form where we use $\delta_c G_i + \eta_c G_{ic} = (\delta_c - \eta_c) G_i + \eta_c (G_{ic} + G_i)$ so we test $\delta_c = \eta_c$ directly using a standard z-statistic.

[35] The results are not very sensitive to changes in the criteria for the control group. If we include all combinations of 1 or 2 countries that export at least 133 goods (5% of the number exported by Mexico) or goods in at least 10 of the sections then 277 out of 300 estimates for the difference are positive and the negative ones are insignificant. If we require countries to export at least 665 goods (20% of the number exported by Mexico) or goods in at least 20 of the sections then 252 out of 273 estimates for the difference are positive and the negative ones are insignificant.

[36] We also include the variable for PTAs other than NAFTA. That coefficient is also always positive and significant and its median value is also within less than a standard error of the 0.71 estimate in table 1.

Finally, our approach has the following application. With a new trade round under way it is important to verify that the behavior of the Uruguay Round is not repeated; both because the stumbling effect is potentially harmful to a number of countries and because it directly violates a WTO decision from 1999 that "Any preferential tariff treatment implemented pursuant to this Waiver [of the MFN rule for the purpose of preferential tariff treatment for LDCs…] *shall not constitute an impediment to the reduction or elimination of tariffs on a most-favored-nation basis*."[37] This was precisely the waiver that the U.S. obtained to implement the ATPA for example. One simple way to ensure that this does not occur in the coming round is to use the empirical model in this paper to evaluate tariff concession proposals.

[37] Emphasis added. In WT/L/304, p. 2, accessed April 2003 on <http://www.wto.org/english/docs_e/legal_e/waiver1999_e.pdf>.

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Appendix A

Table 1

Stumbling block and multilateral negotiation effects (E4)

	OLS		IV-GMM				
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
ANYPTA† (ϕ)	0.982***	0.982***	1.28***	1.33***	1.21***		
	(0.204)	(0.204)	(0.328)	(0.321)	(0.331)		
EVRPTA†		0.004		0.778***			
		(0.139)		(0.219)			
ANYPTAxHI_EX †					0.584***		
					(0.110)		
NONAFTA†						0.710***	
						(0.209)	
NAFTA‡						0.799***	0.815***
						(0.149)	(0.149)
ATPA†							0.398***
							(0.151)
CBI†							0.381***
							(0.141)
GSP†							0.658***
							(0.220)
ISR†							0.010
							(0.097)
BARPOW (β)	0.009***	0.009***	0.009***	0.009***	0.008***	0.009***	0.009***
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
TOTLIB† (ρ)	-0.004	-0.004	0.014**	0.015**	0.017***	0.017***	0.018***
	(0.005)	(0.005)	(0.006)	(0.006)	(0.006)	(0.006)	(0.006)
TOTxNTB†	-0.012	-0.012	-0.071***	-0.070***	-0.072***	-0.071***	-0.072***
	(0.008)	(0.008)	(0.011)	(0.011)	(0.011)	(0.011)	(0.011)
NTB	-0.914*	-0.914*	-4.02***	-3.98***	-4.11***	-3.96***	-3.88***
	(0.480)	(0.480)	(0.676)	(0.668)	(0.676)	(0.673)	(0.670)
DS	-0.635***	-0.635***	-0.676***	-0.708***	-0.748***	-0.704***	-0.753***
	(0.168)	(0.168)	(0.170)	(0.170)	(0.170)	(0.166)	(0.169)
Constant (*a*)	-2.94***	-2.94***	-2.26***	-2.41***	-2.24***	-2.28***	-2.42***
	(0.401)	(0.401)	(0.478)	(0.474)	(0.479)	(0.431)	(0.429)
Observations	5079	5079	5079	5079	5079	5079	5079
Adj. R-sq.	0.18	0.18	-	-	-	-	-
Schwarz criterion	-	-	2.29	2.30	2.29	2.29	2.30
Hansen's J p [1]	-	-	0.64	0.85	0.73	0.54	0.56
C-stat p [2]	-	-	0.39	0.71	0.66	0.49	0.60
Exogeneity test p [3]	-	-	0.00	0.00	0.00	0.00	0.00
Homosk. p [4]	-	-	0.00	0.00	0.00	0.00	0.00
No. of parameters	98	99	98	99	99	99	102

Standard errors in parentheses. HS 2 digit dummies included but not reported. * Significant at 10%; ** significant at 5%; *** significant at 1%.

1. Test of over-identifying restrictions. Probability at which we reject the null hypothesis that the excluded instruments are uncorrelated with the error term, and correctly excluded from the estimated equation.

2. Probability for rejection of the null hypothesis of exogeneity of a subset of *f* instruments marked with "‡" in tables 1 and 2. C-stat = $J^r - J^u \sim \chi_f^2$ where J^r represents the minimized value of the GMM objective function for the restricted and efficient regression (with all overidentifying restrictions) and J^u the value for the unrestricted, inefficient but consistent regression without the questionable instruments.

3. Regression based Hausman specification test for the endogeneity of the variables marked with a "†". Probability value at which we reject the consistency and efficiency of OLS.

4. Identical result for three versions of the Pagan-Hall homoskedasticity test using the levels of the instruments, the "fitted value" of the dependent variable and its square.

Table 2a

First stage regressions

Specification:	Table 1 (3)			Table 1 (6)		
Dependent variable:	ANYPTA	TOTLIB	TOTxNTB	NONAFTA	TOTLIB	TOTxNTB
ANYEXP‡	0.908***	1.23***	0.492*			
	(0.013)	(0.348)	(0.269)			
ANYEXPxNTBALL‡	-0.009	0.613	1.16			
	(0.128)	(3.34)	(2.58)			
EX_NONAFTA‡				0.684***	0.234	0.141
				(0.014)	(0.230)	(0.177)
EX_NONAFTAxNTBALL‡				-0.090	-1.24	-0.705
				(0.119)	(1.90)	(1.47)
NAFTA‡				0.082***	0.669***	0.725***
				(0.014)	(0.225)	(0.173)
NTB_ALL‡	0.051	2.97	3.09	0.158	4.37***	4.65***
	(0.125)	(3.26)	(2.52)	(0.104)	(1.67)	(1.29)
BARPOW	0.001***	0.224***	0.093***	0.001***	0.224***	0.093***
	(0.000)	(0.005)	(0.004)	(0.000)	(0.005)	(0.004)
UNILIB	-0.000	0.694***	0.068***	0.000	0.695***	0.070***
	(0.000)	(0.007)	(0.005)	(0.000)	(0.007)	(0.005)
UNILIBxNTB	0.001**	0.072***	0.592***	0.000	0.070***	0.591***
	(0.000)	(0.011)	(0.009)	(0.001)	(0.011)	(0.009)
NTB	0.007	1.87***	-35.3***	-0.177***	1.81***	-35.3***
	(0.017)	(0.454)	(0.351)	(0.028)	(0.455)	(0.351)
DS	-0.007	-0.446	-0.418	-0.038	-0.471	-0.452
	(0.015)	(0.394)	(0.304)	(0.025)	(0.394)	(0.304)
TC	0.344	8.56	1.91	1.122***	8.15	0.464
	(0.252)	(6.56)	(5.07)	(0.411)	(6.60)	(5.09)
TCsq	-2.71**	-34.6	-15.5	-4.81**	-31.7	-7.32
	(1.34)	(34.8)	(26.9)	(2.18)	(35.0)	(27.0)
TCcb	3.03*	44.9	40.7	5.01*	40.9	32.3
	(1.71)	(44.5)	(34.4)	(2.78)	(44.7)	(34.5)
Constant	0.089***	-26.7***	4.49***	0.184***	-26.3***	4.29***
	(0.030)	(0.792)	(0.612)	(0.047)	(0.761)	(0.587)
Observations	5079	5079	5079	5079	5079	5079
Adj. R-sq.	0.56	0.77	0.97	0.52	0.77	0.97
Partial R-sq.	0.49	0.72	0.59	0.32	0.72	0.59
F-test p	0.00	0.00	0.00	0.00	0.00	0.00

Standard errors in parentheses. HS 2 digit dummies and interactions of transport costs and price changes included but not reported. * Significant at 10%; ** significant at 5%; *** significant at 1%. ‡ Instruments tested for exogeneity, see C-stat in table 1.

1. P-value of the F-test of joint significance of the instruments in first-stage regression

Table 2b
First stage regressions

Specification	Table 1 (7)					
Dependent variable:	ATPA	CBI	ISR	GSP	TOTLIB	TOTxNTB
EX_ATPA ‡	0.690***	-0.001	-0.007**	0.014	0.343	0.349**
	(0.007)	(0.008)	(0.003)	(0.013)	(0.222)	(0.171)
EX_CBI ‡	0.008	0.723***	0.001	0.015	0.221	0.654***
	(0.007)	(0.007)	(0.003)	(0.012)	(0.212)	(0.163)
EX_GSP ‡	-0.000	0.000	-0.002	0.585***	0.096	0.084
	(0.007)	(0.008)	(0.003)	(0.013)	(0.219)	(0.168)
EX_ISR ‡	0.009	0.006	0.986***	0.037***	0.290	0.200
	(0.007)	(0.007)	(0.003)	(0.011)	(0.196)	(0.151)
EX_ATPAxNTBALL ‡	0.035	0.092	-0.004	0.067	1.87	2.54
	(0.075)	(0.078)	(0.030)	(0.132)	(2.27)	(1.74)
EX_CBIxNTBALL ‡	0.089	0.075	0.035	-0.077	0.370	0.220
	(0.069)	(0.072)	(0.027)	(0.120)	(2.074)	(1.60)
EX_GSPxNTBALL ‡	-0.061	-0.133*	-0.041	-0.417***	-1.86	-1.62
	(0.069)	(0.072)	(0.027)	(0.120)	(2.07)	(1.590)
EX_ISRxNTBALL ‡	0.018	0.143	-0.017	0.417**	-2.93	-2.82
	(0.114)	(0.118)	(0.045)	(0.199)	(3.42)	(2.63)
NAFTA ‡	-0.011	-0.017**	-0.001	0.008	0.588***	0.578***
	(0.008)	(0.008)	(0.003)	(0.013)	(0.228)	(0.175)
NTB_ALL ‡	0.082	0.145***	0.057***	0.367***	4.38***	4.80***
	(0.051)	(0.053)	(0.020)	(0.090)	(1.54)	(1.19)
BARPOW	-0.001***	-0.001***	0.000	0.001***	0.223***	0.091***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.005)	(0.004)
UNILIB	-0.001***	-0.001***	0.000	-0.001*	0.695***	0.069***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.007)	(0.005)
UNILIBxNTB	0.001***	0.002***	0.000	0.001	0.073***	0.594***
	(0.000)	(0.000)	(0.000)	(0.001)	(0.011)	(0.009)
NTB	-0.032**	-0.028*	0.004	-0.268***	1.840***	-35.3***
	(0.015)	(0.016)	(0.006)	(0.027)	(0.456)	(0.351)
DS	0.000	0.030**	-0.005	-0.041*	-0.521	-0.532*
	(0.013)	(0.014)	(0.005)	(0.023)	(0.396)	(0.304)
TC	0.499**	0.654***	0.217**	0.896**	3.03	-2.57
	(0.232)	(0.241)	(0.091)	(0.405)	(6.98)	(5.37)
TCsq	-0.322	-1.62	-1.012**	-2.39	-9.44	-0.081
	(1.260)	(1.31)	(0.495)	(2.20)	(37.9)	(29.1)
TCcb	-0.574	1.38	1.04	2.71	18.9	29.1
	(1.66)	(1.72)	(0.650)	(2.89)	(49.8)	(38.3)
PCH9092	0.002	-0.005	0.002	0.032*	-0.099	-0.073
	(0.010)	(0.010)	(0.004)	(0.017)	(0.294)	(0.226)
PCH9092sq	-0.000	0.003	0.001	-0.004	-0.157	-0.098
	(0.004)	(0.004)	(0.001)	(0.006)	(0.111)	(0.086)
Constant	0.030	0.021	0.000	0.267***	-26.2***	4.16***
	(0.025)	(0.026)	(0.010)	(0.044)	(0.765)	(0.588)
Observations	5079	5079	5079	5079	5079	5079
Adj. R-sq.	0.74	0.77	0.98	0.60	0.77	0.97
Partial R-Sq.	0.68	0.70	0.97	0.33	0.72	0.59
F-test IV p [1]	0.00	0.00	0.00	0.00	0.00	0.00

Standard errors in parentheses. HS 2 digit dummies and interactions of the transport cost and price change variables included but not reported. * Significant at 10%; ** significant at 5%; *** significant at 1%. ‡ Instruments tested for exogeneity, see C-stat in table 1.
1. P-value of the F-test of joint significance of the instruments in first-stage regression

Table 3
Quantification of the stumbling block effect

%	Any PTA [a]	Any & every PTA [b]	Any PTA w/ HI_EX [c]	Individual PTAs[d] NAFTA	ATPA	CBI	GSP
Growth of relative U.S. price of PTA goods if p =1, ϕ	1.28	2.11	1.79	0.82	0.40	0.38	0.66
	(0.33)	(0.36)	(0.32)	(0.15)	(0.15)	(0.14)	(0.22)
Relative price growth[e], $1+\phi/\tilde{a}$	52	23	31	68	84	84	74
	(10)	(13)	(11)	(6.6)	(6.3)	(6.1)	(8.7)
Foreign tariff reduction equivalent[f], ϕ/ρ	91	140	104	45	22	21	37
	(39)	(59)	(38)	(16)	(11)	(9.9)	(16)
Foreign bargaining power equivalent[g], ϕ/β	142	234	233	91	44	42	73
	(46)	(60)	(76)	(27)	(20)	(19)	(30)

(a) table 1 (3); (b) table 1 (4), joint effect of ANYPTA and EVRPTA; (c) table 1 (5), joint effect of ANYPTA and ANYPTAxHI_EX; (d) table 1 (7); (e) Represents relative U.S. domestic price effect of preferential vs. non-preferential good if π=1 and world price effect if $\pi_{PTA}\approx\pi<1$. $\tilde{a} \equiv a + (\Sigma_{I\in\iota}a_I)/69$: the estimated average tariff factor growth for non-PTA products over the set ι of 69 industries that contained at least one of them; (f) Average tariff reduction required by an exporter of a PTA type good with share of 1 in order to overcome the preferential effect. If the exporter has a share s then the effect is $\phi/s\rho$; (g) GDP growth required by an exporter of a PTA type good with share of 1 in order to overcome the preferential effect.If the exporter has a share s then the effect is $\phi/s\beta$.
Standard errors in parenthesis obtained using the delta method for the non-linear parameter combinations in the last three rows.

Table 4
Robustness (E4)

	(1)	(2)	(3)	(4)	(5)
ANYPTA † (ϕ)	1.31***	1.32***	1.27***		
	(0.317)	(0.306)	(0.319)		
EVRPTA †		1.26***			
		(0.249)			
ANYPTAxHI_EX†			0.366***		
			(0.111)		
NONAFTA †				0.702***	
				(0.235)	
NAFTA ‡				0.765***	0.755***
				(0.151)	(0.147)
ATPA †					0.566***
					(0.167)
CBI †					0.439***
					(0.158)
GSP †					0.767***
					(0.248)
ISR †					0.098
					(0.100)
BARPOW (β)	0.007***	0.007***	0.006***	0.006***	0.006***
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
TOTLIB† (ρ)	0.015**	0.015**	0.016***	0.016***	0.019***
	(0.006)	(0.006)	(0.006)	(0.006)	(0.006)
TOTxNTB †	-0.057***	-0.056***	-0.057***	-0.056***	-0.059***
	(0.011)	(0.011)	(0.011)	(0.011)	(0.011)
NTB	-2.46***	-2.45***	-2.50***	-2.44***	-2.40***
	(0.696)	(0.693)	(0.696)	(0.696)	(0.694)
DS	-0.401	-0.493**	-0.480*	-0.486**	-0.560**
	(0.248)	(0.251)	(0.250)	(0.246)	(0.249)
Constant	-5.61***	-5.59***	-5.51***	-5.46**	-5.35**
	(2.097)	(2.10)	(2.100)	(2.19)	(2.215)
Observations	4636	4636	4636	4636	4636
Schwarz criterion	2.85	2.86	2.85	2.85	2.86
Hansen's J p [1]	0.630	0.695	0.70	0.311	0.619
C-stat p [2]	0.491	0.249	0.67	0.284	0.556
Exogeneity test p [3]	0.00	0.00	0.00	0.00	0.00
Homosk. p [4]	0.00	0.00	0.00	0.00	0.00
No. of parameters	546	547	547	547	555

Standard errors in parentheses. HS 4 digit dummies included but not reported. * Significant at 10%; ** significant at 5%; *** significant at 1%. Sample reduced because of dropped HS4 categories with only 2 or less observations. Endogenous instrumented variables denoted with "†". First stage regressions available on request.

1. Test of over-identifying restrictions. Probability at which we reject the null hypothesis that the excluded instruments are uncorrelated with the error term, and correctly excluded from the estimated equation.

2. Probability for rejection of the null hypothesis of exogeneity of a subset of f instruments marked with "‡". C-stat = $J^r - J^u \sim \chi_f^2$ where J^r represents the minimized value of the GMM objective function for the restricted and efficient regression (with all overidentifying restrictions) and J^u the value for the unrestricted, inefficient but consistent regression without the questionable instruments.

3. Regression based Hausman specification test for the endogeneity of the variables marked with a "†". Probability value at which we reject the consistency and efficiency of OLS.

4. Identical result for three versions of the Pagan-Hall homoskedasticity test using the levels of the instruments, the "fitted value" of the dependent variable and its square.

Appendix B

Table 1b: Summary statistics

Variable	Mean	St. dev.	Min	Max
USLIB	-2.89	3.21	-32.2	3.17
ANYPTA	0.891	0.312	0	1
EVRPTA	0.049	0.216	0	1
ANYPTAxHI_EX	0.197	0.398	0	1
ANYEXP	0.941	0.236	0	1
EVREXP	0.091	0.287	0	1
HI_EX	0.245	0.43	0	1
NONAFTA	0.605	0.489	0	1
NAFTA	0.826	0.379	0	1
ATPA	0.145	0.352	0	1
CBI [1]	0.186	0.389	0	1
GSP [1]	0.467	0.499	0	1
ISR	0.279	0.449	0	1
EX_NAFTA	0.826	0.379	0	1
EX_ATPA	0.228	0.420	0	1
EX_CBI [1]	0.276	0.447	0	1
EX_ISR	0.285	0.451	0	1
EX_GSP [1]	0.807	0.395	0	1
EX_NONAFTA	0.840	0.367	0	1
EX_ATPAxNTBALL	0.002	0.044	0	1
EX_CBIxNTBALL	0.004	0.059	0	1
EX_ISRxNTBALL	0.001	0.028	0	1
EX_GSPxNTBALL	0.006	0.075	0	1
BARPOW	-30.9	21.5	-70.5	93.0
TOTLIB	-51.3	11.3	-89.5	-0.593
TOTxNTB	-16.2	24.4	-81.9	0
UNILIB	-28.0	15.8	-83.6	-0.115
UNILIBxNTB	-8.68	15.2	-80.6	0
NTB	0.324	0.468	0	1
NTB_ALL	0.009	0.095	0	1
ANYEXPxNTBALL	0.008	0.092	0	1
EVREXPxNTBALL	0.000	0.020	0	1
HI_EX xNTBALL	0.008	0.092	0	1
TC	0.073	0.048	0.003	0.737
TCsq	0.008	0.016	0.000	0.543
TCcb	0.001	0.008	0.000	0.400
TCxPCH9092	0.007	0.063	-0.978	0.967
TCsqxPCH9092	0.001	0.015	-0.292	0.387
TCxPCH9092sq	0.034	0.200	0.000	6.65
TCsqxPCH9092sq	0.004	0.033	0.000	0.9571
PCH9092	0.097	0.678	-6.4023	9.26
PCH9092sq	0.469	2.51	0.000	85.7
DS	0.047	0.213	0	1

Number of observations: 5079. There were 7853 unique 8-digit tariff lines for which the WTO tariff data was available before and after the round once the special sections (21 and 22) are excluded. The sample is reduced once we exclude the lines for which the MFN tariff was zero before the round (n=6800). Lines with missing or zero imports are dropped because the imports are required to construct TOT and BARPOW (n=5227). Missing data on transport costs and preferences further reduce the sample to 5079.

1. Officially the members of the ATPA are also beneficiaries under the CBI and GSP. However, we do not include them in the construction of the CBI and GSP variables both because the benefits under the ATPA are more generous and to minimize multicollinearity problems.

Table 2b
Data description

NAME	Model (E4)	Data/description	SOURCE
U.S.LIB	$\Delta\tau_{it}$	$\ln(1+\tau_{it})$ - $\ln(1+\tau_{it-1})$; τ_{it}: U.S. Ad valorem bound rate (post-UR). τ_{it-1} U.S. Ad valorem base rate (pre-UR)	WTO
ANYPTA	G_i	=1 if exported to the U.S. under any preferential trade agreement before UR implemented (1994)	NBER/Feenstra et al 2002
EVRPTA		=1 if exported to the U.S. under all its preferential trade agreements (1994)	" "
PTA (NAFTA, etc)		=1 if exported to the U.S. under each of the respective PTAs (1994)	" "
ANYPTAxHI_EX		=1 if exported to the U.S. under any preferential trade agreement before UR implemented (1994) and export value of good i from each given PTA exceeds the mean export value of that PTA to the US in all goods.	" "
NONAFTA		=1 if exported to the U.S. under any preferential trade agreement except NAFTA (1994)	" "
ANYEXP		=1 if exported to the U.S. by any PTA partner whether or not the product received a preference (1994)	" "
EVREXP		=1 if exported to the U.S. by all PTA partners whether or not the product received a preference (1994)	" "
EX_NONAFTA		=1 if exported to the U.S. by any PTA partner, other than NAFTA countries, whether or not the product received a preference (1994)	" "
EX_PTA		=1 if exported to the U.S. by each of the respective PTA partners whether or not the product received a preference (1994)	" "
s_{iT}^{k}		Partner k's export share of top 5 exporters of i to U.S.	" "
BARPOW	$\Sigma_k s_{iT}^k\Delta(b_t-b^k_t)$	$\Sigma_k s_{iT}^k\Delta(\ln GDP^{U.S.}_t - \ln GDP^k_t)$	PWT 6.1
TOTLIB	$\Sigma_k s_{iT}^k \Sigma_j \Delta\tau^k_{jt} w^k_{jT}$	$\Sigma_k s_{iT}^k (\Sigma_j \Delta\tau^k_{jt} w^k_{jT})$	
UNILIB		$\Sigma_k s_{iT}^k (\Sigma_j \Delta\tau^k_{jt} - \Sigma_j \Delta\tau^{rk}_{jt}) w^k_{jT}$	
$\Sigma_j \Delta\tau^k_{jt} w^k_{jT}$		where $\Delta\tau^k_{jt} \equiv$ % change 95-86 ~ $\ln(\tau^k_{j95}/\tau^k_{j86})$ $w^k_{jT} \equiv$ Share of good j in k's imports	Finger et al 1999
$\Sigma_j \Delta\tau^{rk}_{jt} w^k_{jT}$		where $\Delta\tau^{rk}_{jt}$ º % change 95-92 ~ $\ln(\tau^k_{j95}/\tau^k_{j92})$	Finger et al 1999
NTB		=1 for HS 8-digit product *i* if the U.S. had any NTB on it applied to *any* country in 1993	UNCTAD TRAINS
NTB_ALL		=1 for HS 8-digit product *i* if the U.S. had any NTB on it applied to *all* countries in 1993	UNCTAD TRAINS
TC, TCsq, TCcb		Average transport cost factor for HS 8-digit product *i* over all exporters of *i* to the U.S. in 1994. TCsq=TC^2, TCcb=TC^3	U.S.ITC
PCH9092, PCH9092sq		Average price change for HS 8-digit product *i* over all exporters of *i* to the U.S. from 90 to 92. PCH9092sq= PCH9092^2	NBER/Feenstra et al 2002
DS		=1 if 8-digit line is subdivided	WTO
VAR1xVAR2		VAR1*VAR2	

Abbreviations: PWT, Penn World Tables; U.S.ITC, United States International Trade Commission; WTO, World Trade Organization; UNCTAD, United Nations Conference on Trade and Development.

Table 3b: Non-PTA control countries for G_{ic} in (E5)

Criteria satisfied (>265 products, or 20 sections)
(528,20) Greece; (777,20) New Zealand; (834,19) Norway; (961,20) Portugal; (999,20) Singapore; (1000,20) Finland; (1131,19) Ireland; (1348,20) Denmark; (1521,20) Australia; (1604,19) Austria; (1814,20) Sweden; (2050,20) Spain; (2087,20) Belgium-Lux.; (2347,20) South Korea; (2362,20) Netherlands; (2566,20) Switzerland; (2783,20) Taiwan; (2898,20) China; (3428,20) Italy; (3541,20) France; (3737,20) Japan; (3878,20) United Kingdom; (3963,20) Germany
Criteria not satisfied
(0,0) St. Pierre; (0,0) Falklands; (1,1) Mauritania; (1,1) Tajikistan; (1,1) Yugoslavia; (4,4) Gabon; (4,2) Fr. Guyana; (4,3) Uzbekistan; (5,4) Iran; (5,3) Azerbaijan; (7,3) Sudan; (8,5) Liberia; (9,6) Afghanistan; (9,6) Turkmenistan; (11,3) Algeria; (12,3) Laos; (12,7) Georgia; (13,5) Armenia; (15,6) Cambodia; (15,5) Moldova; (27,11) Gauadalupe; (37,6) Mongolia; (38,9) Kuwait; (42,3) Qatar; (51,14) Bermuda; (59,11) Burma; (89,14) Vietnam; (90,13) Syria; (92,17) Iceland; (92,12) Belarus; (108,17) Nigeria; (125,16) Saudi Arabia; (208,16)United Arab Emirate

Note: (#,#) are respectively the number of HS 8-digit products and sections the country exported to the U.S. in our sample.

Figure 1
Reductions in U.S. average MFN tariff factors in Uruguay Round by sector (%)

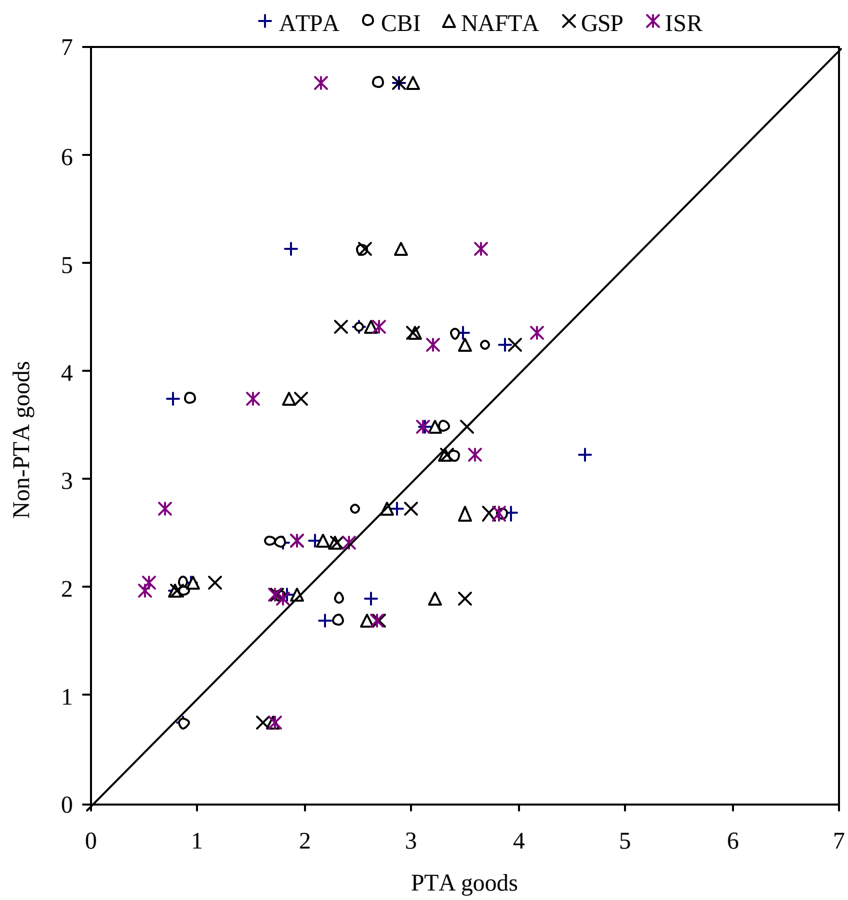

Figure 2

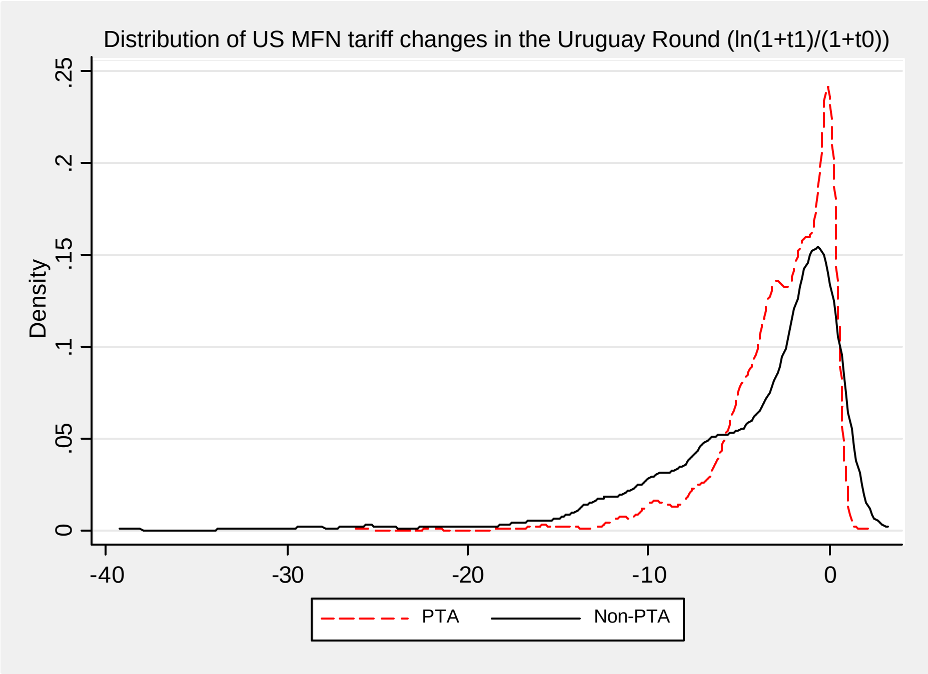

Figure 3