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CESifo Working Paper No. 3985
Category 8: Trade Policy
October 2012

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Abstract

If a free trade agreement (FTA) is characterized by the exchange of market access with a large and competitive trading partner, the agreement can cause a leakage of protectionist benefits to domestic industry from lobbying against external tariff cuts. This rent destruction effect of an FTA can free policy makers to be more aggressive in multi-lateral tariff cuts. We argue that the Canadian-US free trade agreement (CUSFTA) provides an ideal policy experiment to link this mechanism to the data. Exploring the determinants of Canada‘s tariff cuts at the 8 digit HS product level, we find that CUSFTA acted as an additional driver of Canadian multilateral tariff reductions during the Uruguay Round.

JEL-Code: F130, F140.

Keywords: preferences, rent destruction, multilateral liberalization, CUSFTA.

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October 2012
The authors are grateful to Caroline Freund and Emanuel Ornelas for stimulating discussions and comments to earlier version of this paper. They also acknowledge the valuable suggestions of Sohaib Shahid and other participants of the Eleventh GEP Postgraduate Conference in Nottingham. The authors are grateful for financial support from the Leverhulme Trust under Programme Grant F/00 114/AM.

1. Introduction

The substantial increase in recent decades in the number of preferential trade agreements (PTAs) has raised concerns regarding the impact of these agreements on the implementation of multilateral trade liberalization (MTL).[1] This has resulted in a vivid academic debate of whether PTAs are a "building block" or a "stumbling block" to MTL. While the theoretical literature has suggested different channels through which PTAs can affect MTL, the empirical literature is still in its infancy. [2] The few empirical studies have generated contrary findings. Limão (2006) and Karacaovali and Limão (2008) have found empirical evidence for a "stumbling block" effect of PTAs signed by the US and the EU, whereas the results in Estevadeordal et al. (2008) suggest a "building block" effect of PTAs signed in South America. Because these studies take different empirical identification strategies, it is not quite clear whether the contrary results are driven by differences in methodologies or by differences in the underlying policy environments.

Our paper provides theory-based evidence for the view that the effect of a PTA on MTL depends on the motivation for the preferential trade agreement. Employing an identification strategy similar to Karacaovali and Limão (2008), but applying it to Canada's multilateral tariff cuts following the Canadian US free trade agreement (CUSFTA), we find empirical evidence for a building block effect of CUSFTA. We rationalize our finding by the 'exchange of market access objective' of CUSFTA.[3] Assuming that a PTA is motivated by the exchange of market access, Ornelas (2005a) has theoretically shown that a PTA will cause a *leakage* of protectionist benefits to domestic import-competing industries from lobbying against the reduction of external tariffs. As a result, this so-called *rent destruction effect* of a PTA will free policy makers to be more aggressive in multilateral tariff cuts than in the absence of the PTA.

Our paper provides empirical evidence for a new welfare channel of the Canadian-US free trade agreement: the neutralization of inefficiencies created by lobbying activities. This welfare channel of trade agreements goes back to the seminal papers by Maggi and Rodriguez-Clare (1998) and Mitra (2002), but only in a two-country setting. Ornelas (2005a and 2005b) extends the Maggi-Rodriguez-Clare-Mitra rationale to discriminatory liberalization through the

[1] According to the WTO, the number of PTA notifications amounted to 124 in the period 1948-1994. This number increased to over 300 in the time period 1995-2011 (www.wto.org).
[2] Freund and Ornelas (2010) provide a recent review of the PTA literature.
[3] CUSFTA entered into force in 1989 and led to a step-wise phasing out of almost all tariffs between the latter two contracting parties (Romalis, 2007). In 1994 the North American Free Trade Agreement (NAFTA) mainly extended CUSFTA preferences to Mexico.

rent destruction effect. In this setting, the rent destruction feature of CUSFTA has two effects. First, efficiency gains from a lower tariff against non-member countries. Second, since the leakage of protectionist benefits from lobbying is expected to reduce the amount of firm investment in wasteful lobbying, it frees funds for investment in productive capacity. The second effect complements the recent finding of Lileeva and Trefler (2010) that the Canadian-US trade agreement caused Canadian firms to invest in productivity. Our empirical findings are suggestive of a channel where some of the funds for these investments might have come from.

Our theoretical framework is embedded in the modern political economy literature of protection, which goes back to the seminal work by Hillman (1982) and Grossman and Helpman (1994,1995). This literature recognizes that government objectives are affected by campaign contributions of rent-seeking import-competing industries. Extending Grossman and Helpman (1995) by allowing a PTA to affect external tariff formation, Ornelas (2005a) identified the rent destruction effect of a PTA, resulting in a lower external tariff than in the absence of a PTA. Section 2 spells out the main features and assumptions of the underlying theoretical mechanism.

A thorny issue in the empirical literature on the effects of PTAs on MTL is the problem of reverse causality. Section 3 reviews the main features of CUSFTA and argues that it provides a clearly defined policy experiment with the exchange of market access being the prime motive for the agreement. Specifically, policy decisions regarding CUSFTA can be viewed as reasonably exogenous to policy decisions regarding tariff settings during the Uruguay Round.

Causal inference pertaining to the effect of a preferential arrangement (FTA in the present context) on the determination of external tariffs needs to involve some counterfactual reasoning. Since external tariff negotiations are never observed in the presence and the absence of an FTA, we use tariff changes on non-FTA goods as the counterfactual for tariff changes in the absence of an FTA.[4] The identification stems from 8-digit variation in tariff changes in non-FTA goods versus FTA goods, where FTA goods are the subset of goods that a country imports under its FTA. Section 4 of our paper applies this methodology to examine the impact of CUSFTA on Canada's multi-lateral tariff cuts during the Uruguay Round. Using a variety of specifications we find that Canadian preferences under CUSFTA had a statistically and economically significant effect on Canada's tariff reductions during the Uruguay Round. More

[4] To best of our knowledge, this identification strategy was first suggested by Limão (2006).

ambitious tariff cuts of on average 2.2 percentage points on FTA goods relative to non-FTA goods provide strong evidence for Ornelas's (2005a) rent destruction effect.[5]

Our findings are compatible with previous research which provides evidence for preferences promoting external tariff liberalization in Latin America and Asia (cf. Bohara et al., 2004; Estevadeordal et al., 2008; and Calvo-Pardo et al., 2010). Our results are, however, in contrast to Limão (2006) and Karacaovali and Limão (2008) who find the opposite when focusing on the US and the EU. While Karacaovali and Limão (2008) consider a theoretical framework in which a home country grants tariff preferences towards a smaller trading partner in exchange for cooperation agreements in non-trade areas (e.g. environment, immigration, drug trafficking, etc.) - which is the case for a substantial number of PTAs concluded by the US and the EU, our results are based on a PTA-framework of reciprocal market-access with a large and highly developed trading partner.[6] Given the use of a common empirical methodology, the differences in results indicate that it is important to recognise the differences in PTA-settings when analysing the impact of trade preferences on multilateral tariff cuts.

2. Analytical Framework

Our conceptual framework is based on Ornelas (2005a) who showed that a Free Trade Agreement (FTA) can cause leakage of protectionist benefits to import competing industries which may moderate the role of political economy forces and ultimately lead to lower external tariffs than in the absence of an FTA. In what follows we sketch the main features of the model focusing on the key assumptions, underlying mechanism and prediction as relevant to our empirical implementation and refer the interested reader to Ornelas (2005a) for a detailed discussion of the underlying theory. Although Ornelas (2005a) is cast in a specific-factor model, the rent-destruction effect has shown to hold assuming an oligopolistic market structure (Ornelas 2005b) or assuming that countries cooperate multilaterally (Ornelas, 2008). This

[5] It is worthwhile noticing that Trefler (2004) finds also significant short-term adjustment costs for the Canadian economy but long-run gains following the formation CUSFTA. Both aspects tend to support Ornelas's (2005) rent destruction argument as well as his conclusions regarding the political viability of (only) welfare improving FTAs (cf. Ornelas, 2005a).

[6] Both, the EU and the US, have formed numerous trade agreements with smaller trading partners which all include cooperation requirements in certain areas such as intellectual property enforcement, democracy, human rights, labour standards or deeper integration issues. Examples for the EU include the MED, GSP and ACP, preferential trading schemes, whereas PTAs with Andean (ATPA), Caribbean (CBI) and GSP countries may be cited for the US. Canada, however, differs from the latter PTA-setting in the sense that it formed a fully-fledged mutual market-access based FTA with a much larger and highly developed economy-i.e. the US. We argue that a PTA-setting in which substantial market access has been granted to a much larger developed trading partner may give rise to a rent destruction effect and thus more aggressive external tariff liberalisation.

theoretical generality provides further motivation to link the prediction to the data.

Consider a three country, N-sector competitive economy framework that focuses on the external tariff formation of Home against the Rest of the World (ROW) in the presence and absence of Home forming an FTA with a foreign economy (Foreign). The analytics is greatly simplified by assuming that each country is the natural importer of a distinct subset of goods and that tariffs are the only instruments of protection. A key feature of this model is that Home's external tariffs against imports from ROW are endogenous to Home forming an FTA with Foreign.

To understand the mechanism through which an FTA affects Home's equilibrium external tariff, we first lay out the political economy structure of tariff formation. Following Grossman and Helpman (1994, 1995), Home's political objective is to maximize a weighted average of national welfare and campaign contributions from producers in import competing sectors. Assuming both symmetry across sectors and the overcoming of free-rider problems among producers within a sector, the net payoff V of a representative import-competing sector is the difference between aggregate profit Π and campaign contributions T to the home government. The incentive for the industry paying campaign contributions stems from the protectionist benefit of a higher external tariff t, captured by $d\Pi(t)/dt>0$. Hence:

$$V(t,T)=\Pi(t)-T \tag{1}$$

Home's political objective function G is national welfare W, defined as the sum of producers' surplus, consumers' surplus and tariff revenue, plus the weighted sum of campaign contributions:

$$G(t,T)=W(t)+bT, \tag{2}$$

where b (>0) is a parameter capturing the government's political bias.

Following Maggi and Rodriguez-Clare [1988], the equilibrium external tariff, which we denote as the political tariff t^p, is the outcome of a bargaining process between the government and the domestic import-competing industry. Since this political tariff maximizes the joint payoff function, $W(t)+b\Pi(t)$, it is given by:

$$t^P = \arg\max\,[W(t)+b\Pi(t)] \qquad (3)$$

If there is no political bias in tariff formation, i.e. $b=0$, the government will set an optimal tariff t^* which maximizes social welfare W. The deviation of the political tariff t^p from the optimal t^* stems from the weight put on producers' profits.

Consider now the effect of an FTA with Foreign on Home's political equilibrium. The formation of the FTA will eliminate all trade barriers between Home and Foreign, while allowing Home and Foreign to maintain their external tariffs independently. One of the key channels of the FTA is that it reduces the sensitivity of domestic profits to changes in the external tariff, which is captured by:

$$\Pi_{FTA}\,(t)/dt< d\Pi(t))/dt, \qquad (4)$$

where Π_{FTA} denotes the industry's domestic profits in the presence of an FTA. The intuition behind (4) is that the market access granted to foreign producers leads to an increase in competition and a corresponding decrease in the market share of the domestic industry at any given external tariff. Hence, a tariff-induced increase in the domestic price of the import competing sector has a smaller effect on domestic profits compared to a world without the FTA.

This rent destruction feature of an FTA will diminish the political economy forces in the determination of Home's external equilibrium tariff. Lower domestic industry profits will reduce the industry's protectionist benefit from lobbying which will imply a reduction of campaign contributions. On the other hand, lower campaign contributions will diminish the role of the 'political component' in the government's objective function (2), allowing the determination of a tariff closer to the social optimum.

Denoting the optimal political tariff in the presence of an FTA with t^P_{FTA} and the corresponding welfare maximizing tariff with $t^*_{FTA,}$ we obtain our main theoretical relationship:

$$t^{P}_{FTA} - t^*_{FTA} < t^{P} - t^*. \tag{5}$$

The formation of a market-access based FTA reduces the level of politically determined external tariffs thereby reducing the spread between political and socially optimal tariffs. Assuming that the empirically unobservable optimal tariff is not affected by the FTA, i.e. $t^* \approx t^*_{FTA}$, we obtain the prediction which we will bring to the data:

$$\Delta t^{P}_{FTA} > \Delta t^{P}, \tag{6}$$

where $\Delta t^{P}_{FTA} = t^{P}_{FTA} - t^*_{FTA}$ and $\Delta t^{P} = t^{P} - t^*$. Inequality (6) predicts that one should observe higher multi-lateral tariff cuts in the presence of a market access based FTA than in its absence, ceteris paribus. By construction, the derivation of (6) assumed that the formation of the FTA was exogenous to the determination of external tariffs. In the next section we will argue that the nature of the Canadian-US Free trade (CUSFTA) agreement provides a natural testing environment for (6) in light of Canada's tariff adjustments during the Uruguay Round.

3. CUSFTA and Uruguay Round Tariff Reductions

3.1 Canada-U.S. Free Trade Agreement (CUSFTA)

The Canadian-U.S. Free Trade Agreement was signed in 1988 and entered into force on January 1, 1989.[7] Representing a clearly defined natural policy experiment, CUSFTA obliged policy makers to eliminate tariffs on all products over up to ten years. The agreement led to increasing trade flows between the two contracting parties, in particular in those products experiencing the largest tariff cuts (Clausing, 2001). Total bilateral Canadian and U.S. average tariffs of about eight and four percent in 1988 were reduced in line with product-

[7] The negotiations of the Canadian U.S. Free Trade Agreement (CUSFTA) started in 1986 and were accompanied by a heated controversial public debate about the desirability of CUSFTA. In 1994, CUSFTA was extended to Mexico creating the North American Free Trade Agreement (NAFTA).

specific phasing-in periods,[8] revealing on average larger bilateral tariff cuts on the part of Canada relative to the U.S.[9]

Annex Table 1 provides further information on Canada's tariffs against the U.S., across manufacturing industries, for three different years as well as for the changes from 1989 to 1993 and to 1998. The industries which were subject to the largest average tariff cuts in percentage points after the five year phasing-in period were the furniture, wearing apparel and footwear industries with tariff reductions of up to 10 percentage points.[10] The footwear and wearing apparel industries further reduced tariffs to an overall cut of twenty percentage points each by the end of 1998. Analysing Canada's CUSFTA average tariff concessions in percent rather than percentage points further reveals substantial tariff cuts of above 85 percent in the furniture, paper, printing, industrial chemicals, misc. petroleum and machinery sectors until the end of the first major phasing-in period in 1993. Several other industries such as the beverage, wearing apparel, footwear and tobacco industries, on the other hand, were characterized by smaller average tariff cuts over the same time horizon, pointing to the presence of a larger number of products with longer phasing-in periods. By the end of 1998 all industries were characterized by average tariff reductions of 100 percent.[11]

Investigating trade liberalization in the context of the Canadian-U.S. Free Trade Agreement offers a 'clean' policy experiment, in the sense that CUSFTA was neither part of a larger market reform package nor a response to macroeconomic disturbances allowing for a clear identification of trade reform effects (Rodriguez and Rodrik, 2001; Trefler, 2004). Combined with the fact that the Canadian government committed itself to, sooner or later, eliminate tariffs on all manufacturing goods, suggests that the decision to cut bilateral tariff against the U.S. was largely 'exogenous'.[12]

[8] Article 401 (part 2; chapter 4) of the Canadian United States Free Trade Agreement specifies that all tariffs will be withdrawn according to three different reduction schemes by January 1, 1998. While the first reduction scheme immediately eliminated all tariffs for a series of industries as of January 1989, reduction schemes number two and three focused on a step-wise phasing-in of the reductions over five and ten years, respectively. Moreover, CUSFTA also included duty free trade provisions of automotive products as of January 1, 1998.

[9] The latter has been graphically demonstrated by Trefler (2004).

[10] Since US tariff data for 1988 is unavailable, we were only able to calculate the changes from 1989 onwards.

[11] Coefficients of variation displayed in Annex Table 1 also point to considerable variations of 8-digit HS product-level tariff cuts within individual industries; from 1989 to 1993 as well as from 1989 to 1998.

[12] Clausing (2001:678) further points out that "[..] policy makers committed themselves to eliminate tariffs on all goods [..]" highlighting the vast coverage of CUSFTA. The latter tends to suggest that the decision to participate in CUSFTA was either 'in or out'. Moreover, UN-Trains includes only four product lines for which there was an mfn tariff but no Canadian U.S. preferential tariff for 1989 (the first year for which data is available). Since Canada decided to sign the agreement there seems to have been little choice what to do with the bilateral tariffs but to eliminate them supporting the exogeneity assumption of the latter cuts. In addition, analysing industry level tariff reductions under CUSFTA, Gaston and Trefler (1997), as well as Trefler (2004), conduct a series of statistical endogeneity tests and fail to find evidence for the latter.

3.2 Uruguay Round Tariff Concessions

Reducing and 'binding' multilateral (i.e. mfn) tariffs in order to secure and enhance market access for all GATT-contracting parties were the leading themes of the Uruguay Round (UR) of multilateral trade negotiations. The precise reduction modalities of tariff rates were subject to preliminary negotiations among all the GATT-contracting parties. While Canada strongly favoured a formula-based tariff reduction technique, similar to the 'Swiss formula' applied in the previous (i.e. Tokyo) multilateral trade round, the United States fiercely rejected such a procedural method emphasising that it would only engage in item-by-item trade talks (Stewart, 1999; Laird, 1999).[13] Giving more leeway to a potential sorting by sensitive and less sensitive products, the negotiating parties finally agreed on a request and offer approach without preventing countries to apply reduction formulas on their own (offered) tariff cuts (Stewart, 1999; Laird, 1999).[14] The UR participants, further, agreed to reduce tariffs "with a target amount of overall reductions at least as ambitious as that achieved by the [Swiss-] formula in the Tokyo Round" (Hoda, 2001:35), a statement that was generally interpreted as an overall tariff reduction aim of 33.3% (Hoda, 2001; Laird, 1999).[15]

Table 1 summarises Canada's bound ad-valorem mfn tariff rates agreed upon during the Uruguay Round by industry. While the wearing apparel industry shows the largest average MFN tariff protection before and after the Uruguay Round, (with rates of 24 and 17 percent, respectively), the footwear and textile industries are close followers with 22 and 18 percent average protection before, and 17 and 12 percent (respectively) after the Uruguay Round. The lowest average mfn tariff rates before the UR appear in beverages (5 percent) and the wood, paper and non-ferrous metals industries (all 8 percent), while the paper, printing (both duty free) and iron and steel (one percent) sectors were least protected after the UR. In terms of percentage point reductions, printing as well as the iron and steel industries (10 and 8 percentage points, respectively) had the largest average MFN cuts. Beverage and wood products benefited from relatively low average trade barrier reductions (i.e. 2 and 3 percentage points, respectively). Finally, it is also worth noting that the inter-industry variations are

[13] Apart from Canada, the EU and Japan also favoured a formula based reduction approach (Stewart, 1999; Laird, 1999).

[14] Annex Figure 1 provides a graphical analysis of the relationship between initial and final bound rates and tends to confirm that Canada's initially higher (bound) mfn tariff rates were reduced more in the UR, as indicated by the increasing gap between the 45 degree reference line and the two linear regression lines. There is, however, also some indication of the so-called sectoral agreements. In these sectors many tariff lines were reduced to a common (including zero) rate pointing to an alternative reduction approach.

[15] The U.S. implemented the item-by-item request and offer approach by submitting extensive lists of tariff reduction requests to their main trading partners in October 1989 (Laird, 1999).

complemented by considerable intra-industry variations, as illustrated by the coefficients of variation (see Table 1, Column (4)).

Table 1: Canadian industry-level (bound) tariff MFN reductions agreed upon during the Uruguay Round

		(1)	(2)		(3)		(4)			(5)
		Number of HS 8-digit tariff lines per industry	Before Uruguay Round		After Uruguay Round		Change Uruguay Round (Percentage Points)			Change Uruguay Round (Percent)
ISIC code	Sector name		Mean	Std. dev.	Mean	Std. dev.	Mean	Std. dev.	Coef. variation	Mean reductions in % of pre-UR rates
311	Food products	243	0.12	0.04	0.07	0.04	-0.05	0.03	0.58	-38.1
313	Beverages	1	0.05	-	0.03	-	-0.02	-	-	-36.0
314	Tobacco	2	0.15	0.07	0.10	0.04	-0.05	0.02	0.45	-36.1
321	Textiles	503	0.19	0.06	0.12	0.03	-0.07	0.04	0.51	-37.6
322	Wearing apparel except footwear	238	0.24	0.03	0.17	0.02	-0.07	0.02	0.23	-27.8
323	Leather products	25	0.10	0.03	0.07	0.02	-0.03	0.01	0.29	-34.4
324	Footwear except rubber or plastics	14	0.22	0.01	0.17	0.05	-0.06	0.04	0.74	-25.4
331	Wood products except furniture	10	0.08	0.03	0.06	0.02	-0.03	0.01	0.39	-33.9
332	Furniture except metal	1	0.15	-	0.10	-	-0.05	-	-	-35.3
341	Paper and products	122	0.08	0.02	0.00	0.01	-0.07	0.02	0.32	-95.5
342	Printing and publishing	1	0.10	-	0.00	-	-0.10	-	-	-100
351	Industrial chemicals	542	0.11	0.02	0.06	0.01	-0.05	0.02	0.39	-47.6
352	Other chemicals	130	0.11	0.04	0.04	0.03	-0.07	0.04	0.51	-61.1
353	Petroleum refineries	21	0.11	0.02	0.07	0.01	-0.04	0.01	0.33	-36.2
354	Miscellaneous petroleum and coal products	4	0.13	0.00	0.07	0.00	-0.06	0.00	0.00	-48.0
355	Rubber products	34	0.13	0.05	0.08	0.04	-0.05	0.03	0.57	-40.6
356	Plastic products	42	0.17	0.05	0.08	0.04	-0.08	0.05	0.59	-50.7
361	Pottery china earthenware	1	0.11	-	0.08	-	-0.04	-	-	-34.2
362	Glass and products	19	0.09	0.05	0.02	0.04	-0.08	0.03	0.45	-82.0
369	Other non-metallic mineral products	13	0.09	0.02	0.06	0.01	-0.03	0.01	0.23	-33.2
371	Iron and steel	246	0.09	0.02	0.01	0.02	-0.08	0.03	0.34	-92.3
372	Non-ferrous metals	217	0.08	0.03	0.04	0.02	-0.04	0.03	0.64	-53.7
381	Fabricated metal products	89	0.10	0.02	0.05	0.03	-0.05	0.03	0.51	-52.6
382	Machinery except electrical	299	0.09	0.02	0.05	0.03	-0.04	0.03	0.57	-48.2
383	Machinery electric	142	0.10	0.03	0.06	0.02	-0.04	0.02	0.39	-39.2
384	Transport equipment	74	0.11	0.03	0.07	0.03	-0.04	0.02	0.52	-39.2
385	Professional and scientific equipment	60	0.12	0.06	0.06	0.04	-0.05	0.02	0.43	-46.0
390	Other manufactured products	45	0.12	0.04	0.07	0.03	-0.04	0.02	0.43	-37.9
	Total	3138	0.12	0.03	0.07	0.03	-0.05	0.02	0.44	-48.0

Source: Authors' own calculations.

4. Empirical Implementation

4.1 Econometric specification

The theoretical prediction (6) implies larger external tariff cuts in the presence than in the absence of a market-access based preferential trade agreement. In the previous section we have established that CUSFTA provides a fitting policy experiment to examine the impact of CUSFTA on Canada's external tariff cuts during the Uruguay Round. Since it is not possible to observe Canada's external tariff cuts in the absence of CUSFTA, we follow Limão's (2006) identification strategy and use tariff changes on non-FTA goods as the counterfactual for tariff changes in the absence of FTA goods. Contrasting political tariff adjustments in the presence and in the absence of an FTA, equation (6) implies larger external tariff cuts on products imported with preferential market access relative to products not imported under such preferences, ceteris paribus.

Our objective is to estimate the impact of CUSFTA preferences on Canadian multilateral tariff cuts at the product level. The econometric specification is given by

$$\Delta t_i = \alpha + \beta_1 I_i + \beta_2 R_i + \beta_3 P_i + \beta_4 \Delta X_I + \beta_5 t_{i,-1} + \upsilon_i \,. \qquad (7)$$

The dependent variable Δt_i represents the change of the bound MFN tariff negotiated during the Uruguay Round. The analysis is conducted at the 8-digit HS product level and encompasses a sample of 3138 observations. Our sample excludes agricultural products because of the heavy incidence of non-tariff measures in that sector. Product lines with initial zero MFN tariffs are also excluded due to the impossibility to grant tariff preferences on these items.

Our main explanatory variable of interest is the indicator variable I_i which takes the value 1 if product i was granted a specific preferential tariff concession and has also been imported from the US, otherwise it is zero. Canadian tariff preferences, in place at the time of the UR, had been granted under several preferential trading schemes including the General Preferential Tariff (GPT), the Caribbean-Canada Trade Agreements (CARIBCAN) as well as the Canadian-U.S. and later the North American Free Trade Agreement (CUSFTA/NAFTA), the latter two representing the focus of our analysis. We further introduce additional measures for preferential market access using the share of imports originating from North-American trading partners as well as the latter's interaction with the CUSFTA FTA-good indicator (I_i). The impact of the latter two measures on negotiated multilateral tariff reductions may provide additional information regarding the magnitude of tariff cuts relative to the amount of preferential imports. The main theoretical prediction is that the coefficient of I_i is positive and the size of the coefficient captures the magnitude of the building bloc effect of CUSFTA.

The remaining variables in (7) are controls that capture aspects of tariff adjustments which are not captured by the theoretical mechanism and which have been suggested by previous studies in the literature. The variable R_i captures the extent to which Canada lowered its external tariffs in good i to reciprocate tariff reductions of its trading partners. Representing an important element in WTO negotiations, we account for the latter by defining R_i as $R_i = \sum_k s_{it}^k [\sum_i w_i^k \Delta t_i^k / t_i^k]$ where the sum of import weighted percentage tariff concessions (i.e. $\sum_i w_i^k \Delta t_i^k / t_i^k$) of WTO-member country k is aggregated over all products i and further multiplied by either Canada's 1992-import share from country k if the latter is one of Canada's top-5 import suppliers in good i (s_{it}^k) or otherwise by zero.[16] By multiplying country k's average tariff concessions by the import share of Canada's most important trading partners we take into account the agreed tariff concessions of Canada's UR negotiating partners as well as the fact that Canada most likely only negotiated with its most important suppliers. We thereby assume that Canada only engaged in direct trade talks with its top-5 import suppliers in each product line.[17] Aggregation over all principle suppliers k finally delivers a product level proxy measurement for reciprocity-based (bound) MFN tariff reductions.[18]

Potentially lower mfn tariff cuts due to a large number of 'free-riding' countries are accounted for by the variable P_i. Tariff reductions based on reciprocity combined with the GATT's non-discrimination clause may give rise to a so-called MFN externality effect; smaller countries may benefit from the larger traders' (reciprocal) tariff cuts without offering any trade barrier reductions in return.[19] A larger number of 'free-riding' countries may

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[16] Data on the aggregated sum of import weighted percentage tariff concessions on product i (i.e. $\sum_i w_i^k \Delta t_i^k / t_i^k$) of country k stems from Finger et al. (2002), where w_i^k denotes product i's share in total imports of country k and $\Delta t_i^k / t_i^k$ represents k's tariff cuts in product i. Moreover, it is worthwhile noting that reciprocal tariff reductions do not necessarily refer to the same matching set of products. In practise it is more common to reciprocate with tariff reductions on other products which are possibly more important for the partner country. Some authors therefore distinguish between products j and products i, where j denotes products subject to tariff reductions in partner country k, and i to products subject to mfn tariff cuts by the Home country. For simplicity, however, we use the product index i for both trading partners.

[17] Note that information on Canada's direct negotiating partners during the UR is not publicly available.

[18] GATT regulations denote a country as the 'principal supplier' when the latter accounts for the largest share of GATT imports in a specific product of another country. Country k's export share of product i to Canada (s_{it}^k) indentifies Canada's 'principal suppliers' for each good i.

[19] The MFN externality effect becomes more obvious when we consider a scenario in which there is only one exporter of a certain product to Canada, which would then be the only beneficiary of Canada's tariff reductions on a certain product.

therefore translate into smaller Canadian tariff cuts since the latter's government cannot expect extensive reciprocal reductions in return and, therefore, may have a lower incentive to liberalize itself. Given the lack of information on Canada's direct negotiating partners, we capture this effect by constructing a measure based on the share of Canada's non top-5 (i.e. smaller) trading partners per product line i. A significant variation in the latter ratio between 1994 and 1988 may reflect a change in the number of non top-5 exporting countries between the two last successfully concluded trade rounds and thus serve as a proxy measure of an mfn externality effect. [20] We define P_i as an indicator variable taking the value one if the latter mentioned change is larger than the median change and zero otherwise.

The variable ΔX_I introduces political economy forces into the model. As shown in previous studies political economy considerations may be important in shaping a country's trade policy and are likely to result in less ambitious tariff reductions in politically influential sectors. We aim to account for the latter by defining ΔX_I as $\Delta X_I = \Delta(X_I/M_I)/\varepsilon_I$ where $\Delta(X_I/M_I)$ denotes the change in the inverse import penetration ratio between 1992 (final phase UR) and 1978 (end TR) and ε_I reflects the import demand elasticity in the respective ISIC 3-digit industry. Finally, given that larger mfn tariff reductions may be easier to implement on products where pre-UR tariffs were already relatively high and in light of significant differences in average pre-UR tariffs in our sample, we also introduce initial (i.e. pre-UR) tariff rates ($t_{i,t-1}$) as an additional regressor in our model.

[20] Canadian product level import data from 1988 are the earliest ones available from UN-TRAINS. We assume that if the change of small exporters to Canada per product line i was large enough between 1994 and 1988, and therefore mirrors a longer term change between 1994 (end-Uruguay) and 1978 (end-Tokyo), the constructed proxy variable is a valid instrument for the MFN externality effect.

4.2 Data

All variables introduced in the model are listed and defined in Annex Table 2, with the summary statistics for each of the variables set out in Annex Table 3. In this section we provide a short data summary focusing on the most important characteristics. We use 8-digit Harmonized Standard (HS) information on bound mfn advalorem tariffs from the WTO's schedule of concessions and preferential tariff data from the UN-Trains database. The latter database also provides 8-digit HS Canadian import data which we employ to construct the preference indicator variables. Information on Non-Tariff Barriers (NTBs), used to instrument the preference indicators, is not publicly available, but was very helpfully provided by the Trade Information Department of UNCTAD.[21]

Reciprocal tariff reductions have been constructed by using import weighted UR tariff cuts from Finger et al. (2002), who construct an aggregate measure of the participating countries' bound tariff concessions. Using the latter information we combine country level aggregated average concessions with supplier-specific 8-digit HS Canadian import shares using data from the UN-Trains database.

In order to proxy political economy forces we calculate elasticity-weighted inverse import penetration ratios at an ISIC 3-digit industry level. Sector-level import and production data, used to construct the latter ratios, are retrieved from the UN-COMTRADE and UNIDO databases respectively, while the industry-level import demand elasticities are from Kee et al. (2009). Industry-level (i.e. ISIC 3-digit) data on value added and on the number of establishments, employed to construct instruments for the political economy variable, are also from the UNIDO database. In order to take into account different aggregation levels of our independent variables we use clustering of standard errors at the sector level.

[21] We are very grateful to Hiroaki Kuwahara (Chief of the Trade Information Section - Trade Analysis Branch, DITC/UNCTAD), who kindly provided us with the NTB data for the year 1993, which he has extracted from old diskettes and CD versions of TRAINS.

4.3 Estimation and Empirical Findings

Establishing a causal relationship between tariff preferences and multilateral tariff changes is often considered a major challenge. Given that tariff preferences are more valuable the higher the external tariff protection, the level of the MFN tariffs themselves may determine whether a product receives a preference. Smaller expected MFN tariff reduction for certain products may, therefore, strongly enhance a partner country's desire for preferential market access. In light of the overriding influence of CUSFTA preferences for Canada and the presumably greater importance of preferential market access for Canada to the US market than vice-versa, we, however, expect possible endogeneity concerns to be less of an issue. Our preferred empirical specification is therefore based on a non-instrumental OLS regression approach.

Additional IV-GMM estimation techniques are used in order to contrast and test the latter results. We use an instrumental variable technique to tackle possible endogeneity concerns linked to reverse causality. The variables used to instrument the main preference indicator (I_i) include the change in world-prices between 1992 and 1994 as well as two indicator variables reporting whether a product (i) was imported in 1994 (D_i^{94}), and (ii) was subject to an NTB in 1993 (D_i^{ntb93}).[22] While world price changes, between 1992 and 1994, influence a partner country's monetary benefit from a granted preference by increasing the value of the preferentially exported items and hence its attractiveness for preferential trading partners, they are likely to be uncorrelated with the error term as the Uruguay tariff reductions took effect from 1995 onwards.[23] Trading partners are also more likely to ask for preferences on goods which are imported also in order to gain an advantage over other competing exporters. The import dummy D_i^{94} is therefore introduced as an instrument since the latter is unlikely to be correlated with the error term, again due to the timing of the agreed MFN tariff rate changes.[24] Countries are also more likely to ask for preferences on goods which they suspect to be subject to NTBs in the future – as a proxy for future NTBs

[22] World price changes at the 8-digit HS product-level are proxied by calculating unit-values using import value and quantity information available at UN-TRAINS.

[23] Given a fixed amount of exports X_i^S, a country's benefit from a preference for product i can be written as $(t_i^{mfn} – t_i^{pref}) * p_i^w X_i^S$. The latter expression indicates that the higher world prices, the higher the benefit arising from a preference. Increasing world prices may therefore help PTA partner countries to overcome fixed export costs, which could make them more likely to export. Furthermore, the inclusion of unit-values as an instrument reduces our sample from 4742 to 3138 observations. Estimating OLS and IV-GMM with the larger sample (i.e. without unit-values) results in qualitatively identical findings. The latter results are available upon request.

[24] Using a partner-specific import dummy $D_i^{j\,94}$, which is directly linked to the PTA good indicator (i.e. $I_i^j = PR_i^j * D_i^{j\,94}$) instead of an overall import indicator (D_i^{94}) results in qualitatively similar results. The latter results are also available upon request.

data of 1993 is used.[25]

Potential reverse causality problems may also arise from the reciprocity variable. By taking advantage of the timing and mode of the Uruguay Round, we employ an instrumental variable capturing the unilateral external tariff reductions independently undertaken between 1986 and 1992. In light of serious doubts regarding the successful conclusion of the UR before 1992 (Stewart, 1999; Finger et al., 2002) and given that previously undertaken unilateral tariff reductions were later, between 1992 and 1994 - when the final tariff rates were negotiated, explicitly reciprocated, unilateral tariff changes serve as a valid instrument.[26] Finally, given that the political economy variable (defined as the change in the elasticity weighed inverse import penetration ratio between 1978 and 1992) is in all its elements strongly dependent on domestic prices and hence MFN tariff rates, reverse causality issues may also emerge from this variable too. Given that political economy forces may display some persistency over time using lagged values may not fully take into account possible endogeneity issues. The introduction of the change in scale economies (valued added/number of firms) at an ISIC 3-digit level between 1981 and 1992 and the latter's interaction with world prices is based on the intuition that industries with higher fixed entry costs are likely to have a higher inverse import penetration ratio (i.e. X_I/M_I) and that world prices directly impact on domestic prices which in return determine X_I, M_I and ε_I.

Table 2 reports the estimations of equation (7) using heteroscedasticity-robust OLS and two-step efficient generalized methods of moments (IV-GMM) estimators. The standard errors are clustered at the industry level. The results show that Canadian preferences granted under CUSFTA have acted as a 'building block' for more ambitious multilateral tariff reductions agreed upon during the Uruguay Round, with coefficients on the CUSFTA-preference indicator variable of -0.022 and significant at the 1% level; MFN tariff reductions of preferentially imported goods being larger than those of their counterparts, having controlled for other influences. The result is line with the argument that increased internal competition and political rent destruction may have led to a weakening of protectionist forces, resulting in lower multilateral tariffs on products covered by preferences. [27]

[25] We also interact the NTB indicator variable (D^{ntb93}) with the import dummy variable (D_i^{94}) and introduce the combined component as an additional instrumental variable. The intuition is that a country would even be more inclined to ask for a preference if Canada already imported the latter product, which could be subjected to a future NTB.

[26] Finger et al. (2002:121) note that "according to delegations, the informal practice was more or less to count from applied rates in 1986 to the bound rate agreed at the Uruguay Round. By this practice, countries that 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." Moreover, Limão (2006) points out that the data shows that all counties engaged in some unilateral tariff reductions.

[27] In order to take into account potential misclassifications at such detailed levels of product disaggregation we consider a product to be exported to Canada if the latter's trade value is above a certain (low) threshold. In our estimations we

Analysing the results of different econometric specifications shows that OLS as well as IV estimates both report coefficients of -0.022 (Regressions (1) and (5), respectively), for Canadian preferences granted under CUSFTA indicating that bound MFN tariff reductions of preferentially imported goods were on average between 2.2 percentage points larger than on goods imported under the mfn tariff.[28] Moreover, average tariff reductions of around 3.6 percentage points for non-FTA goods tend to highlight the economic importance of the determined 'building block' effect.[29]

Using the share of Canadian U.S. imports at the end of the Uruguay Round (i.e. six years after the implementation of CUSFTA) as an additional measure for preferential market access tends to confirm the latter findings (Regressions (3) and (7)). Coefficients of -0.011 and -0.014 for the OLS and IV estimations, respectively, point to larger tariff cuts on products which were predominantly imported by the US providing suggestive evidence for more aggressive tariff reductions in industries where rents had to be shared (i.e. 'leakage' effect). [30]

The above findings are confirmed when NAFTA preferences are taken into account. With coefficients of -0.022 for the NAFTA preference good specifications (Regressions (2) and (6)), and again of 0.011 and 0.014 for the U.S. import share variables (Regressions (3),(4) and (7),(8)), the results for NAFTA are identical to the findings on CUSFTA, suggesting that CUSFTA preferences are an important driver of Canada's overall preference-tariff relationship.

Political economy factors, included in the regressions as the change in industry-level elasticity-weighed import penetration ratios between 1978 and 1992, are found to be of a minor significance for Canadian (bound) MFN tariff reductions during the UR. Tariff reductions based on reciprocity are shown to have a positive impact in the overall specifications when estimated with OLS (Regressions 1 to 4). Coefficients of around 0.029, significant at the 5% level, may indicate that Canada reduced its tariffs more ambitiously on products imported from trading partners which offered larger tariff concessions by themselves. The IV results, however, do not report a significant influence of the reciprocity variable. The 'MFN externality effect' (β_3) has in all specifications the expected sign and is

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apply a threshold of 1,000 USD (the lowest value recorded in UN-Trains). Using no threshold or thresholds of 2, 3 or 5 thousand USD results in qualitatively identical findings. The results are available upon request.

[28] The average tariff reduction across all product lines amounts to 6 percentage points (see Annex Table 2).

[29] It is interesting to note that Karacaovali and Limão (2008) find a 'stumbling block' of 1.5 percentage points in their main IV specification. Against the backdrop of an almost identical average mfn tariff reduction for non-PTA goods in the latter authors' sample (i.e. 3.4 percentage points), the economic magnitude of the, for Canada determined, 'building block' effect seems to be much larger.

[30] Interacting the U.S. import share with the CUSFTA preference indicator variable (used in regressions (1) and (5)) results in qualitatively similar results. The latter estimations are available upon request.

also highly significant (i.e. at the 1% level). Estimates between 0.008 and 0.011 point to larger Canadian tariff reductions when the number of non top-5 suppliers, and hence potentially free-riding countries, declined by a significant amount between 1988 and 1994.[31] The significant free-riding results may also capture some reciprocity aspects, since a smaller number of potentially free-riding countries provides an enhanced incentive for more meaningful reciprocal tariff reductions. The highly significant MFN externality coefficient in the IV-regressions for the CUSFTA- and NAFTA FTA good specifications may, therefore, partially explain the insignificance of the reciprocity variable in the IV-GMM estimations. Finally, initial (i.e. pre-UR) tariff rates seem to be an important determinant in all specifications. Negative coefficients of between -0.302 and -0.242 tend to point to larger tariff reductions of initially high mfn rates. Given that, on average, pre-UR mfn tariffs display considerable differences when comparing FTA and non-FTA goods, this seems to be an important control variable. For the CUSFTA and NAFTA-good specifications, pre-UR tariffs tend to be markedly higher than their non-FTA good counterparts, leading to the possibility that the 'building block' effect may be driven by larger reductions for higher initial MFN rates when not controlling for the latter.[32] The inclusion of the pre-UR rates may, as a result, also account for the potential use of an implicit reduction formula in the UR negotiations by Canadian negotiators.

The test statistics in Table 2 show that the estimates are reasonably robust. Using 'Hansen's J'-tests to test the excluded instruments' joint significance, we find that the second-stage error terms are uncorrelated with the latter instruments. This is also illustrated by the strong acceptance of the null hypothesis stating an overall good instrument quality.[33] In light of a relative large number of instruments 'Hansen J'-tests may lose some of their explanatory power. We follow Karacaovali and Limão (2008) in testing the subgroup of, *a priori* more endogeneity-prone, instruments (i.e. the NTB and import dummy variables) for orthogonality to the error term. The displayed difference-in-Sargan statistics (C-stats.) reject in all model specifications the null of correlation with the error term for the smaller subset of instruments.[34] In order to test the OLS estimates for inconsistency, Table 2 also reports the results for the standard Hausman tests. The displayed test probabilities for the IV-GMM regressions tend to show no concern for an inconsistency of the OLS estimates at a 5 percent

[31] The 'mfn externality' effect is also significant in the specifications including the Canadian U.S. import share, however, only when estimated with OLS.

[32] The pre-UR average bound tariff levels for the CUSFTA- and NAFTA- good specifications both amount to 12.79 percentage points, whereas the respective non-CUSFTA and non-NAFTA good counterparts are 11.91 and 11.87 percentage points, respectively.

[33] The latter is underscored by a strong rejection of the F-tests in the first stage regressions.

[34] The subset of tested instruments includes all instruments constructed by using the variables D_i^{ntb93} or D_i^{94} (see Annex Table 1).

significance level. Test statistics reported in Table 2, therefore, only provide very weak evidence for the existence of an endogeneity-bias of the OLS estimates.

Table 2: The Impact of Canadian Trade Preferences on Multilateral Tariff Reductions in the Uruguay Round

	OLS				IV-GMM			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	CUSFTA	NAFTA	CUSFTA (US M.-Ratio)	NAFTA (US M.-Ratio)	CUSFTA	NAFTA	CUSFTA (US M.-Ratio)	NAFTA (US M.-Ratio)
$I_i^{‡}$	-0.022***	-0.022***			-0.022***	-0.022***		
	(0.006)	(0.006)			(0.004)	(0.004)		
Im_i			-0.011**	-0.011**			-0.014**	-0.014**
			(0.005)	(0.005)			(0.006)	(0.006)
$R_i^{‡}$	0.029**	0.028**	0.030**	0.028**	-0.004	-0.005	-0.002	-0.005
	(0.013)	(0.013)	(0.012)	(0.012)	(0.015)	(0.015)	(0.016)	(0.015)
$\Delta X_I^{‡}$	0.000	0.000	0.000	0.000	-0.000	-0.000	0.000	0.000
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
P_i	0.011***	0.010***	0.008***	0.008***	0.010***	0.010***	0.006	0.006
	(0.002)	(0.002)	(0.002)	(0.002)	(0.003)	(0.003)	(0.005)	(0.005)
$t_{i,t-1}$	-0.284***	-0.283***	-0.302***	-0.301***	-0.242*	-0.245**	-0.254*	-0.253*
	(0.058)	(0.058)	(0.054)	(0.055)	(0.124)	(0.121)	(0.151)	(0.151)
Constant	0.011	0.011	0.001	0.000	-0.008	-0.008	-0.017	-0.019
	(0.010)	(0.010)	(0.011)	(0.010)	(0.016)	(0.016)	(0.021)	(0.021)
Observations	3138	3138	3138	3138	3138	3138	3138	3138
Number of FTA goods	3011	3021	-	-	3011	3021	-	-
Hansen's J (p-val.)[a]	-	-	-	-	0.673	0.675	0.622	0.618
C-stat (p-val.)[b]	-	-	-	-	0.547	0.550	0.588	0.593
Endogeneity (p-val.)[c]	-	-	-	-	0.055	0.047	0.069	0.069
Heterosked. (p-val.)[d]	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000

Notes. Columns (1) and (2) illustrate the OLS regression results when using a CUSFTA preference indicator variable. Columns (3) and (4) present the regression results focusing on the product-level US import share. Regressions (5)-(8) represent the IV-GMM results for the latter specifications. Regressions (7) and (8) use partner-country specific import dummies D_i^j94 instead of D_i94 as an instrument. All regressions are based on heteroskedasticity robust standard errors and clustering at the 3-digit ISIC industry level. *, **, *** illustrate the 10%, 5%, 1% significance levels, respectively. For the IV-GMM estimations, the instruments exclusion F-tests of the first-stage regression are all rejected either at the 1 or 5 percent threshold level with F-statistics that are considerably larger than 10 for I_i and R_i. For ΔX_I the latter statistic assumes values at around 4. The first-stage regression results are available upon request. (a) Sargan-Hansen test of over-identifying restrictions. Under the null hypothesis all instruments are jointly uncorrelated with the error term of the second stage regression and correctly excluded from the estimated equation (i.e. the instruments are valid instruments). (b) C-statistic (or Difference-in-Sargan statistic) allows for testing the exogeneity of one or a subset of instruments. The null hypothesis states that the tested instruments are exogenous. The subset of tested instruments is: D_i94, Dntball, Dntball*D_i94, Dntb, (Δp9294)avg*Δscale. The C-statistic is defined as the difference of the Sargan-Hansen value of the equation with the restricted (i.e. omitted questionable instruments and, therefore, also smaller) set of instruments and the equation with the unrestricted (i.e. full and larger) set of instruments (i.e. C-stat = Jr - Ju). (c) Endogeneity test for the potentially endogenous regressors marked with ‡. The null hypothesis says that the marked endogenous regressors can actually be treated as exogenous (i.e. OLS estimation is consistent and efficient). The test statistic follows a chi-squared distribution with degrees of freedom equal to the number of regressors tested and is defined as the difference of two Sargan-Hansen values (cf. above). (d) Pagan and Hall's test of heteroskedasticity for estimations using instrumental variables (IV). The null hypothesis is that no heteroskedasticity is present.

4.4 Robustness analysis

We conduct a series of robustness exercises. A summary of the findings from these are in Table 3, the first column reporting the estimated coefficient on the preference good indicator (I_i) from Table 2 (columns 1 and 5) with the estimated coefficients on the other variables suppressed. The other columns, in Table 3, record the estimated coefficient on I_i for the alternative specifications (with the other variables in the regression suppressed, but available on request). The results for the OLS estimations are illustrated in the first row of Table 3, whereas the second row presents the outcome for the IV-GMM specifications.

We first introduce an additional indicator variable at the HS 1-digit level to test whether our 'building block' effect still holds when accounting for unobserved industry effects defined according to the Harmonized System (HS). Column (2) shows that the inclusion of the additional indicator results in the same findings as in the baseline specification when estimating with OLS and a still highly significant (albeit slightly smaller in absolute terms) coefficient of -0.020 when estimating with IV-GMM.

Furthermore, we show that the finding of a 'building block' effect holds if we exclude the product lines affected by the zero-for-zero sectoral negotiations (column 3) and these product lines plus chemicals (column 4). Accounting for sectoral agreements tests whether the 'building block' results may be driven by an alternative tariff reduction rationale.[35]

We also exclude the reciprocity variable and its instrument from our regressions. Column (5) shows that the exclusion of the reciprocity variable results in almost the same findings as in the baseline specification. Karacaovali and Limão (2008) argue that products affected by NTBs towards all trading partners may be characterized by additional common unobserved features which in return may also influence MFN tariff reductions. Although, testing the subset of instruments involving the NTB variable for orthogonality to the error term rejects this possibility, we explicitly drop all instruments including the latter indicator variable. The results are reported in column (6). The findings show a persistent 'building block' effect of -0.021 for the IV regressions, compared to the baseline specification, in column (1), a slightly smaller but still highly significant, 'building block' effect.

[35] Given the varying distribution of PTA and non-PTA goods across industries, we also test whether our main findings are driven by industry-specific characteristics by dropping successively individual industries. The results, not reported in Table 3 but available upon request, show that the findings still hold when excluding all industries individually.

Finally, we check that what we are treating as a CUSFTA tariff preference effect is only due to CUSFTA. Some of the FTA goods under CUSFTA (3011 goods) are also goods that are preferentially treated under Canada's various other preferential trading schemes. If we exclude those goods from the set of FTA goods to identify goods that are CUSFTA only preferential goods, we are left with 1713 goods. In column (7) we report on the estimated coefficient on the preference good indicator (I_i) for this narrower set of CUSFTA-only preferences. Again we find that the estimated coefficient is negative and significant (at a 10% level for OLS and 1% for the IV estimation).

For all of the robustness exercises above 'Hansen's-J' statistics strongly reject the null of a correlation between the instruments and the second-stage error terms.[36] Testing the subgroup of, *a priori* more endogeneity-prone, instruments for a potential correlation to the second-stage error term, also strongly rejects the null of non-orthogonality to the error term, as illustrated by the Difference-in-Sargan statistic p-values which are larger than 0.17 for all regressions.[37]

Table 3: Robustness Analysis

	OLS & IV-GMM						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Robustness test	CUSFTA Preferences	"HS Industry Effects"	"Zero-for-Zero" Sectoral Agreements	"Zero-for-Zero" Agreements incl. Chemicals	Excluding Reciprocity	Exclude all NTB instruments	CUSFTA only Preferences
I_i^{OLS}	-0.022***	-0.022***	-0.016***	-0.014***	-0.021***	-	-0.008*
	(0.006)	(0.004)	(0.004)	(0.004)	(0.004)	-	(0.004)
I_i^{IV}	-0.022***	-0.020***	-0.015***	-0.012***	-0.022***	-0.021**	-0.031***
	(0.004)	(0.003)	(0.004)	(0.004)	(0.004)	(0.004)	(0.009)
Observations	3138	3138	2571	1911	3138	3138	3138
Number of FTA-goods	3011	3011	2447	1791	3011	3011	1713

Notes: Column (1) displays the baseline regression results reported in Table 2 (Column (1)). In all regression concordance tables have been used. Column (2)-(6) illustrate the regression results derived when subjecting the baseline findings, displayed in Column (1), to various robustness tests. Tariff lines covered by so-called sectoral agreements including the 'zero-for-zero' concessions have been excluded in Columns (3) and (4). Information on the product coverage of sectoral agreements is limited to the information provided by the WTO secretariat (WTO, 2005). The baseline results have also been tested to the exclusion of individual industries. The latter results confirm the 'building block' findings and are available upon request. All regressions use heteroskedasticity robust standard errors clustered at the 3-digit ISIC industry level. *, **, *** illustrate the 10%, 5%, 1% significance levels, respectively.

[36] Robustness tests (2), (5) and (6) all report Hansen's-J probability values above 0.62, while the tests reported for Columns (3) and (4) report probabilities of 0.26 and above.

[37] All statistical endogeneity test probabilities show values above 0.18, apart from robustness test (2) which shows a p-value of 0.08.

5. Conclusions

In this paper we examine the impact of tariff preferences on UR multilateral tariff cuts by Canada. While other studies have investigated the effect of preferences given by industrial countries to developing countries, we extend the literature by analysing the effect of trade preferences on the mfn tariff reductions of an industrial country with preferences that include comprehensive preferential trade preferences for a much larger, industrial trade partner. In the present setting we anticipate the exchange of market access to be an important consideration in the setting up of the preferences, and to be much greater scope for a rent destruction effect than where offering preferences to smaller, developing countries. With greater rent destruction from the FTA, we expect there to be reduced political economy resistance to multilateral liberalisation.

Contrary to earlier studies which find evidence for a 'stumbling block' caused by US and EU granted preferences (cf. Limão, 2006; Karacaovali and Limão, 2008), we find a 'building block' effect of Canadian preferences on multilateral tariff reductions agreed upon during the Uruguay Round. Our results show that preferentially imported products were subjected to (bound) MFN tariff reductions which were on average 2.2 percentage points larger than those for non-preferentially imported products.

The identified 'building-block' effect of Canadian preferences is in line with some other findings that preferences act as a catalyst for multilateral tariff cuts (cf. Bohara, Gawande and Sanguinetti, 2004; Estevadeoral, Freund and Ornelas, 2010; Calvo-Pardo et al. 2010). These studies use an alternative empirical methodology to the present study. But our findings stand in stark contrast to Limão (2006) and Karacaovali and Limão (2008), who, using the same methodology as that used here, provide empirical support for the existence of a 'stumbling block' for the US and EU. The contrast in findings cannot therefore be accounted for by differences in the empirical method. It must rather be because of differences in context and influences on the decision-making process. The findings for the US and the EU have been rationalised in the context of a theoretical framework in which smaller trading partners reciprocate tariff preferences with cooperation agreements in non-trade issues. We argue in the present context that PTAs based on mutual market access concessions, may lead to increased intra-bloc competition and a destruction of political rents which in turn facilitate mfn tariff cuts following the formation of a PTA (Bagwell and Staiger, 1999; Ornelas, 2005a). In contrast to the US and the EU, both of which have a substantial number of PTAs with smaller countries, Canada's most important PTA involves predominately preferential market access granted to a much larger economy - the US. The over-riding influence of market access-based preferences in this case can be expected to have reduced internal Canadian protectionist forces and induced our ‘building block’ finding for Canada in the Uruguay Round.

Our finding is consistent with the empirical evidence found for mfn tariff changes in the context of South-South preferential trade agreements, which would also appear to be based on mutual market access concessions (Freund and Ornelas, 2010; Calvo-Pardo et al., 2010). The nature of the PTA formed – i.e. whether it is primarily based on market access concessions or on cooperation in non-trade issues -, as well as the relative competitiveness (i.e. size) of the preferential trading partner(s) are decisive elements in determining the impact of preferences on multilateral tariff cuts.

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Annex

Annex Table 1: Canadian Bilateral Tariff Rates against the U.S. across Manufacturing Industries for different Years

		(1)	(2)	(3)	(4)			(5)		(6)			(7)	
		1989	1993	1998	Bilateral Tariff cuts: 1989-1993			Bilateral Tariff cuts (in%): 1989-1993		Bilateral Tariff cuts: 1989-1998			Bilateral Tariff cuts (in%): 1989-1998	
ISIC code	Sector name	Mean	Mean	Mean	Mean	Std. dev.	Coef. Var.	Mean	Std. dev.	Mean	Std. dev.	Coef. Var.	Mean	Std. dev.
311	Food products	0.07	0.03	0.00	0.04	0.04	0.89	-55.7	22.4	0.07	0.06	0.93	-100.0	0.0
313	Beverages	0.06	0.04	0.00	0.02	0.03	1.95	-44.4	0.3	0.06	0.08	1.45	-100.0	0.0
314	Tobacco	0.14	0.08	0.00	0.06	0.03	0.47	-44.2	0.3	0.14	0.06	0.46	-100.0	0.0
321	Textiles	0.16	0.09	0.00	0.08	0.04	0.48	-47.1	12.0	0.17	0.07	0.43	-100.0	0.0
322	Wearing apparel except footwear	0.20	0.11	0.00	0.09	0.02	0.24	-44.5	0.1	0.20	0.05	0.25	-100.0	0.0
323	Leather products	0.05	0.02	0.00	0.04	0.05	1.19	-72.2	28.2	0.06	0.06	1.09	-100.0	0.0
324	Footwear except rubber or plastics	0.20	0.11	0.00	0.09	0.00	0.05	-44.4	0.0	0.20	0.01	0.04	-100.0	0.0
331	Wood products except furniture	0.05	0.02	0.00	0.03	0.03	0.83	-62.6	26.7	0.07	0.06	0.83	-100.0	0.0
332	Furniture except metal	0.12	0.02	0.00	0.10	0.04	0.34	-88.5	22.9	0.12	0.06	0.48	-100.0	0.0
341	Paper and products	0.05	0.00	0.00	0.05	0.03	0.67	-97.5	12.2	0.06	0.03	0.41	-100.0	0.0
342	Printing and publishing	0.04	0.00	0.00	0.04	0.06	1.75	-100.0	0.0	0.06	0.05	0.91	-100.0	0.0
351	Industrial chemicals	0.07	0.01	0.00	0.06	0.04	0.70	-92.1	20.3	0.06	0.05	0.75	-100.0	0.0
352	Other chemicals	0.07	0.01	0.00	0.06	0.03	0.58	-83.6	25.5	0.07	0.03	0.41	-100.0	0.0
353	Petroleum refineries	0.05	0.01	0.00	0.04	0.04	1.01	-69.1	28.4	0.05	0.05	0.99	-100.0	0.0
354	Miscellaneous petroleum and coal products	0.07	0.00	0.00	0.07	0.04	0.63	-100.0	0.0	0.06	0.05	0.79	-100.0	0.0
355	Rubber products	0.09	0.05	0.00	0.04	0.02	0.56	-44.5	0.3	0.09	0.05	0.61	-100.0	0.0
356	Plastic products	0.14	0.07	0.00	0.07	0.02	0.33	-47.5	12.8	0.12	0.04	0.33	-100.0	0.0
361	Pottery china earthenware	0.09	0.05	0.00	0.04	0.01	0.14	-44.3	0.3	0.10	0.01	0.14	-100.0	0.0
362	Glass and products	0.07	0.03	0.00	0.04	0.03	0.80	-66.3	27.4	0.07	0.06	0.80	-100.0	0.0
369	Other non-metallic mineral products	0.06	0.01	0.00	0.05	0.04	0.80	-79.4	27.1	0.06	0.04	0.72	-100.0	0.0
371	Iron and steel	0.07	0.04	0.00	0.03	0.02	0.55	-48.9	14.2	0.07	0.04	0.53	-100.0	0.0
372	Non-ferrous metals	0.05	0.01	0.00	0.04	0.03	0.84	-76.3	27.6	0.04	0.04	0.97	-100.0	0.0
381	Fabricated metal products	0.08	0.03	0.00	0.05	0.03	0.58	-59.5	24.9	0.08	0.04	0.42	-100.0	0.0
382	Machinery except electrical	0.05	0.01	0.00	0.04	0.04	0.89	-89.5	21.8	0.04	0.04	0.95	-100.0	0.0
383	Machinery electric	0.07	0.02	0.00	0.05	0.04	0.73	-71.8	27.9	0.06	0.04	0.69	-100.0	0.0
384	Transport equipment	0.08	0.03	0.00	0.04	0.04	0.93	-55.5	22.6	0.08	0.07	0.90	-100.0	0.0
385	Professional and scientific equipment	0.06	0.03	0.00	0.03	0.03	0.92	-54.4	22.1	0.06	0.05	0.90	-100.0	0.0
390	Other manufactured products	0.08	0.04	0.00	0.05	0.04	0.79	-56.7	23.1	0.08	0.05	0.68	-100.0	0.0
	Total	0.09	0.03	0.00	0.05	0.03	0.63	-65.7	16.1	0.08	0.05	0.56	-100.0	0.0

Source: Authors' own calculation based on 8-digit HS product level data from UN-Trains. The statistics in columns (1), (2), (3) and (4) are based on 6936, 7118, 7678 and 6670 observations, respectively. Column (5) refers to 4607 observations, while the statistics in Columns (6) and (7) are based on 3891 observations each. The number of observations included is based on data availability (UN-Trains). Columns denoted with 'mean' present the simple mean average over all 8-digit HS product lines pertaining to the respective industry.

Annex Table 2: Description of Variables and Data Sources

Variable	Abbreviation	Exact definition[c]
		Dependent variable
Bound MFN tariff rate	Δt_i	Reduction in bound 'Most Favoured Nation' (MFN) tariffs negotiated during the Uruguay Round.
		Explanatory variables
FTA good dummy variable	I_i	Indicator variable taking the value one if Canada granted (duty-free) preferential market access to the U.S. (and Mexico).
US import ratio	Im_i	Product level Canadian imports from the US over all Canadian imports ('94)
Reciprocity induced changes in market access	R_i	Import weighted percentage tariff reductions of Canada's principal suppliers between 1986 and 1994 multiplied by good i's export share of each principal supplier to Canada; finally, aggregation over all principal suppliers of good i.
Political economy variable	ΔX_I	Change in the elasticity weighted inverse import penetration ratio at an ISIC 3-digit industry level between 1978 (final phase Tokyo Round) and 1992 (final phase Uruguay Round). [a]
MFN externality variable	P_i	Change in the share of small exporters (i.e. non-top 5 exporters/suppliers) of product i to Canada between 1994 and 1988. P_i takes the value one if the above mentioned change is larger than the median change and zero otherwise.
		Instruments
Import dummy variable	D_i^{94}	Dummy variable indicating whether a product was imported by Canada from the US regardless of its preferential status (instrumental variable for I_i).
NTB dummy variable	D_i^{ntb93}	Dummy variable taking the value one if product i was subjected to a Canadian NTB in 1993 (instrumental variable for Ii).
NTB dummy variable	$D_i^{ntball93}$	Indicator variable taking the value one if product i was subjected to a Canadian NTB in 1993 towards all trading partners (instrumental variable for I_i).
NTB & Import dummy variable	$D_i^{ntball93}*D_i^{94}$	Combination of import and NTB indicator variables.
Scale economies	Δscale	Change in value added/number of firms (establishments) between 1981 and 1992 (instrumental variable for the political economy variable)
	Δscale*Δworld price	Interaction of the scale economies instrument with the average world price change per industry between 1992 and 1994 (instrumental variable for the political economy variable).
World prices	$\Delta worldprice_i$, $(\Delta worldprice_i)^2$, $(\Delta worldprice_i)^3$	HS 8-digit world prices changes calculated as changes in unit-values between 1992 and 1994 (instrumental variable for I_i).
Unilateral tariff reductions	R_i^{uni}	Reciprocity measurement as described above but this time focusing on import-weighed unilateral tariff reductions of UR participants undertaken between 1986 and 1992 only (instrumental variable for R_i).

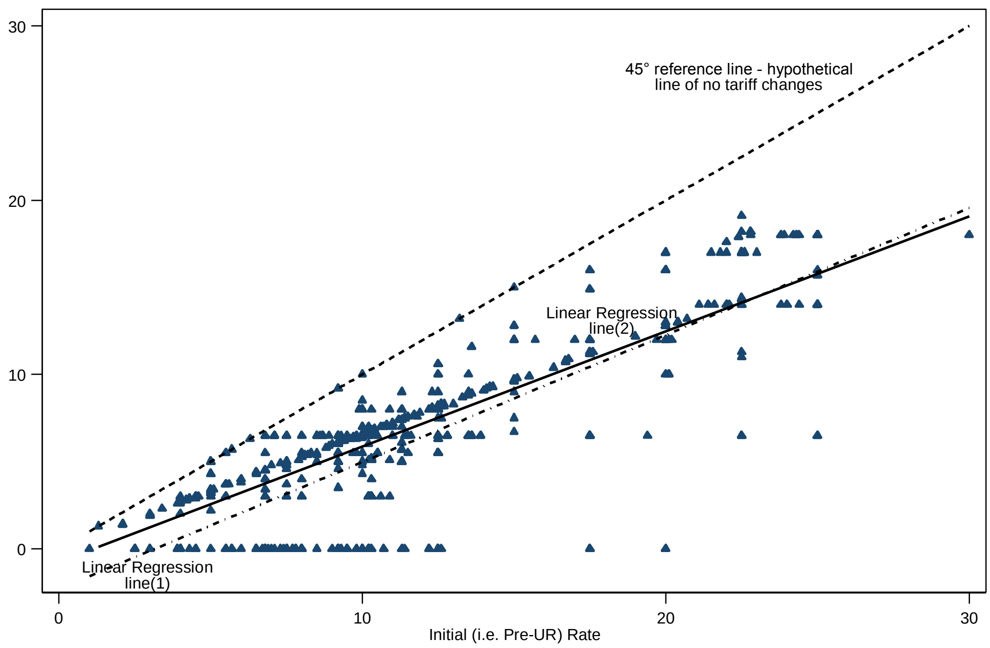
Notes: (a) The change in the elasticity weighed inverse import penetration ratio ΔX_I is calculated as $x_{92} - x_{78}$. (b) The change in the MFN externality effect or the change in the share of small (non-top5 exporters) of product-line i to Canada is calculated as $share_{94}$-$share_{89}$. (c) The variables are based on the authors' own calculations and the following data sources: WTO, TRAINS, COMEXT, UNIDO, as well as Finger et al. (2002) and Kee et al. (2009).

Annex Table 3: Summary Statistics

Variable	Mean	Std. Dev.	Min	Max
Δt_i	-0.06	0.03	-0.20	0.00
I_i	0.96	0.19	0.00	1.00
Im_i	0.66	0.33	0.00	1.00
ΔX_I	-2.27	3.58	-34.07	5.29
R_i	-0.45	0.08	-0.94	-0.02
D_i^{94}	0.99	0.06	0.00	1.00
R_i^{uni}	-0.23	0.07	-0.90	-0.01
D_i^{ntb}	0.38	0.48	0.00	1.00
D_i^{ntball}	0.36	0.48	0.00	1.00
$D_i^{ntball} * D_i^{94}$	0.36	0.48	0.00	1.00
$D_i^{ntball} * D_i^{naftaexp}$	0.35	0.48	0.00	1.00
Δp_{9294}	0.01	0.05	-0.40	1.22
Δscale	1.61	2.22	-13.67	49.11
P_i	0.12	0.32	0.00	1.00
$t_{i,t-1}$	0.13	0.06	0.01	0.30

The summary statistics are based on our dataset of 3138 observations.

Annex Figure 1: Canadian Pre- and Post-UR bound MFN Tariff Rates



Note: The graph illustrates the pre- and post UR (bound) MFN ad-alorem tariff rates and is based on our sample of 3138 observations. The dased 45°line illustrates a hypothetical line of no tariff changes. The dashed linear regression line (1) is based on the whole sample, whereas the solid linear regession line excludes products covered by 'zero-for-zero' concessions and thus by a different reduction rationale.
Source: WTO schedule of concessions and authors' own calculations.