The University of Nottingham

research paper series

Political Economy of Globalisation

Research Paper 2015/03

EU Regionalism and External Tariff Protection: The Role of Initial Tariffs and the Heterogeneity of Preferential Market Access

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Acknowledgements
The author is grateful to Chris Milner and Daniel Bernhofen, as well as Giovanni Facchini and Emanuel Ornelas for stimulating discussions and comments to earlier version of this paper.

This paper reflects the views of the author only and should not be attributed to the European Commission.

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Abstract

This paper provides evidence that the reduction in bound MFN tariffs is significantly larger for products with initially high tariff levels. These findings highlight the sensitivity of the PTA-external tariff question to the inclusion of initial tariffs, putting into perspective the previously identified stumbling block in a European context. Using micro-level trade and tariff data, we also find evidence that the type of trading partner, as well as the consideration of current or anticipated trade preferences matters.

JEL classification: F13, F14

Keywords: Regionalism, Preferential Trade Agreements, External tariff liberalisation, European Union.

Outline

1. Introduction

In view of the increasing popularity of preferential trade agreements (PTAs) globally during the last 20 years, a controversy has developed about whether regionalism promotes or hinders global trade liberalization. Once formed regional trade agreements may give rise to new and changing policy incentives, associated with the effects of PTAs on trade flows and welfare. In light of these consequences a theoretical literature has evolved providing arguments for tariff adjustments in both directions following the formation of a PTA.[1]

Empirical evidence on the subject-matter is at best limited. The small, but in importance growing, empirical literature on the relationship between PTAs and external tariffs finds rather mixed results, with some studies providing support for the 'building block' hypothesis (e.g. Estevadeordal et al., 2008; Calvo-Pardo et al., 2010), and others finding evidence for the contrary (Limão, 2006; Karacaovali and Limão, 2008). Reconciling some of these findings, more recent research argues that the impact of preferences is likely to be affected by the type of trading partner or partners the preferences have been offered to, and hence the associated policy context (Ketterer et al., 2014a&b), with a possible 'non-economic policy concession' motive resulting in smaller MFN tariff reductions, and the political rent destruction dimension of a PTA promoting larger cuts.[2]

In this paper, the primary focus is not necessarily to provide further evidence on these channels, but to examine the importance of existing external protection, and the definition of preferential market access in the context of the hypothesized PTA-external tariff nexus. We argue that the level of existing external (bound MFN tariff) protection may influence the magnitude of the offered tariff concessions, since, apart from accounting for the potential use of formulaic tariff cuts, policy-makers may find it easier to reduce higher initial tariffs more aggressively than lower ones.[3] We also examine the sensitivity of our results when controlling for sector-wide multilateral trade negotiations, which may have created their own dynamism, thereby providing an alternative tariff reduction rationale. Finally, we investigate

[1] For a recent review of the literature see Freund and Ornelas (2010).

[2] While non-economic PTA motives have been modelled by Limão (2007) who shows that PTAs may be used as vehicles to exchange trade preferences for a closer political cooperation on non-trade issues, in particular when granted to smaller trading partners. This exchange is likely to limit governments' incentives to reduce external tariffs, so that sizable preference margins, and hence the willingness to cooperate, can be maintained. The 'domestic political rent destruction' dimension of PTAs has been analysed by Ornelas (2005,a&b, 2008), who shows that PTAs, increase intra-bloc competition and hence the marginal benefit of external tariff protection for the domestic industry, which may, as a result, reduce lobbying efforts for higher tariffs. Ornelas shows that an increasing market share of the partner country's exporting producers amounts to the driving force of such a 'leakage of protectionist benefits' for the domestic industry.

[3] In the context of the PTA-external tariff nexus in Latin America, Crivelli (2012) finds that MFN tariffs are reduced up to three times more aggressively in high MFN tariff PTA members than in low tariff partners.

whether the distinction between existing and anticipated trade preferences plays a role when examining the robustness of the PTA – MFN tariff relationship.

To identify whether initial tariffs, sectoral aspects and the definition of trade preferences matter, we estimate the same model as Karacaovali and Limão (2008), but extend it to a consideration of initial tariffs throughout, sector-specific elements of multilateral trade negotiations, and varying definitions of preferentially traded goods. Our findings point to a less clear-cut support for a 'stumbling block' effect in a European context, when controlling for initial (i.e. pre-UR) tariff rates, and when considering different types of granted trade preferences. The results even suggest a positive, but imprecisely measured, building block effect, when focusing on EFTA trade preferences, where market access may have been more important than non-trade related political PTA motives.

The reminder of the paper is structured as follows. Section 2 outlines the empirical methodology, while section 3 provides a discussion of the results. Section 4 concludes.

2. Empirical methodology

2.1 Identification

The empirical model follows Karacaovali and Limão (2008) who examined Uruguay Round bound MFN tariff cuts for preferentially and non-preferentially traded goods in a European Union context. To identify the impact of initial (i.e. pre-UR) tariff protection we include the level of pre-Uruguay Round bound MFN tariff rates in the model.[4] The econometric specification is given by:

$$\Delta t_i = \alpha + \beta_1 I_i + \beta_2 \Delta x_h + \beta_3 R_i + \beta_4 \Delta P_i + \beta_5 t_{i,-1} + \mu_i \; , \qquad (1)$$

where Δt_i represents the absolute change in bound MFN tariffs negotiated during the UR. The main explanatory variable of interest is denoted by I_i and takes the value one if product i was granted duty-free preferential market access <u>and</u> was imported by the EU from the respective preferential partner country in 1994 – the end of the Uruguay Round.[5] To test

[4] For the anticipated (i.e.1996) preferential tariff specification the level of initial (i.e. pre-UR) tariff rates amounts to, on average, 8.1% for non-PTA goods, and to 7.8% for PTA goods. When focusing on current (i.e. 1994) and expected (i.e. 1996) preferential market access, the level of initial tariffs equals 10.8% for non-PTA goods and to 7.7% for PTA goods, on average.

[5] Note that we a define product to be imported in 1994 only if the registered trade value was above the 5th percentile of a particular country's exports to the EU, to account for potential product misclassifications when using very disaggregated (i.e. 8-digit HS level) information on trade flows and quantities.

whether the timing of the offered trade preferences, or the preferential trade policy setting in which the preferences have been offered matters, we introduce alternative specifications of the PTA good indicator variable. First, we contrast anticipated trade preferences, implemented in 1996, to a combination of current (i.e. 1994, end-UR) and expected (i.e. 1996) preferential market access.[6] We hypothesize that, in principle, the results should not be substantially affected by the varying time-related perspectives of preferential market access. Second, we consider preferences that were *only* granted to more competitive trading partners such as the EFTA countries. In light of recent research (Ornelas, 2005a&b, 2008; Ketterer et al, 2014 a&b) we argue that a leakage of protectionist benefits, and the associated domestic rent destruction, may have enabled more aggressive tariff cuts where trade preferences have been offered to countries where non-economic PTA policy incentives may have played a minor role, and intra PTA-bloc competitive forces were likely to prevail.

To control for alternative aspects, that are likely to influence tariff changes in multilateral trade negotiations, we introduce a set of control variables. Δx_h captures the change in political economy forces, defined as the change in the elasticity weighed inverse import penetration ratio at the ISIC 3-digit industry level between 1992 (final phase UR) and 1978 (end Tokyo Round) (i.e. $\Delta x_h = \Delta((X_h/M_h)/\epsilon_h)$).[7] R_i controls for tariff cuts driven by reciprocal trade concession motives, and has been calculated as $R_i = \sum_k s_{it}^k [\sum_i w_i^k \Delta t_i^k / t_i^k]$, where import weighted UR tariff concessions, aggregated over all products j ($\sum_j w_j^k \Delta t_j^k / t_j^k$) from 'principal supplier' (i.e. top-5 trading partner) country k, have been multiplied by the 1992 8-digit HS EU-import share in good i from country k (s_{it}^k), if the trading partner was one of the EU's top-5 suppliers, and by zero otherwise.[8] Free-riding behaviour based on the 'Most-Favoured-Nation' (MFN) clause may result in limited efforts to liberalize products with a large number of (smaller) foreign exporters. In order to capture this effect the variable P_i is introduced reflecting the change in the number of non-top 5 exporters per product line i between 1989 and 1994.[9] The intuition is that a smaller number of exporters to the EU would translate into larger tariff reductions on part of the EU, which may then, in return, be

[6] Note that when using preferential tariff information for 1996, preferential tariffs are still inter-acted with import values of 1994 in order to construct the PTA good indicator variable. Data on ACP, CACM, GSP and LGSP preferential tariffs granted by the EU are available for the both years (i.e. 1994 and 1996), while information on preferential tariffs offered to the MED, CEC, and EFTA countries is only available for 1996 using UNTRAINS information.

[7] Note that $\Delta(X_h/M_h)$ illustrates the inverse import penetration ratio change between 1992 and 1978, while ε_h denotes the import demand elasticity at the ISIC 3-digit industry level.

[8] It is assumed that the EU was only involved in direct negotiations with its five most important trade partners for each product line.

[9] P_i takes the value one if the above mentioned change is larger than the median change and zero otherwise.

reciprocated by more meaningful (reciprocal) tariff cuts on the part of the EU's direct negotiating partners. The term μ_i in Eq. (1) represents the idiosyncratic error term.

One of the major difficulties in establishing a link between trade preferences and multilateral tariffs is to establish causality, given that MFN tariffs themselves may influence whether a product receives a preference, and the fact that preferential market access is more valuable the higher the external MFN tariff rates, countries may be more inclined to ask for a preference on goods for which they expect smaller MFN tariff cuts. We use IV estimation techniques and a series of instrumental variables to account for potential reverse causality issues.[10] To control for potential endogeneity concerns affecting the reciprocity term unilateral tariff cuts undertaken between 1986 and 1992 are used. These were implemented at a time when the conclusion of the UR was very uncertain, but they were later credited on the negotiated total tariff reductions. Finally, for the political economy variable the change in industry scale economies between 1981 and 1992 and its interaction with world prices are used as instruments.

2.2 Data

The selection of variables and data follows Karacaovali and Limão (2008). Pre-UR and post-UR bound ad-valorem MFN tariff rates are available from the WTO's schedule of concessions database.[11] The PTA-good indicator variables, constructed at an 8-digit HS level, have been built using information on preferential tariff rates provided by the UNCTAD-TRAINS database and information on trade flows provided by Eurostat's Comext database. Sector-level import, production, value added and establishment data, used to control for political economy forces are provided by the UN-COMTRADE and UNIDO databases. Sector-specific import demand elasticicities are from Kee et al. (2009), and import-weighted UR tariff cuts from Finger et al. (2002) have been used to calculate the reciprocity variable. Annex Table 1 provides an overview of the variables used, and their exact definitions and

[10]These include dummy variables indicating whether a product was (i) imported in 1994 (D_i^{94}), (ii) was subject to an NTB in 1993 (D_i^{ntb93}), as well as (iii) the product level change in world prices. While the import dummy D_i^{94} is directly linked to the PTA good indicator (i.e. $I_i = PR_i*D_i^{94}$), it is uncorrelated to the error term since the MFN tariff reductions entered into force from 1995 onwards. Countries are more likely to ask for preferences for good which they suspect to be subject to NTBs in the future – as a proxy for future NTBs 1993 data is used. We use an additional instrument inter-acting the NTB indicator variable (D^{ntb93}) with the export dummy variable (D_i^{94}) World price changes between 1992 and 1994 are introduced since they are likely to determine the pecuniary benefit arising from a granted preference.

[11] Products with zero initial (i.e. pre-UR) MFN tariff rates, and agricultural products were excluded from the analysis.

data sources.[12]

3. Results

Table 1 reports the estimation results for Eq. (1) using heteroscedasticity-robust two-step efficient generalized methods of moments (IV-GMM) estimators, and clustering of standard errors at the sector level. The first two Columns show the findings when considering anticipated preferential market access, while Columns (3) and (4) take into account current as well as expected trade preferences. Not considering the level of initial tariff protection, both time-related specifications show positive parameter estimates pointing to larger tariff cuts for products not traded under preferential market access (Columns 1 and 3). Despite being statistically highly significant the identified stumbling block effects, however, tend to show some divergence in terms of magnitude with larger non-PTA good tariff cuts of about 1.9 percentage points when only considering expected (i.e. 1996) preferences, compared to 1.2 percentage points when allowing for both, the consideration of current (i.e. 1994) and anticipated (i.e.1996) preferential trade concessions at the time of the Uruguay Round.[13]

Columns (2) and (4) report the results when controlling for the level of initial (i.e. pre-UR) tariff protection, and hence the potential use of formulaic tariff cuts, or smaller political economy resistance towards tariff cuts where initial tariffs levels were very high. The results show that, for both time-related preference specifications, the level of initial tariff protection explains a significant part of the variation in external MFN tariff cuts over the PTA and non-PTA good subgroups, as the PTA good indicator variable becomes statistically insignificant, and substantially smaller in size. The reported parameter estimates, point to an about two times smaller stumbling block effect (Columns 2 and 4). The absence of a significant stumbling block effect in these cases is surprising and highlights a notable sensitivity in findings to the inclusion of pre-Uruguay Round MFN tariff levels. Furthermore, we find that the level of initial tariffs is statistically highly relevant in all model specifications, which lends further support to the hypothesized importance to account for initial tariff rates.

In light of recent research, which provides evidence for the view that the effect of trade preferences on multilateral trade liberalization is linked to the motivation and policy context of a PTA (see Ketterer et al., 2014 a&b), we extend our analysis to an examination of

[12] Summary statistics are provided in Annex Table 2.
[13] For comparison, Karacaovali and Limão (2008) find, in their baseline estimations, a stumbling block of 1.5 percentage points when using 1996 preferences.

trade preferences which have been granted to the arguably most competitive preferential trading partners in a European context, at the time of the Uruguay Round, – i.e. the EFTA trade partners.[14] Columns (5) and (6) report the findings when focusing on duty-free preferences that were *only* granted to EFTA countries. The results show negative, but imprecisely measured, parameter estimates pointing to larger tariff cuts for products that were only offered to EFTA member states. Accordingly, these results suggest a building block effect, which is however not statistically significant at the usual thresholds. These findings may provide partial evidence for the argument that a political rent destruction effect, and the associated leakage of domestic political rents, á la Ornelas (2005a&b), may amount to the driving force in determining the PTA-external tariff relationship, in a setting where non-trade related motives (Limão, 2006) may have played a minor role.

The findings for the introduced control variables in Table 1 are largely in line with the previous results in the literature. The findings for political economy factors point to a statistically highly significant impact of political economy forces with positive IV estimates ranging from 0.012 to 0.019, suggesting significantly smaller tariff cuts where domestic industry interests were particularly important. Evidence for reciprocal tariff reduction, and for the hypothesized 'MFN externality effect' remain, however, weak as the respective parameter estimates do not fall within the usual bands of statistical significance.

Annex Table 4 provides a series of sensitivity tests for the results described above; with the findings for the varying PTA good indicator specifications displayed in the table, and the controls supressed (but available upon request). Column (2) reports the findings when focusing on products that actually changed their preference status between the last successfully concluded multilateral trade rounds (i.e. the Tokyo and the Uruguay Round), which, given the econometric specification in first differences, may be more aligned with the empirical identification strategy.[15] The results show an insignificant stumbling block effect when using the anticipated (i.e. 1996) preference good specification, which is not sensitive to the inclusion of initial MFN tariffs. When considering current (i.e. 1994) as well as anticipated (i.e. 1996) preferences, the results show a weakly significant stumbling block effect even when controlling for initial tariffs (Column 2), and may hence point to a certain

[14] The European Free Trade Association (EFTA) includes Iceland, Liechtenstein, Norway, and Switzerland.

[15] To account for the change in preference status, an export dummy for Central European Countries is used as an instrument for the PTA-good indicator variable, instead of the general 1994 import dummy. The intuition is that preferences offered to former Communist (i.e. the CEC) countries were largely unexpected at the time of the Tokyo-, and fully implemented at the time of the Uruguay-Round, highlighting their potential use as an instrument.

sensitivity of the stumbling block findings to the time-related varying perspectives of preferential market access.

Columns (3) to (6) report the results when taking into account the presence of an alternative tariff reduction strategy based on the UR-negotiated sectoral agreement. The results, when excluding product lines that were affected by these negotiations, show in all model specifications a positive and statistically highly significant stumbling block effect when not controlling for the level of initial MFN tariffs. Taking into account the level of existing tariff protection, however, renders the identified stumbling block effect statistically innocuous, and shows, in terms of magnitude, much smaller parameter estimates.

Table 1: External Tariff Cuts and Trade Preferences

	IV-GMM					
	(1)	(2)	(3)	(4)	(5)	(6)
	Anticipated (1996)	Anticipated (1996)	Current & Anticip. (1994 & 1996)	Current & Anticip. (1994 & 1996)	Only EFTA (1994 & 1996)	Only EFTA (1994 & 1996)
$I_i^{‡}$	0.019***	0.008	0.012***	0.006	-0.046	-0.006
	(0.006)	(0.006)	(0.005)	(0.004)	(0.043)	(0.055)
R_i‡	0.0001	-0.005	-0.004	-0.005	-0.007	-0.007
	(0.009)	(0.008)	(0.007)	(0.007)	(0.006)	(0.007)
$\Delta \mathbf{x}$‡	0.015***	0.018***	0.017***	0.019***	0.012**	0.017**
	(0.004)	(0.006)	(0.005)	(0.006)	(0.005)	(0.007)
P_i	-0.002	-0.001	-0.001	-0.0004		-0.205***
	(0.001)	(0.001)	(0.001)	(0.001)		(0.046)
$t_{i,-1}$		-0.160**		-0.161**	-0.001*	-0.0002
		(0.077)		(0.064)	(0.001)	(0.001)
Constant	-0.029***	-0.008	-0.028***	-0.007	-0.020***	-0.001
	(0.006)	(0.012)	(0.006)	(0.010)	(0.005)	(0.011)
Observations	6329	6329	6329	6329	6329	6329
Number of PTA goods	4519	4519	5975	5975	162	162
Hansen's J (p-val.)[a]	0.538	0.591	0.606	0.681	0.714	0.631
C-stat (p-val.)[b]	0.55	0.524	0.644	0.557	0.522	0.504
Endogeneity (p-val.)[c]	0.571	0.657	0.615	0.782	0.239	0.493
Heterosked. (p-val.)[d]	0.000	0.000	0.000	0.000	0.000	0.000

All regressions have been run using heteroskedasticity robust standard errors clustered at the 3-digit ISIC industry level. *,**,*** illustrate the 10%, 5%, 1% significance levels, respectively. The regressions are based on an instrumental variable efficient generalized method of moments' estimator. (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. (c) Endogeneity test of the endogenous regressors marked with ‡. The null hypothesis says that the marked endogenous regressors can actually be treated as exogenous. (d) Pagan and Hall's (1983) test of heteroskedasticity for estimations using instrumental variables (IV). The null hypothesis is that no heteroskedasticity is present.

The results reported in Table 2 account for the magnitude of the granted preferential market access, and hence examine whether the importance of the offered trade preferences matters. Table 2 reports the findings when focusing on the combination of current and anticipated EU duty-free preferences. To identify 'important' preferential market access concessions we impose an additional PTA-partner specific import share threshold of 40% to identify preferentially traded goods.[16] Accounting for 'significant' preferential imports, the results in Columns (1) and (2) tend to provide evidence for larger MFN tariff cuts for non-preferentially traded goods of around 1.3 percentage points when not considering initial tariffs, and of around 0.8 percentage points when accounting for the level of pre-UR tariff protection, which are both statistical significant at levels of 1 and 10%, respectively.[17]

Moreover, we use an alternative measure of significant preferential market access by considering products only to be preferentially imported when the preference margin amounts to at least two percentage points.[18] Following Estevadeordal et al., (2008) and Calvo-Pardo et al. (2010) it is possible that rules of origin may render some preferences innocuous, and that accounting for a significant difference between MFN tariffs and preferential tariffs may represent a more adequate measure of 'significant' market access. The findings are reported in Columns (3) and (4), and show larger tariff reductions for non-preferentially traded goods of around 1.2 and 0.8 percentage points, both significant at the 5% threshold, when omitting, and when accounting for level of initial MFN tariff protection, respectively.

When analysing the importance of the offered preferential market access offered to the EFTA trade partners, previous findings tend to be confirmed. Columns (5) and (6) use the 40% import share to identify preferentially traded goods, while Columns (7) and (8) report the findings when focusing on 'significant' preference margins. The results lend further support to the building block hypothesis in a context of more competitive trade partners. As shown in Table 1, the identified effect, however, again tends to be imprecisely measured at the usual levels of significance.

[16] Qualitatively similar results are obtained when using thresholds of 25%.

[17] Note that when using the anticipated (i.e.1996) PTA good specification we find a stumbling block effect of around 2.1 percentage points, which, however, becomes statistically insignificant, and substantially smaller in size when accounting for the level of initial tariffs. The results are not reported here, but available upon request.

[18] Taking into account the possibility of different preferential tariffs granted for the same product line, we follow Estevadeoral et al. (2008) by using the lowest granted preferential tariff to calculate the preference margin.

A series of robustness tests for the significant preferential market access specifications is reported in Annex Table 5.[19] The estimating results when focusing on goods that switched their PTA status (Column 2) show, for all specifications, an in size substantially larger stumbling block effect, compared to the baseline specifications reported in Column (1). This effect, however, only remains robust to the inclusion of initial tariffs when focusing on significant preference margins, and fails to fall within the usual bands of statistical significance when examining significant import flows. When controlling for the UR sector level negotiations by excluding the relevant product lines, we find for most specifications that the stumbling block effect becomes statistically insignificant when controlling for initial MFN tariff levels; and even shows an opposing negative, and statistically weakly significant, sign when focusing on significant imports and excluding goods that were covered under the Textiles and Clothing Agreement (ATC).[20]

[19] In line with Annex Table 4 we focus on goods which changed their preference status between the Tokyo and the Uruguay Round, and alternative tariff reductions rationales based on the UR sectoral negotiations.

[20] Note that the stumbling block effect remains robust to the inclusion of initial tariff rates, in the model specification which excludes products that were either part of the 'Zero-for-Zero' negotiations or the agreement on chemicals (Column 4).

Table 2: External Tariff Cuts and Trade Preferences - Significant Imports & Preferences

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Signif. IM (1994 & 1996)	Signif. IM (1994 & 1996)	Signif. Prefs. (1994 & 1996)	Signif. Prefs. (1994 & 1996)	Signif. EFTA IM (1994 & 1996)	Signif. EFTA IM (1994 & 1996)	Signif. EFTA Prefs. (1994 & 1996)	Signif. EFTA Prefs. (1994 & 1996)
$I_i^{‡}$	0.013***	0.008*	0.012**	0.008**	-0.004	-0.009	-0.046	-0.006
	(0.005)	(0.005)	(0.005)	(0.003)	(0.029)	(0.027)	(0.043)	(0.055)
R_i‡	-0.015**	-0.009	-0.005	-0.004	-0.009	-0.006	-0.007	-0.007
	(0.007)	(0.007)	(0.008)	(0.007)	(0.007)	(0.007)	(0.006)	(0.007)
Δx‡	0.016***	0.021***	0.015***	0.019***	0.014**	0.020***	0.012**	0.017**
	(0.005)	(0.006)	(0.006)	(0.005)	(0.006)	(0.007)	(0.005)	(0.007)
P_i	-0.001	-0.0003	-0.001	-0.001	-0.001	-0.0001	-0.001*	-0.0002
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
$t_{i,-1}$		-0.180***		-0.131***		-0.196***		-0.205***
		(0.059)		(0.045)		(0.060)		(0.046)
Constant	-0.032***	-0.007	-0.029***	-0.011**	-0.020***	0.001	-0.020***	-0.001
	(0.007)	(0.010)	(0.007)	(0.005)	(0.006)	(0.009)	(0.005)	(0.011)
Observations	6329	6329	6329	6329	6329	6329	6329	6329
Number of PTA goods	5002	5002	5955	5955	27	27	162	162
Hansen's J (p-val.)[a]	0.587	0.806	0.569	0.7500	0.557	0.706	0.714	0.631
C-stat (p-val.)[b]	0.471	0.798	0.558	0.7311	0.453	0.702	0.522	0.504
Endogeneity (p-val.)[c]	0.541	0.754	0.570	0.7434	0.707	0.782	0.239	0.493
Heterosked. (p-val.)[d]	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000

All regressions have been run using heteroskedasticity robust standard errors clustered at the 3-digit ISIC industry level. *.**,*** illustrate the 10%, 5%, 1% significance levels, respectively. The regressions are based on an instrumental variable efficient generalized method of moments' estimator. (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 C-statistic is defined as the difference of the Sargan-Hansen value of the equation with the restricted 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 of the endogenous regressors marked with ‡. (d) Pagan and Hall's (1983) test of heteroskedasticity for estimations using instrumental variables (IV).

7. Conclusions

In light of limited empirical evidence on the external tariff-PTA nexus, this paper examines the role of initial tariff protection, sector level tariff negotiations, and varying definitions of trade preferences in the context of the stumbling block – building block debate of preferential trade agreements. We find that controlling for the level of initial tariff protection explains a considerable part of the variation in bound MFN tariff cuts across the preferentially and non-preferentially traded sub-samples of goods, which partially puts the identified stumbling block effect in a European context into perspective. The absence of a significant 'stumbling block' effect, when considering overall EU preferences, and accounting for the level of initial tariff protection is somewhat surprising, but suggests that the level of initial tariff protection may have played an important role when agreeing upon the final bound MFN tariff cuts during the Uruguay Round.[21]

These findings tend to be confirmed when considering different timeframes of the granted preferential market access; with expected preferences (i.e. preferences granted shortly after the end of the UR in 1996) showing, by trend a smaller and, partially, less robust stumbling block effect, compared to the consideration of both, current (i.e. 1994) and expected (i.e. 1996) preferential market access concessions. When considering the importance of the granted preferential market access – measured by significant product-level import shares, or preference margins – our estimations show more robust stumbling block findings, which are however partially undone when accounting for the UR sector level agreements.

Finally, in line with more recent research (Ornelas, 2005a&b; Ketterer et al., 2014a&b), we also find partial evidence for a positive, however imprecisely measured, building block effect, where preferential market access has only been granted to the arguably most competitive EU preferential trading partners at the time of the UR (i.e. the EFTA countries), pointing to the importance of the PTA policy context when analysing the trade preferences – external tariff nexus. Overall, our findings hence point to a certain sensitivity of results, subject to the heterogeneity in type and importance of the offered preferential market access.

[21] The level initial tariff protection is statistically highly significant in all model specifications, with higher tariffs being reduced more aggressively than smaller ones.

References

Calvo-Pardo H., Freund C., Ornelas, E., 2010, The ASEAN free trade agreement: impact on trade flows and external trade barriers, R. Barro, J. Lee (Eds.), Costs and Benefits of Regional Economic Integration, Oxford University Press, Oxford.

Crivelli, P., 2012, Regionalism and Falling External Protection in High and Low Tariff Members, University of Geneva, working paper series.

Estevadeordal, A., Freund, C., Ornelas, E., 2008, Does regionalism affect trade liberalization towards non-members? Quarterly Journal of Economics, 1531–1575.

Finger, M.J., Reincke U., Castro A., 2002, Market access bargaining in the Uruguay round: how tightly does reciprocity constrain? J. Bhagwait (Ed.), Going alone: The Case for Relaxed Reciprocity in Freeing Trade, MIT Press, London.

Freund, C., Ornelas, E., 2010, Regional Trade Agreements, Annual Review of Economics, 2(1),139-166.

Karacaovali B., Limão N., 2008, The clash of liberalizations: preferential vs. multilateral trade liberalization in the European Union, Journal of International Economics, 74, 299–327.

Kee, H.L., Nicita, A., Olarreaga, M., 2009, Estimating trade restrictiveness indices, Economic Journal, 119, 172–199.

Ketterer, T., Bernhofen, D., Milner, C., 2014, Preferences, Rent Destruction and multilateral liberalization: the building block effect of CUSFTA, Journal of International Economics, 92, 63-77.

Ketterer, T., Bernhofen, D., Milner, C., 2014, The Impact of Trade Preferences on Multilateral Tariff Cuts: Evidence for Japan, GEP working paper.

Limão, N., 2006, Preferential trade agreements as stumbling blocks for multilateral trade liberalization: Evidence for the U.S., American Economic Review, 96, 896–914.

Limao, N., 2007, Are preferential trade agreements with non-trade objectives a stumbling block for multilateral liberalization? Review of Economics Studies., 74, 821–855.

Ornelas, E., 2005a, Rent destruction and the political viability of free trade agreements, Quarterly Journal of Economics, 120, 1475–1506.

Ornelas, E., 2005b, Endogeneous free trade agreements and the multilateral trading system, Journal of International Economics, 67, 471–497.

Ornelas, E., 2008, Feasible multilateralism and the effects of regionalism, Journal of International Economy, 74, 202–224.

Appendix

Annex Table 1: Description of variables and their sources

Variable	Abbreviation	Exact definition	Source
		Dependent variable	
Bound MFN tariff rate reductions	Δt_i	Reduction in bound 'Most Favoured Nation' (MFN) tariffs negotiated during the Uruguay Round and those in place before the Uruguay Round (i.e. Tokyo Round)	WTO + authors' own calculations
		Explanatory variables	
PTA good dummy variable	I_i	Indicator variable taking the value one if a product was granted duty-free preferential access under any of the EU's PTAs in 1994 (and/or) 1996 and was imported by the EU in 1994 (from the respective partner country)	TRAINS + COMEXT
Reciprocity induced changes in market access	R_i	Import weighted percentage tariff reductions of the EU's principal suppliers between 1986 and 1994 multiplied by good i's export share of each principal supplier to the EU; finally aggregation over all principal suppliers of good i	Finger et al. (2002) + COMEXT + authors' own calculations
Political economy variable	Δx_h	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) [22]	COMTRADE + UNIDO + Kee et al. (2009) + authors' own calculations
MFN externality variable	P_i	Change in the share of small exporters (i.e. non-top 5 exporters/suppliers) of product i to the EU between 1989 and 1994. P_i takes the value one if the above mentioned change is larger than the median change and zero otherwise.[23]	COMEXT + authors' own calculations
		Instruments	
Import dummy variable	D_i^{94}	Dummy variable indicating whether a product was imported by the EU from the respective partner country (instrumental variable for I_i)	COMEXT + authors' own calculations
NTB dummy variable	D_i^{ntb93}	Dummy variable taking the value one if product i was subjected to an EU-NTB in 1993 (instrumental variable for I_i)	TRAINS + authors' own calculations
NTB dummy variable	$D_i^{ntball93}$	Indicator variable taking the value one if product i was subjected to an EU-NTB affecting all trading partners in 1993 (instrumental variable for I_i)	TRAINS + authors' own calculations
NTB & Import dummy variable	$D^{ntball93}{*}D^{94}$	Combination of import and NTB indicator variables	TRAINS + authors' own calculations
Scale economies	Δscale	Change in value added/number of firms (establishments) between 1981 and 1992 (instrumental variable for the political economy variable)	UNIDO + authors' own calculations
	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)	UNIDO + COMEXT + authors' own calculations
World prices	Δworld price$_i$, $(\Delta\text{world price})_i^2$, $(\Delta\text{world price})_i^3$	HS 8-digit world price changes calculated as changes in unit-values between 1992 and 1994 (instrumental variable for I_i)	COMEXT + authors' own calculations
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)	Finger et al. (2002) + COMEXT + authors' own calculations

[22] The change in the elasticity weighed inverse import penetration ratio Δx is calculated as $x_{92} - x_{78}$.
[23] The change in the MFN externality effect or the change in the share of small (non-top5 exporters) of product-line i to the EU is calculated as share$_{94}$-share$_{89}$.

Annex Table 2: Summary statistics

Variable	Mean	Std. Dev.	Min	Max
Δt	-0.030	0.022	-0.268	0
I_i^{96}	0.714	0.452	0	1
$I_i^{94\&96}$	0.944	0.230	0	1
I_i^{efta}	0.026	0.158	0	1
Δx_h	-0.906	0.630	-6.887	2.140
R_i	-0.463	0.104	-0.932	0
P_i	0.509	0.500	0	1
t_{t-1}	0.079	0.047	0.005	0.65

The dataset includes 6329 observations.

Annex Table 3: External tariff cuts and trade preferences - Significant imports and 1996 preferential market access

	IV-GMM	
	(1)	(2)
	Signif. Imports (1996)	Signif. Imports (1996)
$I_i^{\ddagger}$	0.021***	0.011
	(0.006)	(0.007)
$R_i^{\ddagger}$	-0.014	-0.009
	(0.009)	(0.008)
$\Delta x^{\ddagger}$	0.012**	0.020***
	(0.005)	(0.007)
P_i	-0.002	-0.001
	(0.001)	(0.001)
$t_{i,-1}$		-0.186**
		(0.080)
Constant	-0.037***	-0.008
	(0.007)	(0.015)
Observations	6329	6329
Number of PTA goods	3838	3838
Hansen's J (p-val.)[a]	0.586	0.708
C-stat (p-val.)[b]	0.749	0.686
Endogeneity (p-val.)[c]	0.627	0.783
Heterosked. (p-val.)[d]	0.000	0.000

All regressions have been run using heteroskedasticity robust standard errors clustered at the 3-digit ISIC industry level. *,**,*** illustrate the 10%, 5%, 1% significance levels, respectively. The regressions are based on an instrumental variable efficient generalized method of moments' estimator.

Annex Table 4: Industry and sector-specific considerations

	IV-GMM					
	(1)	(2)	(3)	(4)	(5)	(6)
Sensitivity test	Baseline	'Sure Switches'	'Zero-for-Zero' Sectoral Agreements	Sectoral Agreements incl. Chemicals	Agreement on Textiles and Clothing (ATC)	Sectoral Agreements incl. Chemicals & ATC
$I^{96\ddagger}_{i}$	0.019***	0.021	0.030***	0.029***	0.018***	0.027***
	(0.006)	(0.013)	(0.005)	(0.005)	(0.005)	(0.005)
$I^{96\ddagger}_{i}$ & initial tariffs	0.008	0.008	0.001	0.00782	-0.003	-0.004
	(0.006)	(0.010)	(0.004)	(0.005)	(0.002)	(0.003)
$I^{94\&96\ddagger}_{i}$	0.012***	0.027**	0.021***	0.029***	0.016***	0.030***
	(0.005)	(0.013)	(0.004)	(0.004)	(0.004)	(0.003)
$I^{94\&96\ddagger}_{i}$ & initial tariffs	0.006	0.019*	0.002	0.014***	0.002	0.003
	(0.004)	(0.011)	(0.004)	(0.005)	(0.002)	(0.004)
Observations	6329	6329	5059	4073	5045	2806

All regressions use heteroskedasticity robust standard errors clustered at the 3-digit ISIC industry level. *,**,*** illustrate the 10%, 5%, 1% significance levels, respectively. Column (1) reports the baseline results of Table 1. Column (2) introduces a new instrument for the duty-free PTA good indicator - i.e. imports from CEC partner countries (D_{cec}) thereby explicitly focusing on products that have switched their PTA-status between the Tokyo- and Uruguay Round. Column (3) to (6) illustrate different regression results when subjecting the baseline results, displayed in Column (1), to various sensitivity tests regarding the potential influence of sectoral agreements. Tariff lines covered by the respective sectoral agreement have been excluded.

Annex Table 5: Industry and sector-specific considerations – Significant imports and preference margins

	IV-GMM					
	(1)	(2)	(3)	(4)	(5)	(6)
Sensitivity test	Baseline	'Sure Switches'	'Zero-for-Zero' Sectoral Agreements	Sectoral Agreements incl. Chemicals	Agreement on Textiles and Clothing (ATC)	Sectoral Agreements incl. Chemicals & ATC
$I^{94\&96\ddagger}_{i}$ (Sign.IM)	0.013***	0.032**	0.024***	0.035***	0.023***	0.033***
	(0.005)	(0.014)	(0.005)	(0.004)	(0.006)	(0.004)
$I^{94\&96\ddagger}_{i}$ (Sign.IM) & initial tariffs	0.008*	(0.020)	0.007	0.022***	-0.004*	0.004
	(0.005)	(0.013)	(0.006)	(0.006)	(0.002)	(0.006)
$I^{94\&96\ddagger}_{i}$ (Sign.Margins)	0.012**	0.025**	0.020***	0.028***	0.014***	0.029***
	(0.005)	(0.012)	(0.005)	(0.004)	(0.004)	(0.003)
$I^{94\&96\ddagger}_{i}$ (Sign.Margins) & initial tariffs	0.008**	0.016**	0.004	0.014***	0.002	0.002
	(0.003)	(0.008)	(0.004)	(0.005)	(0.002)	(0.003)
Observations	6329	6329	5059	4073	5045	2806

All regressions use heteroskedasticity robust standard errors clustered at the 3-digit ISIC industry level. *,**,*** illustrate the 10%, 5%, 1% significance levels, respectively. Column (1) reports the baseline results of Table 2. Column (2) introduces a new instrument for the duty-free PTA good indicator - i.e. imports from CEC partner countries (D_{cec}) thereby explicitly focusing on products that have changed their PTA-status between the Tokyo- and Uruguay Round. Column (3) to (6) illustrate different regression results when subjecting the baseline results, displayed in Column (1), to various sensitivity tests regarding the potential influence of sectoral agreements. Tariff lines covered by the respective sectoral agreement have been excluded.