# The Effect of the Canada-US Free Trade Agreement on Canadaian Multilateral Trade Liberalization

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**Abstract**

In this paper we analyze three different channels for the effect of the Canada-US Free Trade Agreement (CUSFTA) on Canadian trade policy. We first show that the CUSFTA tariff preferences have triggered a decline in Canadian external tariffs. Our results imply that Canadian tariff preferences to the US can explain a two percentage point reduction in the external tariff between 1989 and 1998. Second, we found that Canadian external tariffs declined slower in industries which generate more revenue for the US exporters to Canada. This finding provides the evidence on cooperation in trade policies between the US and Canada. Finally, we estimate the effect of the CUSFTA on the intensity of industrial lobbying for trade policy in Canada and find no relationship between preferential trade iberalization and lobbying activity. Overall, we demonstrate that the CUSFTA generated a considerable reduction in external tariffs and contributed to freer trade.

# 1 Introduction

Preferential trade agreements (PTA) have flourished around the world ever since the first one was established in 1958. Extensive research has been done on the implications of PTAs on multilateral trade liberalization and the welfare of member countries. However, neither theoretical nor empirical analysis has reached a consensus on whether or not joining a PTA would make a country more or less open to trade with non-members. The theoretical literature proposed several channels for the effect of a PTA on external tariffs and while some of them lead to acceleration in trade liberalization towards non-member countries (Bagwell and Staiger, 1997b; Freund, 2000; Bond, Riezman, and Syropoulos, 2004; Ornelas, 2005a), others have the opposite effect (Panagariya and Findlay, 1994; Grossman and Helpman, 1995; Krishna, 1998; Limao, 2006). As the impact of a PTA on trade policy may vary across agreements, it has become necessary to provide more country-specific analysis in order to understand the welfare implication of a PTA for a particular country.

In this paper, we study the effect of the Canada-US Free Trade Agreement (CUSFTA) on Canada's multilateral trade liberalization (MTL). We develop a simple theoretical model with endogenous trade policy formation which incorporates the main channels, identified in previous literature, through which a Free Trade Agreement (FTA) can affect the MTL of its member countries. Using Canadian trade data to test predictions of this model, we find the evidence of both "building block" and "stumbling block" effects of the agreement. Specifically, we demonstrate that the CUSFTA tariff preferences have resulted in considerable reduction in Canadian tariffs, explaining 2.21 out of 4.02 percentage point decline in the average MFN tariff rate between 1989 and 1998. We also show that the MFN tariff reductions were smaller in industries which generate more exports revenue for US exporters. This finding demonstrates the tendency for the Canadian government to maintain higher preference margins for products which have greater potential to generate rent to the partner country. Nevertheless, our results imply that this effect is dominated by the "building block" effect and that, despite slower trade liberalization in some industries, the CUSFTA formation had induced more open MTL policies in Canada.

To outline our theoretical model, we consider an economy of three large countries and monopolistically competitive markets. We start by deriving the effect of an FTA on external tariffs if they are set non-cooperatively by welfare-maximizing governments. In this framework there are two sources of complementarity between external tariffs and preferential tariffs. First, there is a terms-of-trade and tariff revenue effects, similar to the ones obtained by Richardson (1993) and Bagwell and Staiger (1997b), which cause in MFN tariffs to fall alongside the preferential trade liberalization. A reduction in the MFN tariffs following the CUSFTA can moderate some efficiency lost due to the distortionary price discrimination caused by preferential tariffs (Bagwell and Staiger (1997b), Freund (2000), Ornelas (2005b)) and restore part of tariff revenue lost due to the shift in import demand to US goods (Richardson (1993), Bagwell and Staiger (1997b)). Second, there is a market structure effect which stimulates policymakers to raise protection in industries with large domestic presence and low degree of product differentiation in order to redistribute consumer expenditure from foreign to domestic varieties. Since in the presence of an FTA the MFN tariff targets only the rest of the world (ROW) firms and part of the protection benefits flow to the partner country firms, the FTA will reduce the redistributive power of the MFN tariff and contribute to deeper MTL.

Next, we extend the model by introducing cooperative motives in trade policy formation in the spirit of Limao (2007). Limao argued that by taking the interests of FTA partner country's firms into consideration while choosing a trade policy, policymakers may induce the partner country's government to cooperate on non-trade policy issues. Alternatively, the government of one country can set trade policy in the interest of another in expectation that the partner will act similarly. When FTA member countries internalize the effect of their policies on each other in this way, they may want to keep external tariffs high in order to stimulate cooperation by the partner country. Similarly to Limao (2007), who demonstrated that trade agreements with cooperative trade policies tend to become more protectionist in the context of multilateral trade negotiations, this "stumbling block" of an FTA enters the expression for the optimal tariff in additively separable way. This allows to empirically identify its effect independently from the other two described in the previous paragraph.

As a last extension of our model, we incorporate political economy factors in policymakers' preferences based on the works by Grossman and Helpman (1994, 1995). The theory predicts that the strength of domestic lobbying for protection is inversely related to the measure of import penetration. Hence, increased imports from a partner country following formation of an FTA should weaken the lobbying power of domestic special interest groups and reduce the level of protectionism. Ornelas (2005b) calls this effect "rent destruction" since in the presence of an FTA part of the rent from protection will flow to the partner country firms, making lobbying less attractive and weakening political economy distortion.

Our model identifies and generates testable predictions for three factors that lead to complementarity between multilateral and preferential tariffs and one factor contributing to substitutability. While most of the existing literature focus on a few factors contributing to FTA trade policy formation, the richness of our theoretical model allows us to analyze all four factors of FTA's trade policy summarized above in a unified structural framework and compare their relative contribution to observed tariff changes during the implementation stage of the CUSFTA. Estimating the model with Canadian tariff cuts following introduction of the CUSFTA, the main findings of this study are the following. First, the study reveals strong tariff complementarity between Canadian preferential and MFN tariffs which propagates through the terms-of-trade and tariff revenue effects. This positive relationship between external and internal tariffs is robust across all of our specifications. Our results indicate that a one percentage point reduction in preferential tariffs was accompanied by 0.05 percentage points reduction in MFN tariffs in the short run and 0.3 percentage point reduction in the long run. These estimates imply that the CUSFTA tariff preferences generated a decrease in the MFN tariff rate of 2.1% for the average Canadian industry between 1989 and 1998, accounting for 55% of observed external tariff cuts during that period. This result suggests that the size of the partner country may play an important role for the effect of an FTA on MTL because an FTA will have small effect on the terms-of-trade and tariff revenue when the partner country is small.

The evidence on the effect of the CUSFTA on lobbying activity is weak and inconclusive, implying that the "rent destruction" effect was probably not among the main factors of the Canadian trade policy. This result echoes with Ketterer, Bernhofen, and Milner (2012) who also found no effect of political economy factors on Canadian MFN tariff reduction in 1990s.[1] Yet, we believe that our data allows a more reliable analysis of the effect of the CUSFTA on lobbying activity. We move away from the assumption that all industries are equally involved in lobbying for protection and use Canadian Lobbyists Registry data to identify the number of lobbyists representing each industry. With this data, we construct several alternative measures of industrial lobbying activity and run a direct test of the Grossman and Helpman (1995) model. In most specifications we failed to find any evidence for deeper tariff reductions in politically active industries.

Another major finding of this paper is the presence of trade policy cooperation in the CUSFTA. We show that the Canadian government is more reluctant to liberalize industries where it would imply greater preferential rent erosion for the US exporters. Although this result is consistent with the findings by Limao (2006) and Karacaovali and Limao (2008) for the US and the EU, it is in contrast with Ketterer, Bernhofen, and Milner (2012), who document deeper Canadian tariff reductions in industries with the US presence. In these studies the identification of the "stumbling block" effect relies on whether products are imported from the partner country or not. For FTAs with small partners there is likely to be enough cross-industry variation in export status to identify the effect of interest. For the CUSFTA, however, more than 95% of all products are exported by US to Canada, making the identification of the stumbling block effect to rely on a very small number of industries. Based on our model with unilateral trade liberalization, we propose a different identification strategy which relies on cross-industry variation in the preference rent collected by the partner country. We categorize all industries according to the strength of US trade interests in the Canadian market by using industries' shares in total US exports to Canada. If cooperative motives in trade policy exist, Canadian MFN tariffs would fall less in industries that are more important for US exporters to Canada. We show that during 1989-1998 time period, industries with the smallest export shares showed an additional 1.9 percentage point reduction in the MFN tariff relative to industries with the largest shares. These findings are robust to various model specifications and endogeneity problems caused by reverse causality.

[1] Karacaovali and Limao (2008) confirmed the presence of "rent destruction effect" in the EU trade policy during the Uruguay round of the WTO tariff reductions.

Comparing the magnitude of tariff complementarity and substitutability factors, we find that the former dominate the latter so that even industries that contribute the most to US exports into Canada observed net tariff reductions caused by the CUSFTA tariff preferences. This result is in contrast to Limao (2006) and Karacaovali and Limao (2008) who found the net stumbling block effect of trade agreements formed by the US and the EU. The likely reason for this difference in the results is the difference in objectives for countries to enter an agreement. The US and the EU mostly form trade agreements with small countries which are not capable of generating large trade gains for them. However, these agreements often come with provisions which require small countries' cooperation on non-trade issues, and substantial preferential margins may be necessary to provide enough incentives for cooperation. The main purpose of the CUSFTA, on the other hand, was to exchange market access between two large economies rather than to stimulate compliance with non-trade provisions by a prospective FTA partner country. That said, the evidence of trade policy cooperation between the CUSFTA countries suggest that there may exist other reasons for an FTA members to internalize the effect of their trade policies on each other.

## 2 The Theory of FTA Trade Policy

In this section we review the main channels identified in the previous literature through which an FTA may affect external tariffs of its member countries. We present a simple model of monopolistic competition with differentiated products and restricted market entry[2] to illustrate those channels and to derive the equilibrium trade policy of an FTA under different theoretical assumptions. Predictions of this model will lay foundations for our empirical specifications which we use to estimate the effect of the CUSFTA on the Canadian trade policy.

Consider an economy with three countries indexed by $H$, $P$, and $F$, denoting home, FTA partner, and the rest of the world, respectively. All countries produce and trade $N+1$ goods, with the first good being a numeraire, traded at no costs and produced by perfectly competitive

[2]All key predictions of this model hold under alternative market structures as long as the terms-of-trade motive for the trade policy is present.

firms. This assumption fixes wages at the price of the numeraire good, normalized to 1. For all other industries $i$ the number of firms in each country $j$ is fixed and equals to $k_{ij}$, and each firm produces a distinct variety of the good. The representative consumer at Home is characterized by the following quasi-linear utility function:

$$U(X_0, X_i) = X_0 + \sum_{i=1}^{N} a_i \ln X_i, \quad \sum_{i=1}^{N} a_i = 1 \tag{1}$$

where $X_0$ is consumption of the numeraire good. $X_i$ is the constant elasticity of substitution sub-utility derived from consumption of good $X_i$:

$$X_i = \left( \sum_{j=H,P,F} \sum_{f=1}^{k_{ij}} d_{ijf}^{\frac{1}{\sigma_i}} c_{ijf}^{\frac{\sigma_i - 1}{\sigma_i}} \right)^{\frac{\sigma_i}{\sigma_i - 1}} \tag{2}$$

where $\sigma_i$ is the elasticity of substitution between varieties of product $i$ and $d_{ijf}$ denotes the preference or quality parameter for good $i$ produced by firm $f$ in country $j$. Suppose that production costs in country $j$ and industry $i$ are constant and equal to $w_{ij}$ and $d_{ijf} = d_{ij}$ for all $i$ and $j$.[3] Then the profit-maximizing pricing strategy that firm $f$ in industry $i$ sets in the Home country market is

$$p_{ijf} = \left( \frac{\sigma_i}{\sigma_i - 1} \right) (w_{ij} + \tau_{ij}) \tag{3}$$

where $\tau_{ij}$ is the specific tariff imposed by the home country government on imports of good $i$ from country $j$ with $\tau_{iH} = 0$. National welfare $W$, defined as the indirect utility of the representative consumer, is the sum of consumer surplus from consumption of differentiated goods (CS), tariff revenue (TR), and profits of domestic firms ($\pi_H$):[4]

$$W_0(\tau) = CS(\tau) + TR(\tau) + \pi_H(\tau) \tag{4}$$

$$CS(\tau) = U(X_0, X_i, \tau) - \textstyle\sum_{j,i,f} p_{ijf} c_{ijf}$$

$$TR(\tau) = \textstyle\sum_{j=P,F} \sum_{i,f} \tau_{ijf} c_{ijf}$$

$$\pi_H(\tau) = \textstyle\sum_{i,f} \pi_{i,H,f}(\tau_i) = \sum_{i,f} \frac{p_{i,H,f} x_{i,H,f}(\tau_i)}{\sigma_i}$$

where $\tau$ is $2N \times 1$ vector of endogenously determined import tariffs set by the home country government according to some objective function. A common problem in the theoretical literature is that this objective function is unknown and what one assumes about the government's

[3] This assumption implies that all firms within each country and industry are symmetric in terms of costs structure and consumer's demand.

[4] Labor income is normalized to one and is omitted from the expression for welfare for simplicity.

preferences may have important implications for the equilibrium trade policy. In what follows, we consider several specifications of the government's objective function most commonly used in the literature and then use the empirical analysis to differentiate among the alternative specifications.

## 2.1 Non-cooperative trade policy of an FTA

A large body of literature on FTAs with endogenous trade policy assumes that governments of FTA member countries have no political economy motivations and set import tariffs non-cooperatively. With the government's objective function being equal to national welfare, $G_0(\tau) = W_0(\tau)$, the resulting equilibrium ad-valorem import tariff $t_{it}^F$ for imports of product $i$ in year $t$ from country $F$ will maximize national welfare (4) and will take the form[5]

$$\epsilon_i t_{it}^F = (\sigma_i - 1) s_{it}^P t_{it}^P + \frac{\sigma_i - 1}{\sigma_i} s_{it}^H \tag{5}$$

where $\epsilon_i$ is the price elasticity of import demand at Home, $t_{it}^P$ is the preferential ad-valorem tariff rate on imports from the partner country, and $s_{it}^j$ is the share of the Home country's market supplied by firms from country $j = H, P$. The first term on the right-hand side shows that the FTA's external and internal tariffs are positively related. This result was first obtained by Richardson (1993) and later has come to be known as the "tariff complementarity" effect due to Bagwell and Staiger (1997b). Intuitively, a reduction in the tariff rate towards the FTA partner country reduces imports from the ROW, thus reducing tariff revenue proportional to the market share of the partner country firms $s_i^P$. Furthermore, the tariff revenue effect is stronger if the partner country and ROW exports are close substitutes. The second term on the right-hand side reflects government's incentives to protect imperfectly competitive industries. It stems from government's incentives to use trade policy in order to shift consumer expenditure from foreign to domestic producers because only profits of the latter enters the expression for national welfare and the government's objective function. Since the size of the market share of domestic firms affects the share of consumers' expenditure redirected to domestic producers, the strength of reallocating effect of an import tariff is proportional to $s_{it}^H$. Moreover, the ability

[5] See Stoyanov (2009) for complete derivation.

of trade policy to redistribute expenditure from foreign to domestic varieties is stronger when these varieties are close substitutes thus this effect is proportional to $\left(\frac{\sigma_i - 1}{\sigma_i}\right)$.

## 2.2 Cooperative trade policy of an FTA

The literature based on cooperative tariff formation has different predictions about the effect of an RTA on external tariffs. Using different theoretical frameworks, Kennan and Riezman (1990), Bagwell and Staiger (1997a), Bagwell and Staiger (1998), and Ornelas (2007) show that members of customs unions (CU), which set the common external tariff (CET) cooperatively in order to maximize the joint welfare of the union, may want to increase the MFN tariff relative to pre-CU level. There are two main contributing factors to protectionist trade policy of a CU. First, there is a terms of trade argument since the increased economic size of the trading block increases its market power and thereby allows countries in the union to redistribute surplus from their trading partners to themselves. Second, CUs tend to have higher tariffs because their members take into account the effect of CET on each other's welfare. The cooperative trade policy of a CU internalizes the positive effect of a CET on export rents within the block, making CUs more protective than FTAs.

While the first factor simply reflects increasing market power of the trading bloc, the second factor illustrates the role of cooperation on trade policy issues. When members of the trading block coordinate their trade policies, as they do in CUs, the resulting trade policy becomes more protectionist as the member countries internalize the externalities created by their trade policies on each other's firms. Of course, one can argue that members of an FTA may not have enough incentive to cooperate on their trade policies, whereas countries in a CU are forced to cooperatively choose their CET. However, several recent studies suggest that this may be so. Limao (2007) built a theoretical model of an RTA with a public good supplied by individual countries which has regional spillovers. In his model, preferential tariffs can be used indirectly to forge cooperation on non-trade issues between RTA partners and to address the problem of underprovision of athe public good with cross-border spillover effects. Accordingly, RTA members want to maintain the preference margin by keeping the MFN tariff high in order to stimulate economic and political involvement of their partners in non-trade issues, internalizing the decision on the provision of a regional public good.

As in Limao (2007), suppose that in addition to consumption of the numeraire and differentiated goods, consumers also value a public good, provided by both home and partner country governments, which has a positive regional spillover effect. Denote by $e_j$ the expenditure on public good in country $j = H, P$, and by $\Psi(e_H, e_P)$ the additively separable subutility from consuming the public good. Then national welfare and the government's objective functions become:

$$G_1(\tau) = W_1(\tau) = CS(\tau) + TR(\tau) + \pi_H(\tau) + \Psi(e_H, e_P) - e_H \tag{6}$$

Under the assumption that the partner country does not value the public good and sets $e_P = 0$ in the absence of trade preferences, as in Limao (2006) and Karacaovali and Limao (2008), the equilibrium import tariff without FTA will be the same as the non-cooperative tariff in equation (5). In the presence of an FTA, however, the home country's government will take into account the effect of the MFN tariff on the preference margin and hence on the market access rent collected by the FTA partner, which can be used to stimulate expenditure on the public. Therefore, the first-order condition and the optimal import tariff under the FTA becomes:

$$\frac{\partial G(\tau)}{\partial \tau_i} = \frac{\partial W_0(\tau)}{\partial \tau_i} + \frac{\partial \Psi(e_H, e_P)}{\partial e_P}\frac{\partial e_P}{\partial \tau_i} = 0$$
$$\epsilon_i t_{it}^F = SB_i + (\sigma_i - 1)s_{it}^P t_{it}^P + \frac{\sigma_i - 1}{\sigma_i} s_{it}^H \tag{7}$$

where $SB_i = \frac{\sigma_i - 1}{\sigma_i}\frac{1}{X_i P_i'}\frac{\partial \Psi(e_H,e_P)}{\partial e_P}\frac{\partial e_P}{\partial \tau_i}$ and $P_i' > 0$ is the derivative of the CES price index with respect to $\tau_i$. From equations (5) and (7), the effect of an FTA on the home country's MFN tariff becomes:

$$\epsilon_i \Delta t_{it}^F = SB_i + (\sigma_i - 1)\Delta\left(s_{it}^P t_{it}^P\right) + \frac{\sigma_i - 1}{\sigma_i}\Delta s_{it}^H \tag{8}$$

As long as FTA member countries cooperate their policies in the sense that the partner country responds to additional tariff preferences with an increased supply of the regional public good $\left(\frac{\partial e_P}{\partial \tau_i} > 0\right)$, there is an additional stumbling block effect of an FTA ($SB_i > 0$) coming from the home country's government incentives to maintain a large enough preference margin for the partner by increasing the MFN tariff subsequent to FTA formation or not decreasing it too much.[6] The same result is obtained in the absence of the public good if FTA partner countries set trade policies cooperatively, as in the case of a CU, and take into account the effect of their MFN tariffs on each others' welfare: $W_1(\tau) = W_0(\tau) + \gamma W_0^P(\tau)$.

### 2.3 Trade policy of an FTA under political economy

Our third empirical specification follows from a political economy model of trade policy proposed by Grossman and Helpman (1994). This model departs from welfare maximizing trade policy and assumes instead that governments choose the level of tariffs in order to maximize the weighted sum of national welfare $W$ and political contributions $C$ provided by domestic special interest groups:

$$G_2(\tau) = aW_1(\tau) + C \tag{9}$$

where $W_1(\tau)$ is national welfare as defined in the previous section and $a > 0$ is the weight that government attaches to one dollar of welfare relative to one dollar of contributions. As in Grossman and Helpman (1994), we assume that some domestic industries are politically organized and provide the government with contribution schedules, which are contingent on its choice of trade policy, while others are not and do not participate in the tariff-setting process. Let $I_i$ be an indicator variable which takes the value of one if industry $i$ is organized or zero otherwise. Then the objective of the home country government is to choose $\tau$ which maximizes

$$G_2(\tau) = aW_1(\tau) + \sum_i I_i C_i(\tau) \tag{10}$$

Grossman and Helpman (1994) show that with truthful contribution schedules, the optimum trade policy choice maximizes the preference-weighted sum of national welfare and welfare of organized interest groups, which includes profits, consumer surplus and their share in redistributed tariff revenue. In the presence of a preferential trade regime between countries $H$ and $P$,

[6] It is important to note that the expression for the optimal import tariff is slightly different from that in Limao (2007). First, Limao assumed that in the absence of an FTA the MFN tariff will be set cooperatively with the ROW, which is not relevant in the context of the CUSFTA. Accordingly, $SB_i$ term in our model does not include the opportunity costs of missed tariff revenue due to partner country's free riding on liberalization efforts of $H$ and $F$. Second, Limao assumed that the FTA partner country is small so the second term in equation (7) is not present in his model.

the equilibrium tariff imposed on imports from country $F$ takes the following form:

$$\epsilon_i t_{it}^F = \frac{a}{a+\alpha} SB_i + (\sigma_i - 1) s_{it}^P t_{it}^P + \frac{a}{a+\alpha} \frac{\sigma_i - 1}{\sigma_i} s_{it}^H + \frac{1}{a+\alpha} \frac{\sigma_i - 1}{\sigma_i} I_i s_{it}^H \tag{11}$$

where $\alpha$ is the share of population represented by one of the lobbying groups. The last term measures the effect of domestic lobbying on the MFN tariff. The positive coefficient implies that MFN tariffs are higher in industries with the presence of domestic lobbying. As with the third term, the redistributive power of import tariffs depends on the share of domestic firms in the market and on the degree of product differentiation. Equation (11) states that there is an additional channel for the effect of an FTA on the MFN tariff. Since the strength of domestic lobbying is proportional to $s_{it}^H$, an FTA and the following reduction in the market share of domestic firms will weaken lobbying power of the industry and reduce its lobbying intensity for protection. This additional pro-liberalization effect of trade agreements was first identified by Ornelas (2005) who demonstrated that FTAs erode protection rent enjoyed by home country firms.

# 3 Empirical implementation

## 3.1 Econometric specifications

Equations (5), (8), and (11) summarize three main channels for the effect of an FTA on trade policies of member countries and they motivate our main empirical specifications. After the introduction of an error term to the most parsimonious model which excludes cross-border externalities and political economy factors, the empirical specification based on model (5) becomes:[7]

$$Y_{it} = \alpha + \phi_1 X_{it}^1 + \phi_2 X_{it}^2 + \gamma_i + \gamma_t + u_{it} \tag{12}$$

$$Y_{it} = \frac{\sigma_i - 1}{\sigma_i} \epsilon_i t_{it}^F, \; X_{it}^1 = s_{it}^P t_{it}^P, \; X_{it}^2 = s_{it}^H$$

[7] We exclude $\sigma_i$ from $X_{1t}$ in the estimation equation to reduce the chance of having a measurement error.

where $\gamma_i$ and $\gamma_t$ and industry and year fixed effects which capture the effect of time- and industry-invariant factors that are not present in the theoretical model but may influence trade policy formation. Given that the model is static and does not inform us about the dynamic response of the MFN tariffs to changes in the right-hand side variables, we apply two alternative time-difference operators in order to identify coefficients $\phi_1$ and $\phi_2$:

$$\Delta Y_{it} = \alpha + \phi_1 \Delta X^1_{it-1} + \phi_2 \Delta X^2_{it-1} + \gamma_t + u_{it} \tag{13}$$

$$\Delta_9 Y_{i,98} = \alpha + \phi_1 \Delta_9 X^1_{i,1998} + \phi_2 \Delta_9 X^2_{i,1998} + u_i \tag{14}$$

Equation (13) is in first differences and measures contemporaneous relationship between MFN and preferential tariffs where all explanatory variables are lagged by one year to allow for a small delay in response in the dependant variable. Although CUSFTA tariff cuts took place between 1989 and 1998 while most of the MFN tariff reductions negotiated in the Uruguay Round of the WTO happened after 1995, there is still enough variation in both variables prior to 1995 to identify whether or not the presence of a short-term response in MFN tariffs to CUSFTA trade liberalization exists.[8] Yet, Canada does not change its import tariffs frequently and it may take more than one year for the MFN tariff to react to changes in the market environment. Therefore, our second empirical specification is equation (12) differenced over the entire time period of the CUSFTA trade liberalization in order to estimate the long-term response of the MFN tariffs to variation in the right-hand side variables. While model (13) can provide important information on short-term adjustments in trade policy to preferential liberalization, the more general long-run model (14) will be used to gauge the overall effect of the CUSFTA on Canadian multilateral tariff changes during the Uruguay Round of the WTO tariff reductions.

Estevadeordal, Freund, and Ornelas (2008), henceforth EFO, use the intuition behind equilibrium import tariff (5) to test the reduced form relationship between external and internal

[8] More than half of the variation in MFN tariffs during 1990-1994 time period occurs within industry, comparing to three quarters for the time period 1995-1998. These fractions are very similar when calculated using 10-digit HS industry classification at which commodity tariffs are defined. Although the overall variation in MFN tariffs after 1995 is four times greater than before, we believe it is enough to identify $\phi_1$ and $\phi_2$ on both subsamples. In the Section 3.5 we report results estimated from the two subsamples separately.

tariffs for a group of South American countries and confirmed that tariff preferences within RTAs are inversely related to changes in MFN tariff rates. Before presenting results for the structural estimation, we will start with the empirical specification similar to the one suggested by EFO and examine the response of Canadian MFN tariffs to preferential tariff cuts on U.S. imports:

$$\Delta t_{it}^{F} = \alpha + \phi_0 \Delta t_{i,t-1}^{P} + \gamma_t + u_{it} \tag{15}$$

$$\Delta_9 t_{i,1998}^{F} = \alpha + \phi_0 \Delta_9 t_{i,1998}^{P} + u_i \tag{16}$$

Since tariff preferences to a partner country reduce the socially optimal external tariff both through the tariff revenue effect ($X_{it}^1$) and through the the market structure effect ($X_{it}^2$), we expect $\phi_0$ to be positive. Positive $\phi_0$ would imply that preferential tariff cuts are followed by MFN tariff cuts and support the tariff complementarity hypothesis.

To test the hypothesis trade policy cooperation between Canada and the US, we employ the empirical methodology similar to Limao (2006) to estimate equation (8). Since the $SB_i$ term cannot be constructed from the data, it is treated as the parameter of the empirical model, $\beta_0$, which is constant across industries and is expected to contribute positively to the MFN tariff whenever good $i$ is imported from the FTA partner country:

$$Y_{it} = \alpha + \beta_0 D_{it} + \phi_1 X_{it-1}^1 + \phi_2 X_{it-1}^2 + \gamma_t + u_{it} \tag{17}$$

where $D_{it}$ is a dummy variable taking the value of one for goods imported from the US.

However, our preferred model for estimating stumbling block effect is different from Limao (2006) for two reasons. First, Canadian imports from the US are positive for nearly 99% of all 6-digit industries and identification of coefficient $\beta_0$ relies on too few observations. Second, we would expect sectors with greater U.S. involvement to observe less trade liberalization since the stumbling block effect should get stronger for industries that are more important for the partner's exports. The reason is that in such industries a given preference margin applies to a larger volume of exports and generates more rent for a partner country, inducing greater expenditure on the public good, i.e. $\frac{\partial e_P}{\partial \tau_i}$ and $SB_i$ are both increasing in the partner country's exports to Home. Therefore, we differentiate industries according to their importance for the U.S. exports to Canada, and estimate the relationship between $SB_i$ and trade liberalization along the distribution of the share of each industry in total US exports to Canada. In particular, we use quintiles of the distribution of the US export share to Canada to categorize industries into five groups. Denoting by $D_{it}^k$ a dummy variable which takes the value of one if industry $i$ falls into $k$-th quintile, the empirical specification becomes as follows:[9]

$$\Delta Y_{it} = \alpha + \sum_{k=1}^{4} \beta_k D_{it}^k + \phi_1(\sigma_i - 1)\Delta\left(s_{it}^P t_{it}^P\right) + \phi_2 \frac{\sigma_i - 1}{\sigma_i} \Delta s_{it}^H + \gamma_t + u_{it} \tag{18}$$

Similarly to equations (13) and (14), the model (18) is estimated using short and long time differencing. If tariff cooperation exist in the CUSFTA, industries that have higher U.S. representation would have received more protection against foreign competition in the Canadian market. Thus, we expect $\beta_i$ to be negative which would imply that industries that contribute the most to the US exports to Canada are the least liberalized ones (the omitted category is industries with the largest US exports to Canada). Moreover, if industries with larger export shares tend to be more protected, we would expect to find the following ranking of $\beta_k$ coefficients:

$$\beta_{k-1} < \beta_k < 0, \forall k = 2, 3, 4 \tag{19}$$

It is important to note that while the empirical specification (18) is different from Limao (2006) and Karacaovali and Limao (2008), it is a valid test of the stumbling block hypothesis in the context of our theoretical model. In Limao (2007) model Home country and the rest of the world negotiate multilateral tariff cooperatively in order to maximize joint welfare. Under this assumption, the sign of the stumbling block effect is positive only for the corner case when the preferential tariff is zero and can thus be empirically estimated only for a subset of industries with free trade between FTA partners. In the absence of multilateral tariff negotiations, as in the case of our model, the stumbling block effect is always positive for industries with positive imports from the partner country. However, since the US export to Canada is positive for nearly all 6-digit HS product categories, we use equation (18) to identify the cross-industry variation in the strength of the stumbling block effect which may vary with the partner country's gain

[9] We also experimented with the interactions of $D_{it}^k$ and preferential tariff changes to check if the effect of the terms-of-trade channel varies across industries with different US exports. Since these interactions are insignificant in all of our specifications, we do not report the results in the paper but they are available upon request.

from tariff preferences. The benefit of receiving a preference for a parner country is equal to $\left(t_{it}^{F}-t_{it}^{P}\right)p_{it}^{P}c_{it}^{P}$ and is captured by the term $\partial e_P/\partial\tau_i$ in equation 7. Given that the benefit is increasing in the US revenue from exporting to Canada, $p_{it}^{P}c_{it}^{P}$, we would expect $\frac{\partial^2 e_P}{\partial\tau_i\partial\left(p_{it}^{P}c_{it}^{P}\right)}>0$, i.e. that the same tariff preference will generate more surplus to the partner country, and hence more expenditure on public good, in industries that generate more revenue to the US exporters.

Finally, to arrive at our most complete empirical specification with political economy factors, we rearrange equation (11) by adding fixed effects and time differencing it:

$$\Delta Y_{it}=\alpha+\beta_0 D_{it}+\phi_1\Delta X_{it-1}^{1}+\phi_2\Delta X_{it-1}^{2}+\phi_3\Delta X_{it-1}^{3}+\gamma_t+u_{it} \tag{20}$$

$$\Delta_9 Y_{i,98}=\alpha+\beta_0 D_i+\phi_1\Delta_9 X_{i,1998}^{1}+\phi_2\Delta_9 X_{i,1998}^{2}+\phi_3\Delta_9 X_{i,1998}^{3}+u_i \tag{21}$$

where $X_{it}^{3}=I_i s_{it}^{H}$. Positive coefficient $\phi_3$ would imply that while politically organized industries tend to receive more protection from policymakers in general, a reduction in the domestic market share, triggered by the partner country's preferential market access, would cause deeper tariff cuts in those industries.[10] The reason for those deeper cuts is that an FTA leads to a reduction in protectionist rent retained by domestic firms since a part of it will be netted by the partner country's firms. This "rent destruction" effect, originally identified by Ornelas (2005), weakens the incentives of domestic firms to lobby for protection and results in lower levels of external tariffs by moderating political economy distortions.

## 3.2 Estimation issues

Since we are interested in establishing a causal effect of the CUSFTA on Canadian multilateral trade liberalization, it is important to discuss endogeneity concerns with preferential trade liberalization measures and the ways of dealing with them. The primary concern is the potential endogeneity of preferential tariff cuts. The CUSFTA came into force on January 1, 1989, and resulted in elimination of nearly all tariffs by 1998. Trade liberalization between the two countries followed tariff reduction schedules, which were adopted in 1986-1987 when the

[10] Most of the previous empirical literature, e.g. Karacaovali and Limao (2008) and Ketterer, Bernhofen, and Milner (2012), assumed that all industries are equally active in lobbying. Under this assumption, $X_{it}^{1}=X_{it}^{3}$, making it impossible to identify separately the market structure and political economy effects.

CUSFTA was negotiated.[11] Given that most MFN tariff cuts took place after 1995, preferential tariff reductions can be viewed as predetermined relative to subsequent MFN tariffs and the reverse causation from multilateral to preferential trade policy is unlikely. While there can be other factors affecting both preferential and MFN tariffs, such as industry-specific variation in economic and political conditions, to the extent that the Canadian government committed itself to removing tariffs on US imports entirely, preferential tariff cuts seem to be *a priori* exogenous to variation in MFN tariffs.

However, there is one caveat that should be kept in mind when preferential tariff cuts are viewed as exogenous. The fact that tariffs were completely eliminated by 1998 implies that in specifications with changes over the entire CUSFTA phase-out period $\Delta t^P$ will be highly collinear with the initial MFN tariff rate and may thus capture the ease of tariff cut implementation. This should be less of a problem in structural specifications where the interaction of $\Delta t^P$ with the US market share captures the economic value of tariff complementarity effect for tariff revenue. Therefore, as a robustness test we also run specifications where $\Delta t^P$ and $\Delta X^1$ enter separately to isolate the effect of the FTA on the government's economic incentives to change external tariffs from the effect of initial tariff rate on flexibility of trade policy adjustment. In general, however, the causal interpretation of our results should be treated with caution in the absence of good instruments for preferential tariff changes.

The indicator variables for the US presence in the Canadian market can also be endogenous due to reverse causation since the decision to export to Canada and the share of industry in total US exports may depend on the preference margin. To deal with endogeneity of FTA partner's export dummy variable we follow Limao (2006) and use the instrumental variable approach. The first instrument for $D_{it}$ is the dummy variable which takes the value of one for products exported by the US to Canada in 1988, the last year before the first round or the CUSFTA tariff cuts which makes this instrument independent of tariff preferences. The second instrument is the change in the world price for product $i$, measured as the absolute change in price in the previous year for the short-run specifications and as the change between

[11] Tariff reduction schedules classified all products into three groups. Tariffs on products in the first group were eliminated entirely in the first year of the agreement. Tariffs for the other two groups were eliminated in equal annual stages over five and ten years, respectively.

1989 and 1994 for the long-run specifications. While correlated with incentives to export, world price changes occurring prior to the decision to adjust MFN tariffs are likely to be exogenous. Using the same logic, we use quintile dummies for product $i$ in year 1988 and their interactions with prices changes defined above to instrument $D_{it}^k$ variables. Quintile dummies constructed for 1988 represent valid instruments for $D_{it}^k$ because ranking of industries' in US exports to Canada prior to CUSFTA formation is independent of subsequent MFN tariff changes and is highly correlated over time. When interacted with price changes, these variables capture transitions across quintiles of the US export share distribution over time due to exogenous changes in the world prices.

Another challenge with estimation of equations (20) and (21) is the endogeneity problem arising from the simultaneity of market shares and the MFN tariff rate. We address this problem by using a number of different instrumental variables for $s_{it}^H$ and $s_{it}^P$ suggested by previous literature. For the Canadian market share, the list of instruments includes factor shares of physical capital, non-production labour, intermediate inputs, and fuel and electricity in industry output using 6-digit NAICS industry classification. Trefler (1993) suggests that industry factor endowments are independent of the level of protection and thus provide exogenous variation in the Canadian market share. As an additional instrument, we use the revealed comparative advantage index proposed by Balassa (1965).[12] An increase in the revealed comparative advantage index would imply an increase in the competitiveness of Canadian firms, and one would expect to see an increase in the share of domestic and decrease in the share of foreign firms in the Canadian market. At the same time, we found no evidence that Canadian tariff preferences for the US products are related in any systematic way to the growth rate of Canadian exports to other countries thus there are no reasons to believe that the revealed comparative advantage is affected by the Canadian MFN tariff.[13] Similarly, the index of the US revealed comparative

[12] The revealed comparative advantage index is constructed at the product level as $RCA_{it} = \frac{X_{it}/\Sigma_j X_{jt}}{Z_{it}/\Sigma_j X_{jt}}$, where $X_{it}$ is Canadian exports of good $i$ in year $t$ to all countries other than US, and $Z_{it}$ is the corresponding level of exports by all other countries to all destinations excluding the US. The US market is excluded from the calculation since US tariff preferences for Canada, determined simultaneously with Canadian preferences for the US, could have changed the structure of the Canadian exports. In the empirical analysis we use a symmetric index of revealed comparative advantage: $RSCA_{it} = \frac{RCA_{it}-1}{RCA_{it}+1} \in [-1;1]$.

[13] Bown and Crowley (2007 JIE), however, document positive effect of US antidumping duties against China

advantage in the world market excluding Canada is used to instrument the share of US firms in the Canadian market.

To address the issue of measurement error in the political organization dummy variables, constructed from an indirect measure of lobbying activity and discussed in details in the next section, we follow the general approach in the political economy literature by instrumenting them with the market concentration ratio and with the log of average scale. Equation (20) and (21) are estimated by 2-step GMM and all instruments which do not pass the orthogonality to the structural error test at 95% confidence level are excluded from the first stage regression. Since both equations are non-linear in endogenous variables, we also include the cross product of instruments for market shares and political organization dummy variables in the list of instruments.[14] Similarly, to instrument $X_{it}^1$ we use the cross-products of instruments for $s_{it}^P$ and the preferential tariff changes, treating the latter as exogenous. All empirical specifications include Canadian tariff rate in 1988 as an additional regressor to control for the cross-industry variation in the scope of the MFN tariff reductions.

### 3.3 Data

The data used for this paper come from several different sources and cover the time period from 1989 to 1998, the entire phase-out period of import tariffs under the CUSFTA. While trade data is available at 6-digit HS product classification, all industry-level data is only available at 6-digit NAICS. We keep the data at 6-digit HS classification and whenever data are only available at a higher level of aggregation, it is replicated for all 6-digit HS codes within the corresponding aggregate industry.[15] Canadian import and tariff data are obtained from Statistics Canada at HS-6 level. Import tariffs are constructed as a ratio of import duties over the

on Chinese exports to other countries. Therefore, in our empirical analysis we pay close attention to the validity of the exclusion restriction tests.

[14]Wooldridge (2001) shows that cross-products of two sets of exogenous variables are the most relevant instruments when dealing with the product of two endogenous variables.

[15]For this reason, in all regressions where industry-level data is used the standard errors are clustered at the 6-digit NAICS level.

value of imports.[16] The data on output, capital, employment, intermediate inputs, and fuel and electricity consumption is also provided by Statistics Canada. It is recorded at 6-digit NAICS, and we use concordance provided by Industry Canada to make it compatible with the 6-digit HS classification. The home country's market shares were constructed at 6-digit NAICS level as the value of industry shipments (net of exports) relative to total consumption (total shipments minus net exports). The US market share is constructed similarly as the ratio of Canadian imports from the US relative to domestic consumption. The data on Canadian, US, and the ROW's exports, used in the construction of revealed comparative advantage indices, come from the World Bank's World Integrated Trade Solution (WITS) database, and is recorded at 6-digit HS classification. Elasticities of substitution for Canada, $\sigma_i$, were obtained from Broda, Greenfield, and Weinstein (2006) at 3-digit HS industry classification. Import demand elasticities were obtained from Kee, Nicita, and Olarreaga (2009) at 6-digit HS level.

Table 1 provides the summary statistics for the key variables in this study. The average MFN tariff is 5.7% and the average preferential tariff is 2.5% during the phase out period, suggesting that the average preferential margin is equal to 3.2%. The average annual reduction in the MFN tariff is 0.4%, which is 0.3 percentage points less than the average reduction in the preferential tariff. The mean value for the Canadian home market share decreased by approximately 1% annually, from 63% in 1990 to 53% in 1998. During the same period, the US market share in Canada increased by 1% annually, from 21% in 1990 to 29% in 1998.

To construct political organization dummy variables we use data from Stoyanov (2009) and then apply different approaches to categorize industries into politically organized and unorganized ones. The data include information on the number of lobbyists officially registered with the Canadian Registrar of Lobbyists and the firms which recruited them. Each firm is assigned to one 6-digit NAICS industry based on its primary manufacturing activity. We then calculate the total number of lobbyists representing interests of each 6-digit NAICS industry. Since the theory is not very informative about how to classify industries into politically active or not, we construct four different measures of industrial political organization to analyze the sensitivity

[16] Since this is not a perfect measure of import tariffs, we exclude 1% of observations with the highest MFN and preferential tariffs from the data to minimize the risk of measurement error.

of estimation results to the formulation of this generated variable. In our first two measures, we classify an industry as politically active ($I_i = 1$) if it is represented by at least one and at last three lobbyists, respectively. The summary statistics for these two dummy variables, $I_{1i}$ and $I_{2i}$, is presented in Table 1.

To build our third measure of political activity, we follow Gawande and Bandyopadhyay (2000) and regress the number of lobbyists in an industry on the import penetration ratio interacted with 3-digit NAICS dummies and a constant term. All industries with positive coefficients on these interactions are defined as politically active. The intuition behind this definition of political organization is that industries threatened more by import competition will seek more protection from the government. We label this variable as $I_{3i}$.

The last mechanism for constructing political organization dummies follows Matschke (2008) in which the number of lobbyists is regressed on the deadweight loss of protection (normalized by the value added) interacted with 3-digit NAICS dummies. As with the previous measure, all industries with positive coefficients are assumed to be politically organized, while others are not. This specification is motivated by the theoretical prediction that industries with larger welfare losses from protection domestic interest groups should spend more resources on lobbying and recruit more lobbyists.

### 3.4 Estimation Results

In this section, we present the estimation results for the empirical models described in Section 3.1 and discuss the implications of each of them.

Table 2 presents short-run estimation results for reduced form specification (15). The positive and statistically significant estimate of $\phi_0$ coefficient in column (1) confirms the tariff complementarity hypothesis and indicates that tariff preferences granted to the U.S. are associated with reductions in the MFN tariff rate in the following year. The estimate of 0.1053 implies that every one percentage point reduction in preferential tariffs is associated with 0.1053 percentage points reduction in the MFN tariffs, which is considerably below the estimate of 0.57 obtained by EFO for the MERCOSUR. If one believes that the tariff preference schedules, negotiated in 1987-88, are pre-determined, then this relationship can be considered casual unless there are some dynamic factors which had affected the CUSFTA negotiations in 1980s and the propensity to liberalize MFN tariffs in 1990s. Applying our results for an average industry, the reduction in preferential tariff rates caused an additional annual reduction in the MFN rate by 0.08 percentage points and can explain almost 20% of the overall MFN tariff cuts between 1989 and 1998.

To test the hypothesis that Canadian MFN tariff reductions in 1990s were set cooperatively with the US, we regress annual change in the MFN tariff on the indicator variable $D_{it}$ in column (2). Canadian preferential tariffs provide US firms with a comparative advantage against firms from outside of the CUSFTA. Therefore, if the CUSFTA trade policy is set cooperatively, the Canadian government would rely upon trade policy to protect interests of US firms in Canada and we would expect to observe slower MFN tariff reductions in industries with larger share of imports from the US. The OLS estimate of the coefficient on $D_{it}$ in column (2) is insignificant, both statistically and economically, which does not support the hypothesis of cooperative trade policy. However, as it was discussed previously, the construction of this variable may result in a specification problem since for around 99% of all industry-year observations there is a positive value of Canadian imports from the US. Hence, there may not be enough variation in $D_{it}$ to identify the presence of cooperative motives in trade policy formation. Column (3) presents results with $D_{it}$ disaggregated into quintiles of industry $i$'s share in total US exports to Canada. With the fifth quintile being the omitted category, we would expect tariffs to decrease faster in industries in the first four quintiles if tariffs were set cooperatively. Indeed, all coefficients are negative but only one of them is statistically significant at 85% confidence level, providing little evidence of a smaller reduction in Canadian MFN tariffs in industries which are more important for the US exports. Finally, results of a complete specification in column (4) suggest that the size of the US industry is not related to the change in Canadian MFN tariffs, while tariff complementarity effect is still present and statistically significant.

The results with IV estimates in columns (5)-(8), which address simultaneity of MFN tariffs and import indicator variables $D_{it}$ and $D_{it}^k$, also point to the dominance of the “building block” effect of the CUSFTA. The coefficient on preferential tariff change is nearly the same as in the OLS specifications and is statistically significant, indicating that every percentage point increase in tariff preferences is associated with around 0.1 percentage points reduction in the MFN tariff in the following year and by 0.179 percentage points over three years (column 8).[17] At the same time, there is no evidence of slower MFN tariff reduction in industries which have more economic significance for the US.

The results presented so far focus on the reduced-form relationship between MFN and preferential tariffs. In Table 3 we report the regression results for specifications derived from the theoretical model of an FTA with endogenous trade policy. Column (1) illustrates the estimation results for equation (13) derived from the model with non-cooperative trade policy formation (5). The positive and statistically significant estimate of $\phi_1$ indicates that a drop in the preferential tariff is associated with a reduction in the MFN tariff. Comparing the magnitude of this effect with the one in the reduced-form specification, the two are qualitatively similar. The coefficient of 1.6817 in column (1) indicates that an average industry experiences a 0.07 percentage points reduction in the MFN tariff per year due to CUSFTA tariff preferences, which accounts for nearly 17% of observed average MFN tariff reduction over the analyzed period. The coefficient on the Canadian market share, which captures the role of industry structure in imperfectly competitive markets for trade policy, is not statistically significant in all specifications.

Estimates of the model (18) in columns (3) and (4) produce more clear evidence on the presence of trade policy cooperation between the CUSFTA member countries. Analyzing the OLS results in column (3) we find that industries with less exports from the US observe deeper MFN tariff reduction, as predicted by the model with cooperative trade policy. For instance, the coefficient $\beta_1 = -0.0139$ implies that industries in the first quintile of the US import share distribution experience additional 0.4 percentage points decrease in MFN tariff per year relative to industries in the fifth quintile.[18] Furthermore, the ranking of $\beta_k$ coefficients confirms that

[17] The Angrist-Pischke first stage F-test always rejects the null of weak instruments for all endogenous variables in all specifications at 1% confidence level. We also cannot reject the hypothesis of exogeneity of instruments, suggesting that our instruments are overall of a good quality.

[18] With the mean value for the elasticity-adjustment term $\frac{\sigma_i - 1}{\sigma_i}\epsilon_i$ being equal to 3.5, an additional reduction in the MFN tariff for the average industry in the first quintile is $\frac{0.0139}{3.5} \simeq 0.004$.

industries contributing more to the US exports receive smaller reduction in multilateral tariffs, providing further evidence for the presence of the "stumbling block" effect of the CUSFTA. This effect is found to be less significant in results from IV regressions in column (4). However, the Hausman endogeneity test fails to reject the null hypothesis of endogeneity of the explanatory variables. Moreover, the p-value for the test that $D_{it}^{k}$ variables are endogenous is 0.7 and thus we cannot reject the consistency of the OLS estimates in column (3).

The results discussed so far suggest that there is a strong contemporaneous relationship between reductions in MFN and preferential tariffs. We now turn to estimating the long-term effect of the CUSFTA on the Canadian multilateral tariffs. In Table 4 we report the estimates for equation (16) to see how total changes in MFN tariffs between 1989 and 1998 were associated with the overall reduction in preferential tariffs and the accompanying changes in market shares over the entire time period of the CUSFTA trade liberalization. The comparison of the long-run and short-run elasticities of the MFN tariff change with respect to the preferential tariff change reveals considerable differences between them. The coefficient of 0.1996 in the first column of Table 4 indicates that each percentage point reduction in the preferential tariff that took place between 1989 and 1998 is associated with 0.2 percentage points reduction in the multilateral tariff, which is nearly twice as large as the short-run elasticity. Results from the structural estimation, presented in Table 5, point to a similar conclusion: MFN tariff changes are two to four times more responsive to preferential tariff cuts in the long run than in the short run.

Another noticeable difference between the short-run and long-run results are the coefficients on the US import dummy variables which all turn positive in the OLS regressions. This seems to suggest deeper tariff cuts in industries with a larger US presence, however, these results should be treated with caution. When the relationship between the MFN and preferential tariffs is estimated using the structural model (14), the coefficients on $D_{it}^{k}$ become negative (Table 5, column 3). The likely reason is that larger tariff preference may lead to a larger increase in the US market share, making the MFN import tariff less efficient in protecting domestic producers and weakening protectionist forces. Once the effect of the US market size is controlled for through the $X_i^1$ variable in the structural model, the results of the short-run and long-run models become very similar.

In sum, there is strong evidence that Canadian MFN tariff rates feature complementarity with the CUSFTA tariff preferences. Reductions in the preferential tariff rates are always found to induce a reduction in the multilateral tariffs. The evidence on the presence of cooperative motives in trade policy is less clear though. The OLS results provide strong support for the hypothesis that Canadian multilateral tariffs decreased slower in industries which generate more revenue for the US exporters. This suggests that Canadian policymakers at least partially internalize the effect of MFN tariff choice on the US producers. The IV results are less conclusive but the hypothesis of endogeneity of $D_{it}^{k}$ is always rejected suggesting that the OLS results are consistent and efficient. Yet, even in the most complete IV specification (Table 5, column 4) industries in the first quintile of the US export share distribution are found to experience greater (but not statistically significant) reductions in the MFN tariffs than industries in the fifth quintile.

We now turn to empirically testing the final prediction of the theoretical model about the effect of an FTA on MFN tariffs in the presence of political economy factors. According to the model (equation 11), an FTA reduces the share of domestic firms in the market due to an increase in the partner country firms' presence, weakening the redistributive power of the import tariff and reducing incentives of domestic special interest groups to lobby for protection. The estimation results of the full model using IV-GMM are presented in Table 6. The first four columns report results for the short-run model (20) using four different measures of $I_i$, and the last four columns report results for the long-run model (21).

The estimation results provide no evidence for the effect of the CUSFTA on lobbying for protection: the estimates of $\beta_3$ are statistically insignificant and are not robust to the definition of the political organization. Contrary to the model's prediction, a shrinking domestic market share is not found to be associated with a declining lobbying power of special interest groups and with a deeper reduction in the level of protection granted to politically organized industries. The study by Ketterer, Bernhofen, and Milner (2012) arrives at the same conclusion, although they do not attempt to classify industries by the degree of political organization and simply assume that all industries are equally active in lobbying. Therefore, we do not find support for the "rent destruction" effect of an FTA, identified by Ornelas (2005), in the Canadian data.[19]

Turning to other estimates of equations (20) and (21), they are very similar to our previously reported findings. The estimates for $\beta_1$ are positive and statistically significant for all measures of lobbying intensity, pointing to a strong tariff complementarity effect: each percentage point reduction in preferential tariff is associated with approximately 0.05 percentage points reduction in the MFN tariff in the short run and 0.3 in the long run.[20] This difference in the elasticities of the MFN tariff with respect to preferential tariff suggests that a large fraction of the cumulative effect of a one-off preferential tariff cut on the MFN tariff rate is spread across several subsequent years. Taking the sample average of the preferential tariff change and elasticities, the results from column (5) imply a total of 2.21 percentage point reduction in the MFN tariff, which accounts for 55% of MFN tariff cuts between 1989 and 1998. The coefficient on the domestic market share $\beta_2$ is negative but statistically insignificant in the long-run specification, indicating that the MFN tariffs were not adjusted for domestic industries facing sinking market shares and that the market structure is not among the main determinants of the Canadian trade policy.

Consistent with our previous findings, the coefficient estimates on the US import share dummy variables $D_{it}^k$ also remain negative and decrease with the industry's share in the US exports to Canada.[21] This result supports the cooperative trade policy hypothesis, hinting at the reluctance of the Canadian policymakers to erode tariff preferences for products which play a more important role in the US exports. To gauge the importance of this factor for the MFN

[19] This inconsistency with the theory can result from the fact that the protection for sale model deals with the static long-run equilibrium analysis and may not be well suited to analyze the short-run changes in trade policy. Furthermore, FTAs may affect lobbying activity through channels other than "rent destruction" effect. For example, the estimates may also reflect the "surge protection" forces as in the model by Imai, Katayama, and Krishna (2009) where government provides additional protection to politically organized industries when imports surges and the share of domestic firms in the market decline.

[20] With the sample mean value for the elasticity-adjustment term $\frac{\sigma_i - 1}{\sigma_i}\epsilon_i$ being equal to 3.5, and the US market share of 15%, the elasticity of the MFN tariff with respect to preferential tariff can be calculated as $0.04 \cdot \beta_1$.

[21] In Table 6 we do not use instruments for $D_{it}^k$ since in all specifications we can never reject the hypothesis of exogeneity of those variables using the Housman test. Instrumenting $D_{it}^k$ produces less precise results which still indicate more substantial tariff reductions in industries which contribute the least to the US exports to Canada.

tariff changes, we calculate its implied effect for an average industry in each quintile of the US export share distribution in our sample. The estimates in column (5) imply that industries in the first four quintile of the US export share distribution experienced 1.92, 1.51, 0.61, and 0.37 percentage points reduction in the MFN tariff, respectively, in addition to 2.21 percentage reduction in industries in the fifth quintile. This variation in the rate of tariff reduction across industries, systematically related to the importance of those industries for the FTA partner country, provides evidence in favour of the Limao's hypothesis of cooperation in trade policy in the Canadian context.

Overall, our results reveal that the CUSFTA formation induced more open trade policy in Canada. The finding of tariff complementarity between preferential and MFN tariff rates is very strong and persistent. The implied reduction in the MFN tariff in response to a one percentage point decrease in preferential tariff is in the range of $0.05 - 0.35$ percentage points for an average industry. At the same time, we found some support for the trade policies of the CUSFTA member countries to be formed cooperatively as Canada provides more protection to industries with stronger US interests. Finally, our findings suggest no effect of the CUSFTA on the intensity of industry lobbying for protection.

### 3.5 Robustness tests and extensions

In this section we perform several robustness exercises. The first two columns of Table 7 report estimation results for the short-run and long-run models with 2-digit HS industry fixed effects to control for unobserved industry-specific trends which may be related to the pace of trade liberalization. We still obtain positive and highly significant $\phi_1$ coefficient and the expected ranking of $\beta_k$ coefficients, confirming all of our previous findings.

As another robustness test we focus on industries with positive pre-CUSFTA MFN tariffs rates. Given that industries with zero initial tariffs cannot respond to changes in preferential trade, they do not contribute to the identification of the coefficients of our interest. In columns (3) and (4) we drop industries with the MFN tariff rate in 1988 lower than 1% and it does not affect the results. In columns (5) and (6) we estimate the two models on the subsample of industries for which the CUSFTA tariff reductions were scheduled over the entire ten-year phase-out period. Being the most sensitive product categories, they are also more likely to have higher initial tariffs and thus have more room for MFN tariff cuts. The results show that the elasticity of the MFN tariff with respect to preferential tariff is the same as in the benchmark specification,[22] but the effect of the size of US exports becomes more pronounced. For this subsample, industries with the least imports from the US experience additional 3.6 percentage points decline in the MFN tariff relative to industries with the largest US imports.

In columns (7) and (8) of Table 7 we depart from the strict structure of the theoretical model and remove elasticity terms from the construction of the right-hand side variable. Although there are strong reasons to believe that import demand and substitution elasticities may play an important role in trade policy by making it more or less distortive, the elasticities are also likely to be measured with error. Moreover, Gawande and Bandyopadhyay (2000) found that the elasticity terms bear little explanatory power for import tariffs in the protection for sale model. To make sure that our results are not driven by imprecisely measured elasticities, we use the MFN tariff rate as the dependent variable in these two specifications. With this modification the estimates are qualitatively similar to those obtained previously, pointing to the importance of the tariff complementarity effect and industry's share in the US exports. Columns (7) and (8) also reveal considerable difference in the estimate of $\phi_3$ coefficient which is now positive and statistically significant. This result appears persistently in many other specifications when elasticity terms are excluded from the analysis, hinting at the additional trade liberalization force of the CUSFTA that weakens lobbying activity for protection. While this could be a valid interpretation, the difference in $\phi_3$ estimates do not render itself a simple explanation since it is not clear why a potential measurement error in elasticity measures would only affect the estimates of the political economy factors, producing similar results for the effect of the MFN tariff change and the US export share.

As a final robustness test, we exclude the pre-CUSFTA import tariff from the long-run model. Since CUSFTA member countries committed to complete elimination of import tariffs by 1998, the change in preferential tariff between 1989 and 1998 is highly correlated with the

[22] An increase in the $\phi_1$ coefficient is largely offset by a reduction in $\Delta s_{it}^{P}$ in the estimation sample.

starting value of import tariff, which may cause a high degree of multicollinearity between 1989 tariff rate and $X_{it}^1$ variable. The results on column (9) reveal only a marginal increase in $\phi_1$ coefficient, indicating that multicollinearity is unlikely to be a serious problem.

# 4 Conclusion

Whether PTAs induce or deter the incentive of member countries for multilateral trade liberalization has been a central question in the regionalizm literature for the last few decades. So far, no consensus has been reached on the effect of an FTA membership on external tariffs. The theoretical literature on regionalism proposed several channels for the effect of an FTA on multilateral tariffs which can rationalize both falls and rises in the level of protectionism following formation of an agreement. The empirical evidence on the relationship between FTA membership and MTL is inconclusive as well: while some agreements were found to slow down MTL, others resulted in deeper trade liberalization. Identifying the factors associated with one outcome or another is thus an important empirical question. In this paper, we provide further evidence on the relationship between preferential trade liberalization and MFN tariffs by analyzing the effect of the CUSFTA on Canadian external tariffs. To test this relationship, we develop a model of endogenous trade policy formation that combines several forces leading to complementarity and substitutability between FTA internal and external tariffs, which allows us to analyze the relative importance of those forces for Canadian MTL in a unified empirical framework.

The main finding of this paper is that the CUSFTA facilitated greater liberalization of the Canadian multilateral tariffs. The main factor contributing to complementarity between preferential and MFN tariffs operates through the terms-of-trade and tariff revenue effects. We find that a one percentage point reduction in the Canadian preferential tariff rate lead to $0.3 - 0.35$ percentage points reduction in the MFN tariff, which accounts for around 55% of tariff decline observed during the Uruguay round of the WTO negotiations. This result implies that the size of the partner country may play an important role for the effect of an FTA on incentives to liberalize trade multilaterally since the effect of an FTA on the terms-of-trade and tariff revenue is small when the partner country is small.

At the same time, political economy factors do not seem to be an important channel for the effect of the CUSFTA on MTL. We failed to find any consistent evidence on the negative impact of the CUSFTA on lobbying power of domestic special interest groups. Despite the theoretical prediction that intensified competition with the US firms and declining domestic market share should have negative impact on the return to lobbying activity and reduce incentives for lobbying, we do not observe deeper tariff reductions in industries with strong political connections.

Our study also reveals the presence of trade policy cooperation between Canada and the US. We show that industries generating more export revenue for the US were liberalized less during the Uruguay round and tend to remain more protective afterwards. This result is consistent with the hypothesis that the Canadian government is reluctant to erode the rents of the US exporters generated by their preferential treatment. The finding that the CUSFTA partner countries internalize the impact of their MFN tariffs on each other seems to suggest that the agreement was used to stimulate cooperation on non-trade issues. However, the dominance of tariff complementarity effect of the CUSFTA also suggests that the main purpose of the agreement was to exchange market access between the two countries.

# References

Bagwell, K., and R. W. Staiger (1997a): "Multilateral tariff cooperation during the formation of customs unions," *Journal of International Economics*, 42(1-2), 91–123.

——— (1997b): "Regionalism and Multilateral Tariff Cooperation," NBER Working Papers 5921, National Bureau of Economic Research, Inc.

——— (1998): "Multilateral Tariff Cooperation During the Formation of Free Trade Areas," in *International Trade Policy and the Pacific Rim*, ed. by J. Piggott, and A. Woodland. MacMillan.

Balassa, B. (1965): "Trade Liberalisation and Revealed Comparative Advantage1," *The Manchester School*, 33(2).

Bond, E. W., R. G. Riezman, and C. Syropoulos (2004): "A strategic and welfare theoretic analysis of free trade areas," *Journal of International Economics*, 64(1), 1–27.

Bown, C., and M. Crowley (2007): "Trade Deflection and Trade Depression," *Journal of International Economics*, 72(1), 176–201.

Broda, C., J. Greenfield, and D. Weinstein (2006): "From Groundnuts to Globalization: A Structural Estimate of Trade and Growth," (12512).

Estevadeordal, A., C. Freund, and E. Ornelas (2008): "Does Regionalism Affect Trade Liberalization Toward Nonmembers?," *The Quarterly Journal of Economics*, 123(4), 1531–1575.

Freund, C. (2000): "Multilateralism and the endogenous formation of preferential trade agreements," *Journal of International Economics*, 52(2), 359–376.

Gawande, K., and U. Bandyopadhyay (2000): "Is Protection for Sale? Evidence on the Grossman-Helpman Theory of Endogenous Protection," *The Review of Economics and Statistics*, 82(1), 139–152.

Grossman, G. M., and E. Helpman (1994): "Protection for Sale," *American Economic Review*, 84(4), 833–850.

Grossman, G. M., and E. Helpman (1995): "The Politics of Free-Trade Agreements," *American Economic Review*, 85(4), 667–90.

Imai, S., H. Katayama, and K. Krishna (2009): "Protection for sale or surge protection?," *European Economic Review*, 53(6), 675–688.

Karacaovali, B., and N. Limao (2008): "The clash of liberalizations: Preferential vs. multilateral trade liberalization in the European Union," *Journal of International Economics*, 74(2), 299–327.

Kee, H. L., A. Nicita, and M. Olarreaga (2009): "Estimating Trade Restrictiveness Indices," *Economic Journal*, 119(534), 172–199.

Kennan, J., and R. Riezman (1990): "Optimal Tariff Equilibria with Customs Unions," *Canadian Journal of Economics*, 23(1), 70–83.

Ketterer, T., D. Bernhofen, and C. Milner (2012): "Preferences, Rent Destruction and Multilateral Liberalisation: The Building Block Effect of CUSFTA," working paper.

Krishna, P. (1998): "Regionalism And Multilateralism: A Political Economy Approach," *The Quarterly Journal of Economics*, 113(1), 227–250.

Limao, N. (2006): "Preferential Trade Agreements as Stumbling Blocks for Multilateral Trade Liberalization: Evidence for the United States," *American Economic Review*, 96(3), 896–914.

——— (2007): "Are Preferential Trade Agreements with Non-trade Objectives a Stumbling Block for Multilateral Liberalization?," *Review of Economic Studies*, 74(3), 821–855.

Matschke, X. (2008): "Costly revenue-raising and the case for favoring import-competing industries," *Journal of International Economics*, 74(1), 143–157.

Ornelas, E. (2005a): "Endogenous Free Trade Agreements and the Multilateral Trading System," *Journal of International Economics*, 67(2), 471–497.

——— (2005b): "Trade creating free trade areas and the undermining of multilateralism," *European Economic Review*, 49(7), 1717–1735.

——— (2007): "Exchanging market access at the outsiders' expense: the case of customs unions," *Canadian Journal of Economics*, 40(1), 207–224.

Panagariya, A., and R. Findlay (1994): "A political-economy analysis of free trade areas and customs unions," Policy Research Working Paper Series 1261, The World Bank.

Richardson, M. (1993): "Endogenous Protection and Trade Diversion," *Journal of International Economics*, 34(3-4), 309–324.

Stoyanov, A. (2009): "Trade Policy of A Free Trade Agreement in the Presence of Foreign Lobbying," *Journal of International Economics*, 77(1), 37–49.

Trefler, D. (1993): "Trade Liberalization and the Theory of Endogenous Protection: An Econometric Study of U.S. Import Policy," *Journal of Political Economy*, 101(1), pp. 138–160.

Wooldridge, J. M. (2001): *Econometric Analysis of Cross Section and Panel Data*, vol. 1 of *MIT Press Books*. The MIT Press.

Table 1. Summary statistics for key variables

| | Mean | Standard deviation | Minimum | Maximum | Number of observations |
|---|---|---|---|---|---|
| Tariff_row | 0.057 | 0.064 | 0 | 0.249 | 38,445 |
| ΔTariff_row | -0.004 | 0.013 | -0.068 | 0.052 | 38,445 |
| Tariff_us | 0.025 | 0.04 | 0 | 0.222 | 38,445 |
| ΔTariff_us | -0.007 | 0.011 | -0.054 | 0.031 | 38,445 |
| Share_can | 0.578 | 0.196 | 0.066 | 0.999 | 41,204 |
| ΔShare_can | -0.011 | 0.032 | -0.445 | 0.445 | 40,779 |
| Share_us | 0.253 | 0.144 | 0.001 | 0.797 | 41,190 |
| ΔShare_us | 0.009 | 0.026 | -0.392 | 0.376 | 40,758 |
| I1 | 0.506 | 0.501 | 0 | 1 | 243 |
| I2 | 0.239 | 0.427 | 0 | 1 | 243 |
| I3 | 0.235 | 0.425 | 0 | 1 | 243 |
| I4 | 0.453 | 0.499 | 0 | 1 | 243 |
| Elasticity_kno | -2.958 | 4.906 | -37.979 | -0.213 | 4,018 |
| Log scale | 16.143 | 1.286 | 13.593 | 21.805 | 243 |
| Material share | 0.511 | 0.118 | 0.164 | 0.898 | 243 |
| Labour share | 0.191 | 0.073 | 0.015 | 0.37 | 243 |
| Non-prod. Labour share | 0.202 | 0.089 | 0.057 | 0.594 | 243 |
| Fuel share | 0.027 | 0.039 | 0.001 | 0.314 | 243 |

Notes: Summary statistics is calculated for 6-digit HS industries for the time period 1989-1998. Political activity indicators I1 and I2 equal unity if an industry has at least one and three lobbyists, respectively. Political Indicator I3 is constructed as in Gawande and Bandyopadhyay (2000), and I4 is constructed as in Matschke (2008). Import demand elasticities, Elasticity_kno, are obtained from Kee, Nicita, and Olarreaga (2009).

Table 2. Estimation results for the reduced-form short-run model

| | (1) OLS | (2) OLS | (3) OLS | (4) OLS | (5) IV-GMM | (6) IV-GMM | (7) IV-GMM | (8) IV-GMM |
|---|---|---|---|---|---|---|---|---|
| $\Delta tariff\_us_{it-1}$ | 0.1053*** | | | 0.1054*** | | | 0.1000*** | 0.0987*** |
| | (6.69) | | | (6.97) | | | (6.30) | (5.25) |
| $\Delta tariff\_us_{it-2}$ | | | | | | | | 0.0491*** |
| | | | | | | | | (4.33) |
| $\Delta tariff\_us_{it-3}$ | | | | | | | | 0.0316*** |
| | | | | | | | | (4.51) |
| $D$ † | | 0.0009 | | 0.0019 | -0.0067 | | -0.0003 | -0.0017 |
| | | (0.73) | | (1.29) | (-1.30) | | (-0.00) | (-0.01) |
| $D^1$ † | | | -0.0000 | 0.0003 | | -0.0006 | -0.0001 | 0.0010 |
| | | | (-0.04) | (0.83) | | (-1.09) | (-0.06) | (0.49) |
| $D^2$ † | | | -0.0006 | -0.0002 | | -0.0014*** | -0.0006 | -0.0004 |
| | | | (-1.43) | (-0.62) | | (-2.74) | (-0.99) | (-0.42) |
| $D^3$ † | | | -0.0002 | -0.0000 | | -0.0009 | -0.0005 | -0.0003 |
| | | | (-0.71) | (-0.12) | | (-1.63) | (-1.10) | (-0.50) |
| $D^4$ † | | | -0.0001 | 0.0001 | | -0.0004 | -0.0004 | -0.0003 |
| | | | (-0.23) | (0.38) | | (-1.05) | (-1.04) | (-0.61) |
| R-squared | 0.050 | 0.041 | 0.041 | 0.050 | | | | |
| Hansen J-statistics, p-value [(a)] | | | | | 0.38 | 0.52 | 0.49 | 0.48 |
| Endogeneity test, p-value [(b)] | | | | | 0.256 | 0.000 | 0.006 | 0.129 |
| N | 37,508 | 38,854 | 38,854 | 37,508 | 37,170 | 37,170 | 36,190 | 28,390 |

Notes: The dependent variable is the annual change in the MFN tariff. * Significant at 10%, ** significant at 5%, *** significant at 1% confidence level. (a) Test for overidentifying restrictions. The null hypothesis is that instruments are exogenous. (b) Housman specification test for endogeneity of variables marked with "†". Under the null hypothesis the variables are exogenous and the OLS is consistent. Standard errors are clustered at 6-digit NAICS industry level. All specifications include 1988 MFN tariff rate as an additional control.

Table 3. Estimation results for the short-run structural model

| | (1) OLS | (2) IV-GMM | (3) OLS | (4) IV-GMM |
|---|---|---|---|---|
| $\Delta X_{it-1}^{1}$ † | 1.6817*** | 1.6841** | 1.7798*** | 1.4723** |
| | (3.72) | (2.28) | (3.92) | (1.99) |
| $\Delta X_{it-1}^{2}$ † | 0.0362 | -0.2082 | 0.0367 | -0.1014 |
| | (1.09) | (-0.91) | (1.09) | (-0.94) |
| $D$ † | | | 0.0284 | -0.4769 |
| | | | (1.06) | (-0.31) |
| $D^{1}$ † | | | -0.0139*** | -0.0098 |
| | | | (-3.12) | (-0.69) |
| $D^{2}$ † | | | -0.0049** | -0.0002 |
| | | | (-2.11) | (-0.02) |
| $D^{3}$ † | | | -0.0026 | -0.0068 |
| | | | (-1.47) | (-0.90) |
| $D^{4}$ † | | | -0.0025* | 0.0051 |
| | | | (-1.72) | (1.30) |
| R-squared | 0.018 | | 0.019 | |
| Hansen J-statistics, p-value [(a)] | | 0.45 | | 0.50 |
| Endogeneity test, p-value [(b)] | | 0.194 | | 0.401 |
| N | 25,193 | 14,035 | 25,193 | 14,035 |

Notes: * Significant at 10%, ** significant at 5%, *** significant at 1% confidence level. (a) Test for overidentifying restrictions. The null hypothesis is that instruments are exogenous. (b) Housman specification test for endogeneity of variables marked with "†". Under the null hypothesis the variables are exogenous and the OLS is consistent. Standard errors are clustered at 6-digit NAICS industry level. All specifications include 1988 MFN tariff rate as an additional control.

Table 4. Estimation results for the reduced-form long-run model

| | (1) OLS | (2) OLS | (3) OLS | (4) OLS | (5) IV-GMM | (6) IV-GMM | (7) IV-GMM |
|---|---|---|---|---|---|---|---|
| $\Delta tariff\_us_{it}$ | 0.1996*** | | | 0.2009*** | | | 0.1980*** |
| | (6.34) | | | (5.94) | | | (5.90) |
| $D$ † | | -0.0078 | -0.0017 | | -13.1191 | 0.4996 | |
| | | (-0.72) | (-0.15) | | (-1.01) | (0.75) | |
| $D^1$ † | | | 0.0104*** | 0.0104*** | | 0.0238*** | 0.0169*** |
| | | | (4.03) | (4.20) | | (2.90) | (5.05) |
| $D^2$ † | | | 0.0049*** | 0.0052*** | | 0.0007 | 0.0052 |
| | | | (2.66) | (2.88) | | (0.14) | (1.13) |
| $D^3$ † | | | 0.0051*** | 0.0052*** | | 0.0029 | 0.0009 |
| | | | (3.07) | (3.29) | | (0.63) | (0.21) |
| $D^4$ † | | | 0.0027* | 0.0030** | | 0.0022 | 0.0027 |
| | | | (1.75) | (2.01) | | (0.57) | (0.69) |
| R-squared | 0.231 | 0.216 | 0.225 | 0.24 | | | |
| Hansen J-statistics, p-value [(a)] | | | | | 0.00 | 0.13 | 0.17 |
| Endogeneity test, p-value [(b)] | | | | | 0.365 | 0.100 | 0.099 |
| N | 3,864 | 3,887 | 3,887 | 3,864 | 3,785 | 3,785 | 3,764 |

Notes: * Significant at 10%, ** significant at 5%, *** significant at 1% confidence level. (a) Test for overidentifying restrictions. The null hypothesis is that instruments are exogenous. (b) Housman specification test for endogeneity of variables marked with "†". Under the null hypothesis the variables are exogenous and the OLS is consistent. Standard errors are clustered at 6-digit NAICS industry level. All specifications include 1988 MFN tariff rate as an additional control.

Table 5. Estimation results for the long-run structural model

| | (1) OLS | (2) IV-GMM | (3) OLS | (4) IV-GMM |
|---|---|---|---|---|
| $\Delta X_{it}^{1}$ † | 3.9466*** | 7.5771*** | 4.0105*** | 7.7139*** |
| | (3.63) | (3.68) | (3.89) | (3.25) |
| $\Delta X_{it}^{2}$ † | 0.0784*** | 0.0028 | 0.1253*** | -0.0363 |
| | (2.68) | (0.01) | (3.25) | (-0.26) |
| $D^1$ † | | | -0.1522*** | -0.1139 |
| | | | (-3.38) | (-1.30) |
| $D^2$ † | | | -0.0755*** | -0.0254 |
| | | | (-3.25) | (-0.25) |
| $D^3$ † | | | -0.0354* | 0.0106 |
| | | | (-1.77) | (0.19) |
| $D^4$ † | | | -0.0214* | 0.0388 |
| | | | (-1.93) | (1.03) |
| R-squared | 0.076 | | 0.093 | |
| Hansen J-statistics, p-value [(a)] | | 0.73 | | 0.77 |
| Endogeneity test, p-value [(b)] | | 0.027 | | 0.024 |
| N | 3,178 | 2,315 | 3,178 | 2,315 |

Notes: * Significant at 10%, ** significant at 5%, *** significant at 1% confidence level. (a) Test for overidentifying restrictions. The null hypothesis is that instruments are exogenous. (b) Housman specification test for endogeneity of variables marked with "†". Under the null hypothesis the variables are exogenous and the OLS is consistent. Standard errors are clustered at 6-digit NAICS industry level. All specifications include 1988 MFN tariff rate as an additional control.

Table 6. Estimation results for the political economy model (11)

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
|---|---|---|---|---|---|---|---|---|
| Model: | | Short run | | | | Long run | | |
| $\Delta X_{it}^{1}$ † | 1.3720* | 1.3907* | 1.3231* | 1.3709* | 6.2005*** | 5.9994*** | 6.3499*** | 6.3903*** |
| | (1.93) | (1.95) | (1.85) | (1.90) | (3.08) | (2.89) | (3.19) | (3.22) |
| $\Delta X_{it}^{2}$ † | -0.3303 | -0.2243 | -0.5124 | -0.1670 | 0.0247 | 0.0117 | -0.0506 | -0.0284 |
| | (-0.90) | (-0.82) | (-1.27) | (-0.39) | (0.15) | (0.07) | (-0.43) | (-0.21) |
| $\Delta X_{it}^{3}$ † | 0.1650 | -0.1081 | 0.3036 | 0.0482 | -0.1286 | -0.2016 | -0.0053 | -0.0539 |
| | (0.44) | (-0.25) | (0.79) | (0.11) | (-0.75) | (-0.61) | (-0.03) | (-0.36) |
| $D^{1}$ † | -0.0123*** | -0.0118*** | -0.0121*** | -0.0119*** | -0.0671* | -0.0613* | -0.0647* | -0.0627* |
| | (-3.02) | (-2.91) | (-2.88) | (-3.05) | (-1.90) | (-1.66) | (-1.86) | (-1.73) |
| $D^{2}$ † | -0.0029 | -0.0026 | -0.0033 | -0.0025 | -0.0529** | -0.0522* | -0.0509* | -0.0510* |
| | (-1.01) | (-0.92) | (-1.07) | (-0.92) | (-1.99) | (-1.96) | (-1.89) | (-1.91) |
| $D^{3}$ † | -0.0073*** | -0.0072*** | -0.0073*** | -0.0072*** | -0.0210 | -0.0184 | -0.0178 | -0.0175 |
| | (-3.54) | (-3.40) | (-3.45) | (-3.48) | (-0.96) | (-0.88) | (-0.84) | (-0.85) |
| $D^{4}$ † | -0.0031** | -0.0029* | -0.0035** | -0.0029* | -0.0131 | -0.0097 | -0.0110 | -0.0101 |
| | (-2.00) | (-1.74) | (-2.04) | (-1.95) | (-0.74) | (-0.58) | (-0.66) | (-0.63) |
| Hansen J-statistics, p-value [(a)] | 0.26 | 0.27 | 0.33 | 0.23 | 0.44 | 0.46 | 0.26 | 0.29 |
| Endogeneity test, p-value [(b)] | 0.02 | 0.019 | 0.006 | 0.072 | 0.001 | 0.001 | 0.002 | 0.001 |
| N | 14,035 | 14,035 | 14,035 | 14,035 | 2,315 | 2,315 | 2,315 | 2,315 |

Notes: * Significant at 10%, ** significant at 5%, *** significant at 1% confidence level. (a) Test for overidentifying restrictions. The null hypothesis is that instruments are exogenous. (b) Housman specification test for endogeneity of variables marked with "†". Under the null hypothesis the variables are exogenous and the OLS is consistent. Standard errors are clustered at 6-digit NAICS industry level. All specifications include 1988 MFN tariff rate as an additional control.

Table 7. Robustness tests and extension

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
|---|---|---|---|---|---|---|---|---|---|
| Model: | short run | long run | short run | long run | short run | long run | short run | long run | long run |
| $\Delta X_{it}^{1}$ † | 1.4036* | 5.4073*** | 1.4460** | 5.9109** | 2.5364** | 9.3757* | 0.1839*** | 1.3507*** | 7.8504*** |
| | (1.85) | (2.66) | (2.04) | (2.48) | (2.45) | (1.85) | (2.63) | (6.68) | (4.18) |
| $\Delta X_{it}^{2}$ † | 0.8553 | -0.2759 | -0.3516 | 0.0799 | 0.0296 | 0.3795 | 0.0511 | -0.0293* | -0.0157 |
| | (1.55) | (-1.48) | (-0.82) | (0.34) | (0.06) | (0.91) | (0.95) | (-1.91) | (-0.10) |
| $\Delta X_{it}^{3}$ † | -1.0583 | 0.3300 | 0.1201 | -0.2339 | -1.0918 | -0.7995 | -0.0409 | 0.0434** | -0.1394 |
| | (-0.66) | (1.04) | (0.29) | (-0.99) | (-1.20) | (-1.38) | (-0.80) | (2.24) | (-0.82) |
| $D^{1}$ † | -0.0189*** | -0.0864** | -0.0173*** | -0.0935** | -0.0161*** | -0.1475** | -0.0018*** | -0.0085** | -0.0618* |
| | (-3.16) | (-2.32) | (-3.55) | (-2.36) | (-2.67) | (-2.39) | (-3.50) | (-2.09) | (-1.69) |
| $D^{2}$ † | -0.0062 | -0.0700** | -0.0049 | -0.0759** | -0.0077* | -0.1053** | -0.0008*** | -0.0037 | -0.0501* |
| | (-1.43) | (-2.24) | (-1.35) | (-2.48) | (-1.69) | (-2.21) | (-2.75) | (-1.33) | (-1.92) |
| $D^{3}$ † | -0.0079** | -0.0342 | -0.0093*** | -0.0356 | -0.0116*** | -0.0902* | -0.0006** | -0.0057** | -0.0161 |
| | (-2.47) | (-1.52) | (-3.54) | (-1.35) | (-2.81) | (-1.72) | (-2.03) | (-2.44) | (-0.73) |
| $D^{4}$ † | -0.0023 | -0.0201 | -0.0039** | -0.0217 | -0.0044 | -0.0704 | -0.0003 | -0.0027 | -0.0100 |
| | (-0.57) | (-1.51) | (-2.07) | (-1.01) | (-1.36) | (-1.34) | (-0.99) | (-1.56) | (-0.58) |
| Condition | HS2 fixed effects | | No industries with zero tariffs in 1988 | | Only industries with gradual preferential tariff reductions | | No elasticities in the dependent variable | | No initial tariff |
| Hansen J-statistics, p-value (a) | | | 0.27 | 0.51 | 0.48 | 0.90 | 0.02 | 0.24 | 0.38 |
| Endogeneity test, p-value (b) | | | 0.563 | 0.036 | 0.667 | 0.058 | 0.087 | 0.055 | 0.004 |
| N | 14,035 | 2,315 | 11,378 | 1,895 | 6,671 | 1,104 | 17,435 | 2,883 | 2,317 |

Notes: * Significant at 10%, ** significant at 5%, *** significant at 1% confidence level. (a) Test for overidentifying restrictions. The null hypothesis is that instruments are exogenous. (b) Housman specification test for endogeneity of variables marked with "†". Under the null hypothesis the variables are exogenous and the OLS is consistent. Standard errors are clustered at 6-digit NAICS industry level. Columns (1)-(8) include 1988 MFN tariff rate as an additional control. Columns (1) and (2) include 2-sigit HS industry fixed effects. Column (3) and (4) exclude industries with zero MFN tariff in 1988. Column (5) and (6) are estimated on industries with 10-year phase-out periods for the CUSFTA preferential tariff reductions. In columns (7) and (8) the dependent variable is the change in the MFN import tariff.