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

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Abstract

In this paper show that the Canada-US Free Trade Agreement (CUSFTA) tariff preferences have triggered a decline in Canadian external tariffs, explaining a two percentage point reduction in the average tariff between 1989 and 1998. Next, we found that industries which generate the least export rent to the US firms experienced deeper tariff cuts in Canada; this result provides evidence of 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 liberalization and lobbying activity.

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 concerning 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, 2007). 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 specific nations.

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. This model 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 demonstrate that the CUSFTA tariff preferences have resulted in considerable reduction in Canadian tariffs, explaining a 2.21 out of 4.02 percentage point decline in the average most favoured nation (MFN) tariff rate between 1989 and 1998. We also show that the MFN tariff reductions were deeper in industries that generated the least export revenue for US exporters. This finding suggests that the Canadian government may be more inclined to reduce preference margins for products which have smaller potential to generate rent to the partner country.

To outline our theoretical model, we consider the economy of three large countries and monopolistically competitive markets. We start by deriving the effect of an FTA on external tariffs if they were set non-cooperatively by welfare-maximizing governments. In this framework there are two sources of complementarity between external and preferential tariffs. First, there are the terms-of-trade and tariff revenue effects, similar to the ones obtained by Richardson (1993) and Bagwell and Staiger (1997b), which lead to a decline in MFN tariffs. A reduction in the MFN tariffs following the CUSFTA can moderate some efficiency loss caused by the distortionary effect of preferential market access on the relative price of imports (Bagwell and Staiger, 1997b; Freund, 2000; Ornelas, 2005b) and restore part of tariff revenue loss caused by the shift in import demand towards the US-produced goods (Richardson, 1993). Second, there is a market structure effect which stimulates policymakers to raise protection in industries with a large domestic presence and a 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 thus 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 by assuming that when the government of one country sets trade policy, it takes into account the effect on welfare of the partner country. When FTA member countries mutually internalize the effect of their policies, they keep external tariffs high in order to generate more export rent to their FTA partners. In a similar vein, Limao (2007) demonstrated that trade agreements with cooperative trade policies tend to become more protectionist in the context of multilateral trade negotiations. This "stumbling block" effect of an FTA enters the expression for the equilibrium tariff rate in additively separable way, which allows us to empirically identify its effect on the external tariffs independently from other influences.

As a last extension of our model, we incorporate political economy factors in policymakers' preferences using the protection-for-sale framework of Grossman and Helpman (1994). 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 would weaken the lobbying power of domestic special interest groups and reduce the level of protectionism. This effect is identified by Ornelas (2005b) as "rent destruction" since in the presence of an FTA a part of the rent from protection will flow to the partner country firms, making lobbying less attractive and weakening political economy distortions.

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. Estimating the model with Canadian tariff data during the time period of the CUSFTA implementation, the main findings of this study are as follows. First, the study reveals a strong tariff complementarity between Canadian preferential and MFN tariffs which propagates through 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 $0.3 - 0.35$ percentage points 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 results on the effect of the CUSFTA on lobbying activity, however, is inconclusive. Using various measures of industrial lobbying activity we failed to find any relationship between preferential tariff liberalization and MFN tariff changes in politically organized industries. These results echo Ketterer, Bernhofen, and Milner (2012) who found no effect of political economy factors on the Canadian MFN tariff reductions in 1990s.[1] Unlike most of the previous studies that estimate the effect of FTAs on the political economy of trade policy, we move away from the assumption that all industries are equally involved in lobbying. We differentiate industries by lobbying intensity using the data from the Canadian Lobbyists Registry in order to identify the number of lobbyists representing each industry. Through the use of these data, we construct several alternative measures of industrial lobbying activity and use the modified protection for sale model to structurally estimate the effect of the CUSFTA on lobbying for MFN tariffs. In most specifications we unable to find any evidence of deeper tariff reductions in politically active industries. While this result may imply that the "rent destruction" effect was not among the main factors of the Canadian trade policy, it may as well be driven by lack of reliable measures of sectorial lobbying intensity.

The evidence on the presence of trade policy cooperation in the CUSFTA is mixed. On one hand, we show that the Canadian MFN tariffs declined deeper in 20% of the industries comprising the smallest US exports to Canada. The result that the Canadian government is more open to trade liberalization in industries where it does not erode exports rent of the US exporters is consistent with the hypothesis of trade policy cooperation. On the other hand, we do not find any effect of the US exports rent on the Canadian MFN tariffs in the remaining 80% of the industries. This partial evidence of trade cooperation is consistent with the findings by Limao (2006) and Karacaovali and Limao (2008) for the US and the EU, however 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 99% of all 6-digit HS products are exported by the US to Canada, which makes the identification of the stumbling block effect difficult as it relies on a very small number of industries. Our identification strategy, derived from the model of cooperative trade policy formation, relies on a richer cross-industry variation in the preference rent collected by the partner country. This allows us to run a more general test of the trade policy cooperation hypothesis.

The rest of the paper is organized as follows. Section 2 introduces a theoretical model of endogenous trade policy in the presence of an FTA, presenting the results on the effect of trade agreements on non-cooperative (Section 2.1), cooperative (Section 2.2), and political (Section 2.3) trade policies. Sections 3.1 and 3.2 develop the analytical framework for estimating the effect of FTAs on external import tariffs, and Sections 3.4 and 3.5 present regression results and extensions. Finally, section 4 concludes with a summary of our findings.

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 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 n_{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:

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

where σ_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 τ_{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 (π_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 τ 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 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 rely on empirical analysis to differentiate amongst 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]

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

where ε_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 termed as the "tariff complementarity" effect due to Bagwell and Staiger (1997b). Intuitively, a decline in the tariff rate towards the FTA partner country reduces imports from the ROW, thus reducing tariff revenue proportional to the partner country's market share s_i^P. Furthermore, the tariff revenue effect is stronger if varieties imported from the partner country and from the ROW exports are close substitutes. The second term on the right-hand side reflects the government's incentives to protect imperfectly competitive industries. This term 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 the reallocating effect of an import tariff is proportional to s_{it}^{H}. Moreover, the ability of trade policy to redistribute expenditure from foreign to domestic varieties is stronger when these varieties are close substitutes.

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, tend to increase the MFN tariff relative to pre-CU level. There are two main contributing factors to the protectionist trade policies of a CU. Firstly, there is the terms of trade argument that arises from an increase in the economic size of the trading block, which in turn increases its market power and thereby allows member countries to redistribute surplus from their non-member trading partners. Secondly, CUs tend to have higher tariffs because their members take into account the effect of a CET on each others welfare. The cooperative trade policy of a CU internalizes the positive effect of a CET on export rents within the block, thus making CUs more protective than FTAs.

While the first factor simply reflects the 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, then the resulting trade policy becomes more protectionist as the member countries internalize the externalities created by their trade policies. 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 have suggest that this may be the case. The theoretical model of an RTA, constructed by Limao (2007), features a public good supplied by individual countries which generates positive 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 the public good with cross-border spillover effects. Accordingly, RTA members want to maintain their preference margins by keeping the MFN tariff high in order to stimulate economic and political involvement of their partners in non-trade issues, thus internalizing the decision on the provision of the regional public good.

We model trade policy cooperation by assuming that FTA member countries take into account the effect of their trade policies on each other and set import tariffs in order to maximize the sum of their welfare. With segmented markets, the objective function of the home country government becomes the sum of national welfare and profits earned by the partner country's firms:

$$G_1(\tau) = W_0(\tau) + b\pi_P(\tau) = CS + TR + \sum_i (n_{iH}\pi_{iH} + bn_{iP}\pi_{iP}) \tag{6}$$

Parameter $b \in [0;1]$ measures the degree of trade policy cooperation between the two countries. When $b = 0$, $G_1(\tau) = G_0(\tau)$ and there is no trade policy cooperation. When $b = 1$, home country policymakers internalize the effect of τ on partner country's welfare completely. Differentiating the wights on foreign and domestic welfare in the objective function allows us to empirically test the trade cooperation hypothesis against the alternative of no cooperation.

In the presence of trade policy cooperation, the equilibrium import tariff with and without the agreement becomes:[6]

$$
\begin{aligned}
\varepsilon_i t_i^F &= (\sigma_i - 1)\, t_i^P s_{iP} + \frac{\sigma_i - 1}{\sigma_i} s_{iH} + b\frac{\sigma_i - 1}{\sigma_i} s_{iP} && \text{with FTA} \\
\varepsilon_i t_i^F &= (\sigma_i - 1)\, t_i^P s_{iP} + \frac{\sigma_i - 1}{\sigma_i} s_{iH} - b\frac{\sigma_i - 1}{\sigma_i} \left(\frac{s_{iP}}{s_{iP} + s_{iF}}\right) && \text{without FTA}
\end{aligned}
$$

and the effect of the agreement on the MFN tariff is

$$
\frac{\sigma_i}{\sigma_i - 1} \varepsilon_i \Delta t_{it}^F = \sigma_i \Delta \left(s_{it}^P t_{it}^P\right) + \Delta s_{it}^H + b\left(s_{iP}^1 + \frac{s_{iP}^0}{s_{iP}^0 + s_{iF}^0}\right) \tag{7}
$$

As long as FTA member countries cooperate their policies, there is an additional stumbling block effect of an FTA (the last term on the right-hand side of equation 7). This effect originates from the incentive of the home country's government to maintain a large enough preference margin for the partner by increasing the MFN tariff subsequent to FTA formation (or by decreasing it insubstantially).

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:

$$
G_2(\tau) = aW_1(\tau) + C \tag{8}
$$

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 τ which maximizes

$$G_2(\tau) = aW_1(\tau) + \sum_i I_i C_i(\tau) \tag{9}$$

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

$$\frac{\sigma_i}{\sigma_i - 1}\varepsilon_i t_{it}^F = \sigma_i s_{it}^P t_{it}^P + \frac{a}{a+\alpha} s_{it}^H + \frac{1}{a+\alpha} I_i s_{it}^H + \frac{b}{a+\alpha} s_{iP} \tag{10}$$

where α is the share of population represented by one of the lobbying groups. The third term on the right-hand side 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.[7] As with the second 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 (10) 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 (2005b) who demonstrated that FTAs erode protection rent enjoyed by home country firms.

3 Empirical implementation

3.1 Econometric specifications

Equations (5), (7), and (10) summarize three main channels for the effect of an FTA on trade policies of member countries and 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:[8]

$$Y_{it} = \alpha + \phi_1 X^1_{it} + \phi_2 X^2_{it} + \gamma_i + \gamma_t + u_{it} \tag{11}$$

$$Y_{it} = \frac{\sigma_i - 1}{\sigma_i}\varepsilon_i t^F_{it},\ X^1_{it} = s^P_{it} t^P_{it},\ X^2_{it} = s^H_{it}$$

where γ_i and γ_t are industry and year fixed effects which capture the influences of time- and industry-invariant factors that are not present in the theoretical model but may affect trade policy formation. Given that the model is static and does not inform us about the dynamic response of MFN tariffs to changes in the right-hand side variables, we apply two alternative time-difference operators in order to identify coefficients ϕ_1 and ϕ_2:

$$\Delta Y_{it} = \alpha + \phi_1 \Delta X^1_{it-1} + \phi_2 \Delta X^2_{it-1} + \gamma_t + u_{it} \tag{12}$$

$$\Delta_9 Y_{i,98} = \alpha + \phi_1 \Delta_9 X^1_{i,1998} + \phi_2 \Delta_9 X^2_{i,1998} + u_i \tag{13}$$

Equation (12) is in first differences and measures short-term relationship between MFN and preferential tariffs (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 whilst most of the MFN tariff reductions, negotiated at the Uruguay Round of the WTO, occurs after 1995, there is still enough variation in both variables prior to 1995 to identify the presence of short-term response in MFN tariffs to CUSFTA trade liberalization.[9] 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 (11) 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 (12) can provide important information on short-term adjustments in trade policy to preferential liberalization, the more general long-run model (13) 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) in order to test the reduced form relationship between external and internal tariffs for a group of Latin American countries. Their findings confirm 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 US imports:

$$\Delta t_{it}^F = \alpha + \phi_0 \Delta t_{i,t-1}^P + \gamma_t + u_{it} \tag{14}$$

$$\Delta_9 t_{i,1998}^F = \alpha + \phi_0 \Delta_9 t_{i,1998}^P + u_i \tag{15}$$

Since tariff preferences to a partner country reduce socially optimal external tariffs both through the tariff revenue effect (X_{it}^1) and through the market structure effect (X_{it}^2), we expect $\phi_0 > 0$. Positive ϕ_0 would imply that reductions in preferential tariff are followed by MFN tariff cuts and support the tariff complementarity hypothesis.

The intuition for the test of the trade policy cooperation hypothesis comes from equation (7). When countries set their trade policies cooperatively, an FTA will have an additional positive effect on the MFN tariff which operates through two channels. First, cooperation implies lower tariffs in industries with large US market shares prior to the establishment of an FTA, and this effect vanishes once US firms receive preferential market access in Canada. Second, MFN tariffs generate rent to the partner country firms under the FTA so that the home country government will tend to provide more protection to industries which yield more rent to the partner country's firms. Together, these two effects induce an increase in the external tariffs amongst industries that have a large US presence in the Canadian market.

However, with trade policy restrictions imposed by the WTO, identification of cooperation in trade policy from equation (7) becomes problematic. First, the WTO tariff ceiling binding will either prevent tariffs from raising entirely or narrow the scope for increase to the gap between the binding and the applied tariff rates. Second, at the Uruguay Round (UR) of the WTO negotiations, the Canadian government committed to a 33% reduction in the average MFN tariff between 1995 and 1999, and by 1999 the applied MFN tariff rate had increased in only 3% of all industries relative to pre-CUSFTA levels. These two exogenous constraints on tariff adjustments imply that the effect of trade policy cooperation on changes in the MFN rates may not propagate through the US market share in the way predicted by equation (7).[10]

To identify the trade policy cooperation effect, we rely on the variation in MFN tariffs generated by the UR of tariff reductions. By varying the depth of tariff cuts across industries, countries had a considerable degree of flexibility in achieving the UR target of the reduction in the average MFN tariff by one-third. Although a common presumption is that the GATT trade liberalization is based on reciprocity, Finger, Reincke, and Castro (2002) point out that there was no specific formula applied to achieve the average tariff reduction target, giving negotiators some freedom in applying discretionary tariff cuts in different industries. Since tariff concessions are the outcome of a bargaining process in which every country protects its national interests, negotiators would trade-off interests of different industries and could extend protection to some industries at expense of deeper concessions in others. Therefore, the observed tariff concessions must reflect those national interests, and in particular the trade cooperation motive, if it is present in the objective function of a policymaker.

To understand how tariff concessions will differ across industries in the presence of trade policy cooperation, we differentiate the rent earned by the partner country exports to the home country with respect to the MFN tariff:

$$b\frac{\partial \Sigma_i \left(n_{iP}\pi_{iP}\right)}{\partial \tau_{iF}} = bn_{iP}\frac{\partial \pi_{iP}}{\partial \tau_{iF}} = b\xi_i \frac{\sigma_i - 1}{\sigma_i}\left(n_{iP}p_{iP}q_{iP}\right)$$

where $\xi_i = \frac{\partial P_i}{\partial \tau_{iF}}\frac{\tau_{iF}}{P_i}$ is the elasticity of the price index with respect to the MFN tariff. The MFN tariff affects a partner country's export rent through ξ_i (when ξ_i is large, tariffs have stronger effect on foreign firms' prices), elasticity of substitution (large σ_i implies stronger redistributive effect of price changes on consumer expenditure), and the value of imports from the partner country ($n_{iP}p_{iP}q_{iP}$). Therefore, when $b > 0$ and the partner country's rents enter the objective function of the home country's policymaker, a commitment to reduce the average MFN tariff by a certain amount will stimulate the home country to cut tariffs deeper in industries with smaller imports from the partner country, all else being equal. This is because in those industries a given preference margin applies to a smaller volume of exports and as such generates less rent to the partner country. In the absence of reliable data on the elasticity of substitution and price elasticity with respect to tariff, we rely on the variation in the value of US exports to identify the effect of trade policy cooperation.

The above result implies that if cooperation motives are present in the objective function of a policymaker, they should only play a role in industries where the partner country earns non-zero rent. We therefore introduce an indicator variable D_{it} which takes the value of one for goods imported from the US in specifications (12) and (13):

$$\Delta Y_{it} = \alpha + \beta_0 D_{it} + \phi_1 \Delta X^1_{it-1} + \phi_2 \Delta X^2_{it-1} + \gamma_t + u_{it} \tag{16}$$

As with equations (12) and (13), the model (16) is estimated using short and long time differencing. If $b > 0$, we would expect to find $\beta_0 < 0$. It should be emphasized that this test is identical to Limao (2006) and Karacaovali and Limao (2008) who derive the relationship between tariff reduction and preferential import indicator variables from the model where countries choose cooperative tariff rates in order to generate more rent to their FTA partner and to stimulate provision of the regional public good.

However, our preferred method for estimating the stumbling block effect is different from model (16) for two reasons. First, Canadian imports from the US are positive for nearly 99% of all 6-digit HS industries and identification of coefficient β_0 relies on too few observations. Second, in the presence of trade policy cooperation we would expect sectors with greater US involvement to observe less trade liberalization since the tariff concessions in those industries have stronger negative impact on the partner's export rent. Therefore, we differentiate industries according to their importance for US exports to Canada, and estimate the relationship between the volume of exports 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^k_{it} a dummy variable which takes the value of one if industry i falls into k-th quintile, the empirical specification becomes as follows:[11,12]

$$\Delta Y_{it} = \alpha + \sum_{k=1}^{4} \beta_k D^k_{it} + \phi_1(\sigma_i - 1)\Delta\left(s^P_{it} t^P_{it}\right) + \phi_2 \frac{\sigma_i - 1}{\sigma_i} \Delta s^H_{it} + \gamma_t + u_{it} \tag{17}$$

If tariff cooperation exists in the CUSFTA, industries that have higher US representation should receive more protection against foreign competition in the Canadian market. Thus, we expect all β_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 β_k coefficients:

$$\beta_{k-1} < \beta_k < 0, \forall k = 2, 3, 4 \tag{18}$$

Therefore, since the US exports to Canada are positive for nearly all 6-digit HS product categories, we use equation (17) to identify the cross-industry variation in the strength of the trade policy cooperation effect which varies with the partner country's gain from tariff preferences.

Finally, to arrive at our most complete empirical specification with political economy factors, we rearrange equation (10) by adding fixed effects and time differencing it:

$$\Delta Y_{it} = \alpha + \beta_0 D_{it} + \phi_1 \Delta X^1_{it-1} + \phi_2 \Delta X^2_{it-1} + \phi_3 \Delta X^3_{it-1} + \gamma_t + u_{it} \tag{19}$$

$$\Delta_9 Y_{i,98} = \alpha + \beta_0 D_i + \phi_1 \Delta_9 X^1_{i,1998} + \phi_2 \Delta_9 X^2_{i,1998} + \phi_3 \Delta_9 X^3_{i,1998} + u_i \tag{20}$$

where $X^3_{it} = I_i s^H_{it}$. Positive coefficient ϕ_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.[13] The reason for deeper tariff cuts is that FTAs lead to a reduction in protectionist rent retained by domestic firms since a part of this rent will be netted by the partner country's firms. This "rent destruction" effect, originally identified by Ornelas (2005b), 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. Preferential tariff cuts is the primary concern for potential endogeneity. The CUSFTA came into force on January 1, 1989, and resulted in the 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 CUSFTA was negotiated. 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, Δt^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 Δt^P with the US market share captures the economic value of tariff complementarity effect for tariff revenue. Yet, as a robustness test, we also run specifications where Δt^P and ΔX^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 the 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 the endogeneity of the FTA's partner 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 of the CUSFTA tariff cuts, which makes this instrument independent of tariff preferences. Our second instrument is the dummy variable which takes the value of one for products exported by the US to the ROW in 1988. This is a valid instrument because, on one hand, the US export structure to other countries prior to the CUSFTA formation is independent of Canadian trade policy during the years 1989-1998, and, on the other hand, positively correlated with the structure of the US exports to Canada. The third instrument is the change in the world price for product i, which is measured as the absolute change in price in the previous year for the short-run specifications and as a change between 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 price changes as 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 but 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 the estimation of equations (19) and (20) 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's output using 6-digit NAICS industry classification. Trefler (1993) suggests that industry's 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).[14] 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 firms and a 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 index is affected by the Canadian MFN tariff.[15] Similarly, the index of the US revealed comparative 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 (19) and (20) are estimated by 2-step GMM and all instruments which do not pass the orthogonality to the structural error test at a 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.[16] 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, which is 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 the 6-digit HS classification, and whenever data are available only at a higher level of aggregation, it is replicated for all 6-digit HS codes within the corresponding aggregate industry.[17] 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 value of imports.[18] 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 a 6-digit HS classification. Elasticities of substitution for Canada, σ_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 lobbyists officially registered with the Canadian Registrar of Lobbyists, the subject-matter for communicating with government officials, and the firms which recruited them. Working with only those lobbyists who contact policymakers regarding international trade policy issues allows us to construct political economy variables with a more pronounced relationship to trade policy formation. 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 non-active, we, thus, construct four different measures of industrial political organization to analyze the sensitivity 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_1 and I_2, 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 a 3-digit NAICS dummy variables and a constant term.[19] 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 greater protection from the government. We label this variable as I_3.

In construction of the fourth measure of political organization dummy, we follow 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 ($I_4 = 1$), while others are not. This specification is motivated by the theoretical prediction that in industries with larger welfare losses from protection domestic interest groups should spend more resources on lobbying and recruit more lobbyists.

The last mechanism for constructing political organization relies on the rates of Canadian preferential liberalization. Tariff reduction schedules between Canada and the US classified all products into three categories. Tariffs on products in the first category were eliminated entirely in the first year of the agreement, and tariffs for the other two groups were eliminated in equal annual stages over five and ten years, respectively. Assuming that the most politically active industries would be sheltered by more protectionist tariff reduction schedules, we classify all industries in the third category as politically organized. In other words, we use information on the observed changes in trade policy to reveal "sensitive" industries, which makes this measure better connected to lobbying for trade policy than the ones based on lobbyist headcount. At the same time, we should keep in mind that industries may be sensitive for a variety of factors unrelated to lobbying.

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 (14). The positive and statistically significant estimate of ϕ_1 coefficient in column (1) supports the tariff complementarity hypothesis and indicates that tariff preferences granted to the US 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 nearly identical to the estimates of $0.1 - 0.12$ obtained by EFO for Latin American countries in comparable empirical specifications. 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 changes in the MFN tariff on the indicator variable D_{it} in column (2). Canadian preferential tariffs provide US firms with a competitive 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 the 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 more than 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).[20] At the same time, the evidence on slower MFN tariff reduction in industries which have more economic significance for the US is weak as only one of the dummy variables has expected sign and is statistically significant.

The results presented so far focus on the reduced-form short-run relationship between MFN and preferential tariffs. In Table 3 we report the regression results for short-run specifications derived from the theoretical model of an FTA with endogenous trade policy. Column (1) illustrates the estimation results for equation (12) derived from the model with non-cooperative trade policy formation (5). The positive and statistically significant estimate of ϕ_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 (17) 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 reductions, 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.[21] Furthermore, the ranking of β_k coefficients confirms that industries contributing relatively more to the US exports to Canada receive smaller reduction in multilateral tariffs. In results from IV regressions in column (4) this ranking is not preserved, however the hypothesis that the industries with less imports from the US are liberalized at a faster rate still cannot be rejected.[22]

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 Canadian multilateral tariffs. In Table 4 we report the estimates for equation (15) 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 CUSFTA trade liberalization time period. 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 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 result seems to suggest deeper tariff cuts in industries with a larger US presence. However, these aforementioned results should be treated with caution. When the relationship between the MFN and preferential tariffs is estimated using the structural model (13), the coefficients on D_{it}^k become negative (Table 5, column 3). The likely reason for this is that larger tariff preference may lead to a larger increase in the US presence in Canadian markets, making the MFN import tariff less efficient in protecting domestic producers and, thus, weakening protectionist forces. Once the effect of the US market share in Canada 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 the most complete specification (column (4) of Table 5), industries in the first quintile of the US export share to Canada experience additional 4.57 percentage point reduction in the MFN tariff relative to industries in the fifth quintile over the period 1989-1998. In fact, it is only 20% of industries with the smallest US exports to Canada which observe larger reductions in the MFN tariff. The effect of the size of the US exports does not vary across the remaining industries.

In sum, there is strong evidence that Canadian MFN tariff rates feature complementarity with 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 US exporters. This suggests that Canadian policymakers at least partially internalize the effect of MFN tariff choice on US producers. The IV results are less conclusive, but we can never reject the hypothesis that industries in the fifth quintile of the US export share distribution experience greater reductions in the MFN tariffs than industries in the first three quintiles.

We now turn to empirically testing the final prediction of the theoretical model concerning the effect of an FTA on MFN tariffs in the presence of political economy factors. According to the model (equation 10), 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 the incentive for 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 five columns report results for the short-run model (19) using five different measures of I_i, and the last five columns report results for the long-run model (20).

The estimation results provide no evidence for the effect of the CUSFTA on lobbying for protection: the estimates of ϕ_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 decline in lobbying power of special interest groups and a deeper reduction in the level of protection granted to politically organized industries. The study by Ketterer, Bernhofen, and Milner (2012) also find no effect of the

CUSFTA on lobbying for protection against outsiders, 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.[23] It is important to note that although we do not find support for the hypothesis of the negative effect of the CUSFTA on lobbying for protection, the power of our test can be low due to poor measurement of industrial lobbying activity. This problem, which is common to the political economy of trade literature, can make it difficult to pick up the effect of our interest in noisy data.[24]

Turning to other estimates of equations (19) and (20), they are very similar to our previously reported findings. The estimates for ϕ_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 0.06 percentage points reduction in the MFN tariff in the short run and 0.31 in the long run.[25] 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, ϕ_2, is negative but statistically insignificant in the long-run specification, indicating that the MFN tariffs were not adjusted for domestic industries facing shrinking 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 remain negative but only β_1 is statistically significant. This result provides some support for the cooperative trade policy hypothesis, indicating that Canadian policymakers were more willing to liberalize industries which play the least important role in the US exports. To gauge the importance of this factor for the MFN 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 (6) imply that industries in the first three quintile of the US export share distribution experienced a respective 4.53, 0.92, and 0.31 percentage point reduction in MFN tariffs in addition to the average 2.21 percentage reduction in industries in the fifth quintile. Since we do not find statistically significant variation in the rate of tariff reduction across industries in the last fourth quintiles of the US export share distribution, deeper tariff cuts in the first quintile provide only partial support to 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 $0.3 - 0.35$ percentage points for an average industry. At the same time, we found only limited support for the trade policies of CUSFTA member countries to be formed cooperatively. While Canada provides less protection to industries with the least imports from the US, changes in the MFN tariff rates are not systematically related to export rents generated by those industries to US exporters. Finally, we do not find any effect of the CUSFTA on the intensity of industrial lobbying for trade protectionism.

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 ϕ_1 and β_1 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 again 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.[26]. For this subsample, industries with the least imports from the US experienced additional $5.1 - 5.3$ percentage points decline in the MFN tariff relative to industries in the top three quintiles of the US import share distribution.

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 market shares of US exporters.

As a final robustness test, we exclude the pre-CUSFTA import tariff from the long-run model. Since CUSFTA member countries had committed to a complete elimination of import tariffs by the year 1998, the change in preferential tariff between 1989 and 1998 is highly correlated with the starting value of import tariff, which may cause a high degree of multicollinearity between the 1988 tariff rate and the X^1_{it} variable. The results in column (9) reveal a marginal increase in ϕ_1 coefficient, indicating that multicollinearity is unlikely to be a serious problem.

4 Conclusions

Whether FTAs induce or deter the incentive of member countries for multilateral trade liberalization has been a central question in the regionalism 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 rises and falls in the level of protectionism following the formation of an agreement. Furthermore, the empirical evidence on the relationship between FTA membership and an MTL is inconclusive: 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 developed 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 did in fact facilitate a greater liberalization of 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 leads to a $0.3 - 0.35$ 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 a 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.

In this study we failed to find any consistent evidence on the negative impact of the CUSFTA on the lobbying power of domestic special interest groups. Despite the theoretical prediction that intensified competition with US firms and declining domestic market share should have had a 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. Yet, it is hard to obtain reliable measures of lobbying intensity, and this finding can be derived by the poor quality of our political economy variables.

Our study also provides weak evidence on the presence of trade policy cooperation between Canada and the US. We show that industries that generate less export revenue for the US had experienced deeper tariff cuts during the Uruguay round. This result is consistent with the hypothesis that the Canadian government is reluctant to erode the rents of US exporters generated by their preferential treatment. However, contrary to the tariff cooperation hypothesis, we failed to find any relationship between changes in MFN tariffs and US export rents among industries with large exports to Canada. Overall, the dominance of a tariff complementarity effect of the CUSFTA suggests that the main purpose of the agreement was to exchange market access between the two countries.

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Notes

[1] On the other hand, 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.

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

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

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

[5] See Technical Appendix for complete derivation.

[6] See Technical Appendix for complete derivation.

[7] It is important to note that under the assumption that all industries are politically organized ($I_i = 1$ for all i), used in most of the political economy of trade literature, equation (10) becomes

$$\frac{\sigma_i}{\sigma_i - 1}\epsilon_i t_{it}^F = \sigma_i s_{it}^P t_{it}^P + \frac{a+1}{a+\alpha} s_{it}^H + \frac{b}{a+\alpha} s_{iP}$$

and the effect of lobbying cannot be identified separately from the market structure effect in the regression analysis.

[8] We exclude σ_i from X_{1t} in the estimation equation to reduce the chance of having a measurement error.

[9] 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 ϕ_1 and ϕ_2 on both subsamples. In the Section 3.5 we report results estimated from the two subsamples separately.

[10] It is important to note that the two other three effects of FTA on external tariffs studies in this paper are negative and can thus be identified in the presence of ceiling tariff bindings.

[11] 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.

[12] It is important to note that while the empirical specification (17) is different from Limao (2006) and Karacaovali and Limao

(2008), introduction of the regional public good in our theoretical model will lead to the same prediction that industries with more imports from the US should experience smaller MFN tariff reductions. 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 effect of tariff preferences on public good provision is positive only for the corner case when the preferential tariff rate 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 effect of tariff preferences on the public good provision by the partner will be increasing in the US revenue from exporting to Canada. Derivations are available upon request.

[13] 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.

[14] 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: $RSCA_{it} = \frac{RCA_{it}-1}{RCA_{it}+1} \in [-1;1]$.

[15] Bown and Crowley (2007 JIE), however, document positive effect of US antidumping duties against China on Chinese exports to other countries. Therefore, in our empirical analysis we pay close attention to the validity of the exclusion restriction tests.

[16] 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.

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

[18] 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.

[19] Since we do not model endogenous lobby formation, we do not allow industrial political activity to vary over time. Hence, in construction of the third and fourth measures of political organization we use the data for 1988 and then extrapolate the results for the rest of the sample. Focusing on the lobbying stucture in the year leading up the the CUSFTA formation has the advantage that these measures of lobbying activity are unlikely to be driven by factors related to the agreement.

[20] The Angrist-Pischke first stage F-test always rejects the null of weak instruments for all endogenous variables in all specifications at 99% confidence level. We also cannot reject the hypothesis of exogeneity of instruments, suggesting that our instruments are overall of a good quality.

[21] 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 $\frac{0.0139}{3.5} \simeq 0.004$.

[22] Moreover, the Hausman endogeneity test fails to reject the null hypothesis of endogeneity of D_{it}^k variables and thus we cannot reject the consistency of the OLS estimates in column (3).

[23] If we also assume that all industries are politically oranized, as in Ketterer, Bernhofen, and Milner (2012) and Karacaovali and

Limao (2008), specification (20) becomes

$$\Delta_9 Y_{i,98} = \alpha + \beta_0 D_i + \phi_1 \Delta_9 X^1_{i,1998} + (\phi_2+\phi_3)\, \Delta_9 X^2_{i,1998} + u_i$$

and the effects of lobbying and market structure become separately unidentifiable. Under this assumption IV results in Table 5 demonstrate that $(\phi_2+\phi_3)$ is not statistically different from zero and do not support the hypothesis of the effect of the CUSFTA on lobbying for protection.

[24] Another source of inconsistency with the theory can be the static nature of the PFS model and the long-run equilibrium analysis and may not be well suited to describe 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.

[25] 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 $0.04 \cdot \phi_1$.

[26] An increase in the ϕ_1 coefficient is largely offset by a reduction in Δs^P_{it} in the estimation sample.

Table 1. Summary statistics for key variables

	Mean	Standard deviation	Minimum	Maximum	Number of observations
MFN Tariff	0.057	0.064	0	0.249	38,445
Δ(MFN Tariff)	-0.004	0.013	-0.068	0.052	38,445
Preferential tariff	0.025	0.04	0	0.222	38,445
Δ(Preferential tariff)	-0.007	0.011	-0.054	0.031	38,445
Canadian market share	0.578	0.196	0.066	0.999	41,204
Δ(Canadian market share)	-0.011	0.032	-0.445	0.445	40,779
US market share	0.253	0.144	0.001	0.797	41,190
Δ(US market share)	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
I5	0.520	0.500	0	1	5020
Import demand elasticity	-2.958	4.906	-37.979	-0.213	4,018
Log of firm scale	16.143	1.286	13.593	21.805	243
Material share	0.511	0.118	0.164	0.898	243
Labor share	0.191	0.073	0.015	0.37	243
Non-prod. Labor share	0.202	0.089	0.057	0.594	243
Fuel and electricity 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 take the value of one 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). I5 is equal to one for industries which received the most protection during the CUSFTA grace period. Import demand elasticities are obtained from Kee, Nicita, and Olarreaga (2009).

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

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	OLS	OLS	OLS	OLS	IV-GMM	IV-GMM	IV-GMM	IV-GMM
$\Delta tariff\_us_{it-1}$	0.1053***			0.1054***			0.0996***	0.1067***
	(6.69)			(6.97)			(6.92)	(6.01)
$\Delta tariff\_us_{it-2}$								0.0540***
								(4.87)
$\Delta tariff\_us_{it-3}$								0.0316***
								(4.59)
US imports indicator (D) †		0.0009		0.0019	-0.0067		0.0358	0.1816**
		(0.73)		(1.29)	(-1.30)		(0.90)	(2.22)
Exp. Shr, quintile 1 (D^1) †			-0.0000	0.0003		-0.0001	0.0008	0.0010
			(-0.04)	(0.83)		(-0.29)	(1.45)	(1.42)
Exp. Shr, quintile 2 (D^2) †			-0.0006	-0.0002		-0.0014***	-0.0010**	-0.0012**
			(-1.43)	(-0.62)		(-2.65)	(-2.01)	(-2.17)
Exp. Shr, quintile 3 (D^3) †			-0.0002	-0.0000		-0.0003	-0.0002	-0.0000
			(-0.71)	(-0.12)		(-0.79)	(-0.57)	(-0.01)
Exp. Shr, quintile 4 (D^4) †			-0.0001	0.0001		-0.0005	-0.0004	-0.0003
			(-0.23)	(0.38)		(-1.14)	(-0.99)	(-0.74)
R-squared	0.050	0.041	0.041	0.050				
Hansen J-statistics, p-val. [(a)]					0.38	0.68	0.50	0.69
Endogeneity test, p-val. [(b)]					0.256	0.000	0.006	0.036
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
L. Δ(US market share*Pref. tariff), (ΔX^1_{it-1}) †	1.682***	1.684**	1.780***	1.231*
	(3.72)	(2.28)	(3.92)	(1.79)
L.ΔCanadian market share, (ΔX^2_{it-1}) †	0.036	-0.208	0.037	-0.154
	(1.09)	(-0.91)	(1.09)	(-1.29)
US imports indicator, (D) †			0.028	-1.874
			(1.06)	(-1.38)
Exp. Shr, quintile 1, (D^1) †			-0.014***	-0.023**
			(-3.12)	(-2.08)
Exp. Shr, quintile 2, (D^2) †			-0.005**	0.006
			(-2.11)	(0.64)
Exp. Shr, quintile 3, (D^3) †			-0.003	-0.010*
			(-1.47)	(-1.71)
Exp. Shr, quintile 4, (D^4) †			-0.003*	0.006
			(-1.72)	(1.51)
R-squared	0.018		0.019	
Hansen J-statistics, p-val. [(a)]		0.45		0.69
Endogeneity test, p-val. [(b)]		0.194		0.237
N	25,193	14,035	25,193	14,035

Notes: The dependent variable is the elasticity-adjusted 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 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
$\Delta tariff\_us_{it}$	0.120***			0.201***			0.205***
	(6.34)			(5.94)			(6.14)
US imports indicator, (D) †		-0.008	-0.002		-13.119	0.076	
		(-0.72)	(-0.15)		(-1.01)	(0.32)	
Exp. Shr, quintile 1, (D^1) †			0.010***	0.010***		0.020***	0.019***
			(4.03)	(4.20)		(5.56)	(6.53)
Exp. Shr, quintile 2, (D^2) †			0.005***	0.005***		0.002	0.005
			(2.66)	(2.88)		(0.41)	(1.01)
Exp. Shr, quintile 3, (D^3) †			0.005***	0.005***		0.003	0.001
			(3.07)	(3.29)		(0.56)	(0.34)
Exp. Shr, quintile 4, (D^4) †			0.003*	0.003**		0.002	0.002
			(1.75)	(2.01)		(0.54)	(0.57)
R-squared	0.231	0.216	0.225	0.24			
Hansen J-statistics, p-val. [(a)]					0.00	0.22	0.24
Endogeneity test, p-val. [(b)]					0.365	0.016	0.020
N	3,864	3,887	3,887	3,864	3,785	3,785	3,764

Notes: The dependent variable is the change in the MFN tariff between the years 1998 and 1989. * 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
Δ(US market share*Pref. tariff), (ΔX_{it}^{1}) †	3.947***	7.577***	4.011***	7.821***
	(3.63)	(3.68)	(3.89)	(3.61)
ΔCanadian market share, (ΔX_{it}^{2}) †	0.078***	0.003	0.125***	0.008
	(2.68)	(0.01)	(3.25)	(0.06)
Exp. Shr, quintile 1, (D^{1}) †			-0.152***	-0.160**
			(-3.38)	(-2.02)
Exp. Shr, quintile 2, (D^{2}) †			-0.076***	-0.028
			(-3.25)	(-0.29)
Exp. Shr, quintile 3, (D^{3}) †			-0.035*	-0.009
			(-1.77)	(-0.18)
Exp. Shr, quintile 4, (D^{4}) †			-0.021*	0.038
			(-1.93)	(0.96)
R-squared	0.076		0.093	
Hansen J-statistics, p-val. [(a)]		0.73		0.54
Endogeneity test, p-val. [(b)]		0.027		0.003
N	3,178	2,315	3,178	2,315

Notes: The dependent variable is the elasticity-adjusted change in the MFN tariff between the years 1998 and 1989. * 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)	(9)	(10)
Model:			Short run					Long run		
Δ(US market share*Pref. tariff), (ΔX_{it}^{1}) †	1.57**	1.65**	1.52**	1.73***	1.49**	7.80***	7.79***	7.71***	7.86***	7.36***
	(2.40)	(2.55)	(2.30)	(2.62)	(2.25)	(3.57)	(3.61)	(3.58)	(3.66)	(2.82)
ΔCanadian market share, (ΔX_{it}^{2}) †	-0.079	-0.084	-0.163	0.214	-0.227*	0.036	-0.022	-0.028	0.020	-0.099
	(-0.52)	(-0.85)	(-0.96)	(0.67)	(-1.67)	(0.17)	(-0.14)	(-0.17)	(0.13)	(-0.34)
Δ(Canadian market share)*I, (ΔX_{it}^{3}) †	-0.074	-0.382	-0.095	-0.412	0.328	-0.043	0.101	0.065	-0.029	0.185
	(-0.36)	(-1.40)	(-0.46)	(-1.42)	(0.92)	(-0.21)	(0.50)	(0.40)	(-0.17)	(0.49)
Exp. Shr, quintile 1, (D^{1}) †	-0.008	-0.009	-0.009	-0.010	-0.008	-0.159**	-0.166**	-0.154**	-0.157*	-0.151*
	(-1.09)	(-1.28)	(-1.10)	(-1.46)	(-1.10)	(-2.02)	(-2.11)	(-1.97)	(-1.93)	(-1.75)
Exp. Shr, quintile 2, (D^{2}) †	-0.004	-0.001	-0.004	-0.002	-0.003	-0.032	-0.024	-0.016	-0.029	-0.033
	(-0.37)	(-0.06)	(-0.37)	(-0.21)	(-0.31)	(-0.33)	(-0.25)	(-0.17)	(-0.31)	(-0.35)
Exp. Shr, quintile 3, (D^{3}) †	-0.005	-0.006	-0.005	-0.005	-0.004	-0.011	-0.007	-0.005	-0.009	0.008
	(-0.89)	(-0.97)	(-0.98)	(-0.84)	(-0.79)	(-0.21)	(-0.13)	(-0.10)	(-0.17)	(0.12)
Exp. Shr, quintile 4, (D^{4}) †	0.004	0.005	0.004	0.004	0.004	0.036	0.037	0.041	0.038	0.043
	(0.99)	(1.20)	(0.93)	(0.97)	(0.92)	(0.85)	(0.95)	(1.03)	(0.98)	(1.00)
Hansen J-statistics, p-val. (a)	0.45	0.51	0.53	0.61	0.59	0.56	0.64	0.58	0.51	0.73
Endogeneity test, p-val. (b)	0.367	0.341	0.248	0.374	0.200	0.005	0.003	0.003	0.005	0.002
N	14,035	14,035	14,035	14,035	14,035	2,315	2,315	2,315	2,315	2,315

Notes: The dependent variable is the elasticity-adjusted 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 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
Δ(US market share*Pref. tariff), $(\Delta X_{it}^{1})^{\dagger}$	3.031**	4.828**	1.724***	7.769***	2.495***	11.188**	1.565**	7.796***	9.341***
	(2.32)	(2.07)	(2.65)	(3.03)	(3.04)	(2.38)	(2.40)	(3.57)	(4.80)
ΔCanadian market share, $(\Delta X_{it}^{2})^{\dagger}$	-0.949	-0.213	0.019	0.050	0.076	0.164	-0.079	0.036	0.017
	(-0.44)	(-1.08)	(0.07)	(0.20)	(0.16)	(0.49)	(-0.52)	(0.17)	(0.08)
Δ(Canadian market share)*I, $(\Delta X_{it}^{3})^{\dagger}$	-0.246	0.282	-0.309	-0.095	-1.106*	-0.372	-0.074	-0.043	-0.055
	(-0.21)	(1.24)	(-1.23)	(-0.36)	(-1.67)	(-0.75)	(-0.36)	(-0.21)	(-0.26)
Exp. Shr, quintile 1, $(D^{1})^{\dagger}$	-0.150*	-0.110	-0.009	-0.174**	-0.017	-0.186*	-0.008	-0.159**	-0.162**
	(-1.68)	(-1.35)	(-1.05)	(-2.29)	(-1.58)	(-1.82)	(-1.09)	(-2.02)	(-1.99)
Exp. Shr, quintile 2, $(D^{2})^{\dagger}$	0.016	-0.103	-0.007	-0.073	0.006	-0.097	-0.004	-0.032	-0.040
	(0.26)	(-1.03)	(-0.65)	(-0.68)	(0.62)	(-0.77)	(-0.37)	(-0.33)	(-0.41)
Exp. Shr, quintile 3, $(D^{3})^{\dagger}$	0.111**	0.016	-0.004	-0.008	-0.012	0.004	-0.005	-0.011	-0.001
	(2.00)	(0.29)	(-0.65)	(-0.13)	(-1.13)	(0.04)	(-0.89)	(-0.21)	(-0.03)
Exp. Shr, quintile 4, $(D^{4})^{\dagger}$	-0.055	0.009	0.005	0.032	0.001	0.006	0.004	0.036	0.038
	(-0.93)	(0.21)	(1.25)	(0.64)	(0.10)	(0.09)	(0.99)	(0.85)	(0.93)
Condition	2-digit HS 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-val [(a)]			0.72	0.49	0.83	0.40	0.45	0.56	0.58
Endogeneity test, p-val [(b)]			0.466	0.024	0.302	0.216	0.331	0.025	0.002
N	14,035	2,315	11,378	1,895	6,671	1,104	14,035	2,315	2,317

Notes: The dependent variable is the elasticity-adjusted 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. 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.