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Trade Policy Determinants and Trade Reform in a Developing Country

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Abstract

In this paper, I start out with a standard political economy of trade policy model to measure the determinants of trade policy in a developing country. I carefully test the model empirically with Colombian data from 1983 to 1998 accounting for endogeneity and omitted variable bias concerns and then expand it in several directions.. I show that it is important to control for the impact of a drastic trade reform shock that affects all sectors and disentangle its effect from preferential trade agreements (PTAs). I find that protection is higher in sectors that are important exports for preferential partners which may be seen as a stumbling effect of PTAs for Colombia. I also relax the assumption of fixed political weights that measure the extra importance of producers' welfare relative to consumers in the government objective. I measure the impact of sectoral characteristics on tariffs indirectly through political weights as a novel alternative to nonstructurally estimating them as determinants of protection. Accordingly, I obtain more realistic estimates for the political weights further contributing to the literature.

JEL Classification: F13, F14, F15.

Keywords: Political economy of trade policy, trade liberalization, preferential trade agreements, empirical trade.

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1 Introduction

Although according to trade theory the optimal trade policy for a small open economy is free trade, in reality, trade protection in small developing nations is higher and more widespread than the rest. Using a standard political economy of trade policy model, I start out by showing that protection in a small open economy will be inversely related to import penetration (imports/domestic production) and import demand elasticity which is a common result in several different models (Findlay and Wellisz 1982; Hillman 1982; Mayer 1984; Grossman and Helpman 1994; and so on). Then, I use this parsimonious model to estimate the 4-digit industry level (ISIC) tariff rates in Colombia from 1983 to 1998 and confirm the prediction of the model which is consistent with the evidence in the empirical literature such as Goldberg and Maggi (1999) for US, Mitra et al. (2002) for Turkey, McCalman (2004) for Australia, and Karacaovali and Limão (2008) for EU among others.

Next, I expand the benchmark model in various directions. First, I empirically model a unilateral trade liberalization shock which affects all sectors to capture the Colombian experience of drastic trade reform in the early 1990s similar to Chile, Mexico, Turkey, and India. I model this common shock with an overall reduction in the additional political weight the government places on producers' welfare (who lobby for protection) relative to the welfare of average citizens (who are unorganized).

Second, I relax the assumption of fixed political economy weights attributed to producers and allow them to vary based on three sectoral characteristics that might mark up or discount these weights: 1) Share of employment in a sector, 2) firm share as a proxy for concentration, and 3) labor to output ratio as a proxy for labor intensity. I rely on a short list of variables that were identified to affect trade policy in the earlier literature such as in Baldwin (1985), Trefler (1993), and Gawande (1998). The novelty of the estimation approach in this paper is that rather than assuming a nonstructural relationship between protection and these variables, I empirically model them as factors directly influencing cross-industry political weights (and hence indirectly affecting protection). I find that political weights are discounted for sectors with higher share of employment while they are marked up for labor intensive and concentrated sectors in Colombia. I also obtain more realistic estimates for the political weights by allowing them to vary across sectors over time.

Third, I consider the impact of the preferential/regional agreements by controlling for the sectoral share of imports from preferential partners and find that the protection is higher for sectors with higher share of preferential imports from the Andean Group. This evidence supports the findings in Limão (2006) for the US and Karacaovali and Limão (2008) for the EU who identify a slowing down effect of preferential trade agreements (PTAs) on multilateral tariffs. This finding is in contrast with Bohara et al. (2004) for Argentina and Estevadeordal et al. (2008) for ten Latin American countries. However, given that the proliferation of PTAs and the rise in their intensity coincide with a period of much unilateral trade liberalization in these economies, accounting for trade reform as I do in this paper becomes very important.

Finally, I carefully address the potential endogeneity issues in the econometric model using an instrumental variables approach and perform several robustness checks.

The paper is organized as follows. In the next section, I present the basic theoretical framework that guides the estimations and then in Section 3, I develop the econometric model and offer the extensions to the theoretical setup as well as discuss specification issues. In Section 4, I describe the data and present the estimation results and robustness checks. Section 5 concludes.

2 Theoretical Framework

I rely on a standard political economy model of trade policy that can be interpreted as the reduced form of a model where special interest politics is given micro-foundations like in Grossman and Helpman (1994). I assume a small open economy where output and factor markets are perfectly competitive. The numeraire good $i = 0$ is produced with labor only,

$X_0(p_0) = L_0$, whereas the other goods, $X_i(p_i)$ for $i = 1, ..., n$, are produced with labor and a sector specific factor (that is immobile across sectors). The population and world prices of all goods are normalized to one, $p_i^w = 1 \ \forall i$, and the numeraire good is traded freely. Therefore, the wage rate also equals one given a competitive labor market and assuming there is enough labor for the numeraire good to be always produced in equilibrium.

While consumers fail to overcome the collective action problem and organize for free trade (Olson 1965), specific factor owners that constitute a negligible share of the population get organized and lobby for protection in their own sector. Tariffs are assumed to be the only form of protection for simplicity so the domestic price of nonnumeraire goods is $p_i = 1 + \tau_i$, where τ_i stands for both advalorem and specific tariff rates.[1]

The government determines tariffs by maximizing the following political support function

$$G(p) = \sum_{i=1}^{N} \left(\int_{1+\tau_i}^{\infty} D_i(\tau_i) d\tau_i + (\pi + 1) \int_{0}^{1+\tau_i} X_i(\tau_i) d\tau_i + \tau_i M_i(\tau_i) \right) \tag{1}$$

which is a weighted sum of aggregate consumer and producer surplus as well as tariff revenue. $D_i(\tau_i)$ denotes aggregate demand, $X_i(\tau_i)$ denotes aggregate supply, and $M_i(\tau_i) = D_i(\tau_i) - X_i(\tau_i)$ is the aggregate import demand. Assuming away wasteful government expenditures, the tariff revenue, $\sum_i^N \tau_i M_i(.)$, is rebated back to public in its entirety. $\pi > 0$ measures the additional political weight the government places on the welfare of specific factor owner lobbies relative to an average voter. In the absence of the political weight, $\pi = 0$, equation (1) boils down to a standard social welfare function without lobbying.

Maximizing equation (1) with respect to τ_i we obtain the following first order condition for an interior solution

$$\frac{\partial G}{\partial \tau_i} = -D_i(\tau_i) + (\pi + 1) X_i(\tau_i) + M_i(\tau_i) + \tau_i M_i'(\tau_i) = \pi X_i(\tau_i) + \tau_i M_i'(\tau_i) = 0 \tag{2}$$

[1] This is because world prices are normalized to one. Furthermore, trade is balanced through movements of the numeraire good.

Therefore, the equilibrium advalorem/specific tariff rate for good i is implicitly defined by

$$\tau_i = -\pi \frac{X_i(\tau_i)}{M_i'(\tau_i)} \equiv \pi \frac{X_i(\tau_i)/M_i(\tau_i)}{\varepsilon_i(\tau_i)} \tag{3}$$

where $\varepsilon_i(.)$ stands for the elasticity of import demand.[2] This expression is similar to those obtained in various political economy models as shown in Helpman (1997). The tariff rate for sector i increases in the additional political weight placed on the well-being of producers, π, while decreases in the import demand elasticity, ε_i, and the import penetration ratio, M_i/X_i. A tariff is a tax on imports so the deadweight loss from taxing imports is lower for more inelastic import demand. A relatively larger market for imports creates a greater price distortion potential putting a downward pressure on tariffs, whereas the marginal benefit of a tariff to a producer lobby is higher when it applies to more units.

3 Econometric Model

3.1 The Benchmark

As a benchmark, I first assume that tariffs are determined by equation (3) for sectors $i = 1, ..., N$ and over years $t = 1, ..., T$ which in log linear and error form can be re-expressed as

$$\log \tau_{it} = \alpha + \beta_1 \log \frac{X_{it}/M_{it}}{\varepsilon_{it}} + u_{it} \tag{4}$$

where $\hat{\alpha} = \log \hat{\pi}$. Given the parsimonious nature of the model, to account for other industry specific characteristics that might make tariffs differ across sectors in a systematic way, I then augment this model with industry fixed effects

$$\log \tau_{it} = \alpha + \beta_1 \log \frac{X_{it}/M_{it}}{\varepsilon_{it}} + \theta_i \boldsymbol{\beta}_2 + u_{it} \tag{5}$$

[2] Import demand elasticity is defined as $\varepsilon_i = -M_i' p_i^w / M_i$.

where θ_i is a $1 \times N$ vector of industry dummies[3], u_{it} is the error term, α and β_1 are scalars, and $\boldsymbol{\beta}_2$ is an $N \times 1$ vector of coefficients.

3.2 Trade Reform

I estimate tariffs at the industry level over the 1983 to 1998 period in Colombia which like many other developing countries (e.g. Brazil, Turkey, India, etc.) went through significant unilateral trade liberalization in the early 1990s (see Figure 1). The average tariff rate went from 44% in 1983 down to 14% after the reform and given that there were not any financial crises during this time period that could potentially interfere with the analysis, Colombia provides a natural experiment environment for studying determinants of trade policy and trade reform in a developing country. Import licenses were another common measure used along with tariffs prior to trade reform but these were almost eliminated together with tariff liberalization (Edwards 2001). Therefore, the reduction of tariff protection was not replaced by a new form of protection. Tariff rates are also better measured and they are positively correlated with import licenses. Nevertheless, as a robustness check, I use effective rate of protection (ERP), which is based on value added, as an alternative protection measure in Section 4.4 and show that the results with tariffs hold under ERP.

It is important to account for the common trade reform shock across sectors while I maintain the hypothesis that political economy of trade policy still matters. The variation in tariffs across sectors is expected to depend on political economy forces even in the face of unilateral trade liberalization that prevails in all sectors. In the late 1980s and early 1990s there was a change in the economic consensus where old import substitution policies were abandoned for more liberal trade policies, perhaps with the encouragement of World Bank research and policy dialogue (Edwards 1997). I model this change in view leading to a trade reform with a common decline in the additional political economy weight attributed to producer lobbies, π. Edwards (2001) indicates that César Gaviria (President of Colombia

[3]The i^{th} column of θ_i is 1 and the rest are zeros.

from 1990 to 1994) "developed from early on a critical view regarding CEPAL's [Economic Commission for Latin America] import substitution development strategies." I capture this effect with a period dummy that measures the shift in the intercept starting from 1990

$$\log \tau_{it} = \alpha + \beta_1 \log \frac{X_{it}/M_{it}}{\varepsilon_{it}} + \beta_2 REF_t + \theta_i \boldsymbol{\beta}_3 + \nu_{it} \tag{6}$$

where $REF_t = 1$ for $t \geq 1990$ and zero otherwise, θ_i is a $1 \times N$ vector of industry dummies, and ν_{it} is the error term.[4]

Based on theory, the expected sign for β_1 is positive indicating that tariffs are inversely related to elasticity adjusted import penetration ratio, $(M/X)\varepsilon$. REF_t points to a common decline in tariffs across sectors due to the trade reform put in place in 1990 and onwards so β_2 is expected to be negative. In effect, equation (6) allows the political weight π to differ before and after the reform eras. More specifically,

$$\log \hat{\pi}_t = \begin{cases} \hat{\alpha} \text{ for } t < 1990 \\ \hat{\alpha} + \hat{\beta}_2 \text{ for } t \geq 1990 \end{cases} \tag{7}$$

Finally, the industry dummies account for fixed sectoral characteristics that might explain further cross-industry variation in tariffs that are not already captured by the benchmark model.

3.3 Political Weights

Several other variables have been identified as potential factors affecting protection in the earlier empirical studies (see Baldwin 1985 for example). Based on this observation, it is plausible to argue that political weights may vary from one sector to the other over time (Karacaovali and Limão 2005). I conjecture that the value of contribution from lobbies may be discounted or marked up based on sectoral characteristics. Therefore, I relax the assumption of fixed political economy weights and empirically model them to vary based on

[4] α, β_1, and β_2 are scalars and $\boldsymbol{\beta}_3$ is an $N \times 1$ vector of coefficients.

some alternative industry variables:

$$\log \tau_{it} = \alpha_0 + \alpha_{it} + \beta_1 \log \frac{X_{it}/M_{it}}{\varepsilon_{it}} + \beta_2 REF_t + \theta_i \boldsymbol{\beta}_3 + \upsilon_{it} \tag{8}$$

where α (estimating the effect of the fixed political weight on producer welfare, $\log \pi$) is broken into fixed and variable (across sectors over time) portions: α_0 and $\alpha_{it} = \sum_k \gamma_k Z_{kit}$. Here, Z_{kit} is the k^{th} factor measuring sectoral variation in political weights. Therefore, the political weights are estimated as follows

$$\log \hat{\pi}_{it} = \begin{cases} \hat{\alpha}_0 + \sum_k \hat{\gamma}_k Z_{kit} \text{ for } t < 1990 \\ \hat{\alpha}_0 + \hat{\beta}_2 + \sum_k \hat{\gamma}_k Z_{kit} \text{ for } t \geq 1990 \end{cases} \tag{9}$$

Keeping a parsimonious approach, I focus on $k = 3$ industry level variables for Z_{kit}: 1) Share of employment (ratio of employment in the sector to total employment in the economy); 2) firm share (ratio of total number of firms in the economy to the number of firms in the sector) as a proxy for firm concentration; 3) labor to output ratio as a proxy for labor intensity of a sector.

An industry with a higher share of employment commands more votes and may thus be more likely to be favored by politicians (Caves 1976). However, with more workers, there might also arise a free-rider problem and therefore a weaker organization to demand protection in an industry (Trefler 1993). Consequently, the expected sign of the coefficient on share of employment is ambiguous:$\gamma_1 \lesseqgtr 0$. A higher ratio of total number of firms in the economy to the number of firms in an industry is a proxy for firm concentration. A more concentrated sector indicates a stronger organizational power asking for protection (Olson 1965) so we expect $\gamma_2 > 0$. Finally, labor intensive sectors may be favored based on a social justice motive as they may be impacted more adversely from import competition (Baldwin 1985). Accordingly, we expect $\gamma_3 > 0$.

3.4 Preferential Trade Agreements

Preferential Trade Agreements (PTAs) encompassing both Free Trade Agreements (FTAs) and Customs Unions (CUs) are expected to affect the MFN tariffs that apply to countries outside the PTA. Karacaovali (2010) shows that once a free trade agreement (FTA) is in place and it leads to some trade being diverted away from non-member nations into member nations, external tariffs are expected to decline under an endogenous political economy model of trade policy and FTAs. Bohara et al. (2004) find that "over the period $1991 - 1996$...the increasing penetration of imports from Brazil and the resulting 'decline' of industries in Argentina led...to the lowering of external tariffs in these industries" (p. 85). Estevadeordal et al. (2008) look at ten Latin American countries from 1990 to 2001 and similar to Bohara et al. (2004) find that "preferential tariff reduction in a given sector leads to a reduction in external (MFN) tariff in that sector" (p. 1531). However, Karacaovali and Limão (2008) show that the European Union (EU) has reduced its multilateral tariffs less in products imported duty-free from preferential partners but not in products imported from new EU members. Limão (2006) finds a similar effect for the U.S. Therefore, there is mixed evidence lending support for both the stumbling block and building block (Bhagwati 1991) effects of PTAs on global free trade.

Although it is possible that a PTA may exert a downward pressure on external tariffs as predicted in Karacaovali (2010), there might be cross-industry differences over time in terms of the effect of PTAs. In the spirit of the argument in Limão (2007), we may expect countries to hold back reducing tariffs in sectors that are important for PTA partner countries because each time MFN tariffs are liberalized, the preferential access is eroded. If MFN tariffs were to be eliminated, it would also annihilate the preferential agreements which the countries presumably value in the first place. I control for such an effect of PTAs by including the share of PTA imports relative to total imports in an industry as an additional regressor in equation (8). While REF_t takes care of the common decline in all tariffs due to unilateral liberalization, share of PTA imports capture the stumbling versus building block effects across industries.

I focus on the Andean Group PTA of Colombia originally established by the Cartagena Agreement in 1969 with other founding members Bolivia, Chile, Ecuador, and Peru. Venezuela became a member in 1973 while Chile withdrew in 1976. Andean Group is the second biggest trade bloc in South America after Mercosur and it is the most comprehensive regional/preferential trade agreement Colombia was involved in for the sample period of this study. Colombia was also a member of Latin American Integration Association (LAIA) which was established in 1980 and was limited in scope. Although it was augmented by some further bilateral agreements with Chile, Mexico, and Mercosur countries, none of these agreements provided noteworthy preferential access as compared to Andean Group. Furthermore, the Andean agreement is particularly strengthened to eliminate barriers to virtually all intra-regional trade coinciding with the period of general trade reform in Colombia (World Trade Organization 1996).

After controlling for PTA effects, equation (8) can be modified as

$$\log \tau_{it} = \alpha_0 + \alpha_{it} + \beta_1 \log \frac{X_{it}/M_{it}}{\varepsilon_{it}} + \beta_2 REF_t + \beta_3 ShM\_ANDE_{it} + \theta_i \boldsymbol{\beta}_4 + \upsilon_{it} \tag{10}$$

where ShM_Ande_{it} is the share of imports from Bolivia, Ecuador, Peru, and Venezuela to total imports in industry i, year t.

3.5 Specification Issues

All estimations including the benchmark econometric model are potentially subject to endogeneity given the fact that elasticity adjusted inverse import penetration, $X/M\varepsilon$, (the main right-hand-side variable) is a function of domestic prices, hence tariffs. Therefore, OLS estimation is expected to produce biased results. As a way to get around the problem of endogeneity, I use one period lags of all right-hand-side variables. Although this may alleviate the bias, it would not totally eliminate it given the persistence of the dependent variable (tariffs) over time. Therefore, I consider an Instrumental Variables (IV) approach. While the validity and strength of instruments will be discussed in Section 4.4, here I provide a brief intuition behind the choice of instruments.

First, I use import unit values as a proxy for world prices at the border which are correlated with domestic prices by definition but not tariffs so they are useful to instrument for $X(.)/M(.)\varepsilon(.)$. Second, I use a measure of scale (value added/number of firms) as an instrument for import penetration given that scale is likely to be correlated with fixed costs of entry to an industry, hence affect import penetration. However, scale is an inherent characteristic of a sector and once we account for industry size in the protection equation, its effect is only indirect and it can be correctly excluded from the protection equation as done in Goldberg and Maggi (1999) and Gawande and Bandyopadhyay (2000). Third, I rely on upstream total factor productivity (TFP) to instrument for the TFP of a sector, hence $X(.)$. Productivity in a sector is expected to be affected by the average productivity of upstream sectors but upstream TFP is likely to be independent from sector's own tariffs.

Despite relying on a theoretical model and addressing several factors that might define tariffs, the estimations may still suffer from an omitted variable bias. Therefore, I use industry fixed effects in all specifications while the instrumental variables approach is also expected to reduce such bias. Finally, other econometric concerns are addressed in Section 4.4 after estimation results are discussed next.

4 Empirics

4.1 Data

The data for the estimations cover 1983, 1985, and 1988 through 1998 and the definitions of all the variables used in the empirical analysis are provided in Table A.1. Here, I present the data sources.

Tariff data are obtained from the National Planning Department (DNP) of Colombia at the 8-digit product level (Nabandina code), which are aggregated to the 4-digit ISIC (International Standard Classification, United Nations) industry level by using simple averages.[5] Effective rate of protection (ERP) figures are also available from the same data source which I use to test the sensitivity of the results in Section 4.4.

The main production data (total output, value added, number of employees) are available at the 4-digit ISIC level through UNIDO's Industrial Statistics Database while bilateral and aggregate imports data are from COMTRADE, UN Statistics Division (again at the 4-digit ISIC level).

Import demand elasticity is obtained by combining the structural estimates in Kee et al. (2004) with GDP data from World Bank's World Development Indicators (WDI) and import data from COMTRADE, at the 3-digit ISIC level. An alternative time-invariant import demand elasticity measure (3-digit ISIC) is obtained from Nicita and Olarreaga (2007) as a robustness check. Import unit values (dollars per kilogram) are an average at the 3-digit ISIC level proxying for world prices at the border and are taken from Nicita and Olarreaga (2007) as well.

Scale is measured as value added (UNIDO) divided by the number of firms (Eslava et al. 2004) at the 4-digit ISIC level. Upstream total factor productivity (TFP) measures the weighted average of the TFP[6] (Eslava et al. 2004) of the upstream sectors excluding itself where weights for the share of inputs per upstream industry is obtained from the input-output tables at the 3-digit ISIC level from Nicita and Olarreaga (2001). The union dummy variable which measures labor union activity at the 3-digit ISIC level is obtained

[5] I thank Marcela Eslava for providing this data. Using simple shares is common in other papers as well. Alternatively, one could use production or import shares as weights but such data are not available at the disaggregate level.

[6] TFP is obtained at the firm level using a production function residual approach from a two-stage least squares estimation and then aggregated to the 4-digit ISIC level using production shares as weights (Eslava et al. 2004).

from Quintero (2006) and finally, 4-digit industry level employment hours to compute the labor to output ratio is from Eslava et al. (2004).

4.2 Descriptives

Table 1 lists the average tariff rates and their dispersion across 4-digit industries for the main sample. There is a sharp reduction in the average tariff rates starting in 1990 while the dispersion declines to a lesser extent which can be observed from the coefficients of variation.[7] The same trend can also be observed in Figure 1 which depicts the tariff rates at the 4-digit industry level over time. The trade reform affects all sectors, yet there is cross-industry variation which I conjecture to be attributed to political economy forces based on the econometric model developed in Section 3. Therefore, various specifications of the model are tested formally in the next section. Table A.2 provides descriptive statistics for all the variables used in the estimations.

4.3 Estimation Results

4.3.1 The Benchmark Model and Trade Reform

As discussed in Section 3.5, endogeneity of the main right-hand-side (RHS) variable, elasticity adjusted inverse import penetration ratio $X/M\varepsilon$, is a valid concern so I first confirm that endogeneity is present with a Durbin-Wu-Hausman test. Then, as an initial step to address this concern, I use one-period lags of the right-hand-side variables. However, given the persistence in variables, this will be a weak method to address the endogeneity so I resort to an instrumental variables (IV) approach next. More specifically, I use the two-step efficient generalized method of moments (IV-GMM) estimator which is robust to heteroskedasticity of unknown form due to its use of an optimal weighting matrix (Cragg 1983). I test for

[7] Coefficient of variation (CV) is defined as standard deviation divided by the mean, hence takes into account the differences in the magnitude of average tariffs across the periods.

heteroskedasticity using a Pagan-Hall (1983) test and find it to be a problem so this further justifies the use of an IV-GMM estimator.

Although the estimates will be biased, I present the results from Cragg's heteroskedastic ordinary least squares estimator (HOLS) in the first four columns of Table 2 for comparison with the IV-GMM results in the last four columns. In Columns 1 and 5, I test equation (5) and in columns 2 and 6, I retest the same benchmark specification using one period lags of the main RHS variable and its instruments instead. There is a strong support for the benchmark political economy model, where elasticity adjusted import penetration is found to be inversely related to tariffs at the 1% level of significance. This result is in line with the aforementioned empirical evidence in the literature (e.g. Goldberg and Maggi 1999; Karacaovali and Limão 2008).

Next, I account for the common unilateral trade liberalization reform shock affecting all sectors with the trade reform dummy, REF_t which has a negative and significant coefficient (at the 1% level) as expected. Under the HOLS estimations, the R-squared values are significantly higher when I control for trade reform.[8]

4.3.2 Specification Issues

The results from both HOLS and IV-GMM estimations are consistent in general, albeit with smaller coefficients under HOLS. However, given the presence of endogeneity, IV-GMM is the preferred methodology. Hansen's (1982) test of overidentifying restrictions point that the instruments (import unit values, log of scale, and upstream TFP) are orthogonal to the error term and correctly excluded from the estimated equations. The probability values for the Hansen's J test are reported in the last rows of all tables for IV-GMM specifications.

For instance, in Table 2, column 8, under the preferred benchmark specification with lagged $X/M\varepsilon$ and trade reform dummy, the p-value for Hansen's J test is 0.351 so we fail to reject the validity of instruments. Furthermore, the excluded instruments are jointly

[8] Under the IV-GMM methodology R-squared is not a meaningful measure, hence not reported.

significant and the Kleibergen-Paap (2006) test, which is robust to heteroskedasticity, rejects that the model is underidentified. However, weak identification may be a concern for IV estimations in general (c.f. Baum et al. 2007). For the same benchmark specification, the Cragg-Donald (1993) statistic is 9.284 which lets us reject the presence of weak instruments at the 10% level using Stock and Yogo (2005) critical values. The first-stage regressions are presented in Appendix Table A.3. Finally, the Andersen-Rubin (1949) test (which is robust to the presence of weak instruments) indicates that the endogenous regressor, $X/M\varepsilon$, in the structural equation is significant at the 1% level.

4.3.3 Political Weights

In Table 2, columns 5 and 6, I estimate equation (5) and its variant with the one-period lag of X/M, respectively. As indicated in equation (7), the constant term provides an estimate for the fixed political economy weight, π, which is equal to 0.016 in column 5 and 0.008 in column 6. Similarly, equation (6) estimates are in columns 7 and 8 where I allow the political economy weight, π, to vary before and after the reform capturing the common unilateral trade liberalization shock in all sectors with a drop in π. The estimates for π are 0.104 (0.041) before 1990 and 0.053 (0.030) afterwards in the column 8 (7) specification. This indicates that the government values producer welfare 10% more than an average citizen and this figure goes down to 5% after the trade reform. These estimates, although arguably small (c.f. Gawande and Krishna 2003; and Imai et al. 2009), are significantly higher (when trade reform is accounted for) than the ones in the earlier literature. For example, Goldberg and Maggi (1999) estimate π to be 0.014 and Gawande and Bandyopadhyay (2000) 0.0003 for the US, whereas Mitra et al.'s (2002) estimates for Turkey range between 0.010 and 0.013.

It is actually plausible to think that the industry fixed effects in the benchmark model will partially capture the sectoral variation in political weights which is fixed over time. If we assume that the industry effects fully capture such variation, the estimates for political weights in equation (6) can be obtained as follows

$$\log \hat{\pi}_{it} = \begin{cases} \hat{\alpha} + \hat{\beta}_{i3} \text{ for } t < 1990 \\ \hat{\alpha} + \hat{\beta}_2 + \hat{\beta}_{i3} \text{ for } t \geq 1990 \end{cases} \tag{11}$$

where $\hat{\beta}_{i3}$ is the i^{th} row of $\hat{\boldsymbol{\beta}}_3$, the $N \times 1$ vector of coefficients for industry dummies θ_i. The estimates in this framework average 0.177 before 1990 and 0.092 afterwards. However, these estimates should be viewed with caution since industry fixed effects control for alternative determinants of cross-industry variation in tariffs. Therefore, we now turn to the estimates for equation (8) where the political weights are specifically designed to vary across sectors over time based on three sectoral characteristics: 1) Share of employment, 2) firm concentration, and 3) labor intensity.

In the first three columns of Table 3, each variable is first considered one at a time. We see that the political weights for concentrated and labor intensive sectors are marked up while they are discounted for sectors with a high share of employment. As discussed in Section 3.3, more concentrated sectors will have stronger producer lobbies demanding protection so the political weights and hence protection is higher in them. Labor intensive sectors are more adversely affected from increasing import competition so they are given a higher weight and protected more.[9] The expected effect of employment share is ambiguous, since more workers have a bigger voting power, yet they might find it more difficult to get organized. In Colombia, sectors with a lower share of employment have a bigger weight. All estimates are significant at the 1% level in the first three columns.

Although the negative coefficient for share of employment may seem counterintuitive, it is highly and negatively correlated with firm concentration so it might indeed be capturing the lack of organizational power in a sector. As a matter of fact, when we use both variables in the same specification (column 5), firm concentration becomes insignificant. In column 4, I control for the presence of union activity at the 3-digit level along with share of employment

[9]This result is similar to Yotov (2010) in spirit who finds that in the U.S. politicians attach a four times higher weight on trade-affected workers.

and find a negative effect of unionization as well. This finding is actually similar to Matschke and Sherlund (2006) who theoretically and empirically show that trade protection is lower when capital owners lobby irregardless of trade union activity but protection is higher when unions lobby while capital owners do not. In columns 6 and 7 we have the remaining two combinations and in column 8 we have all three variables plus the union dummy where the results are similar to the previous ones in columns 1 through 5.

Applying equation (9), we can estimate the political weights that vary across sectors over time controlling for all three variables. The average political weight estimate before 1990 is 0.246 which decreases to 0.115 afterwards. In Figure 2, the variation in political weights is illustrated over the sample period. The cross-industry variation is noteworthy. In Table A.4, I list the highest five and lowest five sectors in terms of political weights along with the average tariff rates at the 4-digit sector level. Prior to 1990, manufacture of musical instruments sector (ISIC 3902) has a political weight of 0.708, whereas manufacture of drugs and medicines has 0.145. Furthermore, the variation in these weights is not correlated with the variation in tariff rates indicating that we do indeed capture the effect of sectoral variables on tariffs indirectly through the political weights.

Mitra et al. (2006) in their search for more realistic estimates for political weights find that their estimates of π range between 0.02 and 0.03 when they assume 10% of the population is organized, and they range between 0.21 and 0.42 when they assume 90% of the population is organized (Table 2, p. 201). In this respect, my approach in this paper not only provides a plausible alternative to fixed political weights but also produces more realistic estimates for them as compared to earlier studies in the literature.

4.3.4 Preferential Trade Agreements

In Table 4 column 1, I present direct estimates of equation (8) with an additional variable: (Lag of) Share of Imports from Andean Group, L.ShM_ANDE_{it}. I find that in sectors where Andean Group exports relative to total exports to Colombia are higher, there is more protection. This result indicates that Colombia tries to slow down the erosion of preferences in sectors that matter for its preferential partners despite a common unilateral trade liberalization shock that affects all sectors. In order to test whether this relationship holds after the trade reform, I interact the Andean export share variable with the reform period dummy in columns 2 and 3. I find that although there is no significant incremental effect after the trade reform (column 2), the relationship still holds (column 3). This is in line with the fact that preferences were deepened around the same time as the trade reform (World Trade Organization 1996).

One may suspect whether ShM_ANDE_{it} could be endogenous to tariffs. The fact that I use the lag of ShM_ANDE_{it} should alleviate such a potential problem. However, I also specifically test the exogeneity/orthogonality of this variable with a C-test (Baum et al. 2007) and confirm that it is not endogenous. Furthermore, there does not appear to be a correlation between tariffs and this share so its impact on tariffs can be estimated maintaining the assumption of its orthogonality to the error term.

In columns 4 through 6, I estimate variants of equation (10) where I allow political weights to vary across sectors over time as in the previous subsection. The results are the same and the coefficient on ShM_ANDE_{it} remains positive and significant at the 1% level in all specifications.

The results support the stumbling block findings for the US (Limão 2006) and EU (Karacaovali and Limão 2008). However, given that they contrast the building block findings for Argentina (Bohara et al. 2004) and ten Latin American countries (Estevadeordal et al. 2008), my results point to the importance of explicitly accounting for the impact of trade reform as well as political economy factors on trade policy.

4.4 Robustness

In Table 5, I present the results for different robustness checks. In column 1, I use effective rates of protection (ERP) as opposed to tariffs as the dependent variable and find that the results are consistent. ERP measures protection on value added by considering the effect of tariffs on inputs as well. Since ERP was computed by National Planning Department (DNP) and I do not have access to its computation procedure, I use this measure only to check sensitivity.

In column 2, I check the robustness of the results to the use of year effects instead of the trade reform period dummy and show that results are not affected qualitatively. However, for purposes of this study, it is important to directly account for the effect of trade reform which I modeled as a common shock affecting all sectors due to a change in the government perception about the value of import substitution policies, as discussed in Section 3.2.

In column 3, I use an alternative time-invariant import demand elasticity measure from Nicita and Olarreaga (2007) and in column 4, I apply an errors-in-variables correction to this measure following Gawande and Bandyopadhyay (2000) given that elasticity is a generated regressor and may be mismeasured. We see that the results are robust to using these alternative measures and the IV-GMM approach should further alleviate the measurement problem. Therefore, my original time-varying import demand elasticity measure is the preferred one.

Finally, tariff rates in general are censored from below given that they cannot be negative so in column 5, I test the robustness of the results to IV-GMM procedure by considering Newey's two-step tobit estimator (IV-Tobit) instead. The results are not sensitive to using IV-Tobit and also given the fact that all tariff rates are actually positive both before and after the trade reform in Colombia, I do not expect the potential censoring from below to be a problem for my data set.

5 Conclusion

Based on a standard political economy of trade policy model, tariff rates are expected to be inversely related to elasticity adjusted import penetration ratio in a small open economy. I confirm this finding for Colombia and also expand the benchmark model in several directions.

First, I model a common trade reform shock that affects all sectors in a developing economy and leads to a drastic trade liberalization episode. The experience in Colombia is similar to many other developing countries that have abandoned import-substitution policies and structurally high protection rates across the board in the late 1980s and early 1990s. Second, I empirically model political weights that measure the extra importance of the well-being of producers (who are organized) relative to average voters (who are unorganized). This weight may decrease over time when modeled as a common shock where the government objective changes and the usefulness of protection as a development policy is downplayed. I relax this structure further by allowing the political weights to vary across sectors over time and find that they are marked up for concentrated and labor intensive sectors while discounted for sectors with a high share of employment in the economy. The novelty of the approach in this paper is that I capture the effect of these sectoral variables on tariffs indirectly through the political weights different from the estimations in the earlier literature. Third, I account for the effect of preferential trade agreements (PTAs) on tariffs by controlling for the export share of the Andean Group countries to Colombia and find that protection is higher in sectors which are important for the preferential partners. This is in line with the stumbling block rationale such that erosion of preferential benefits will be slowed down because the elimination of preferences would mean the end of the PTA itself.

Given that Colombia experienced a substantial trade reform on a unilateral basis, it is important to disentangle the effect of PTAs from the effect of unilateral trade liberalization by explicitly accounting for it. Moreover, one way to solve the puzzle of unrealistically low estimates of the political weights in the literature may be to relax the assumption of fixed weights and allow them to vary based on sectoral characteristics as I do in this paper.

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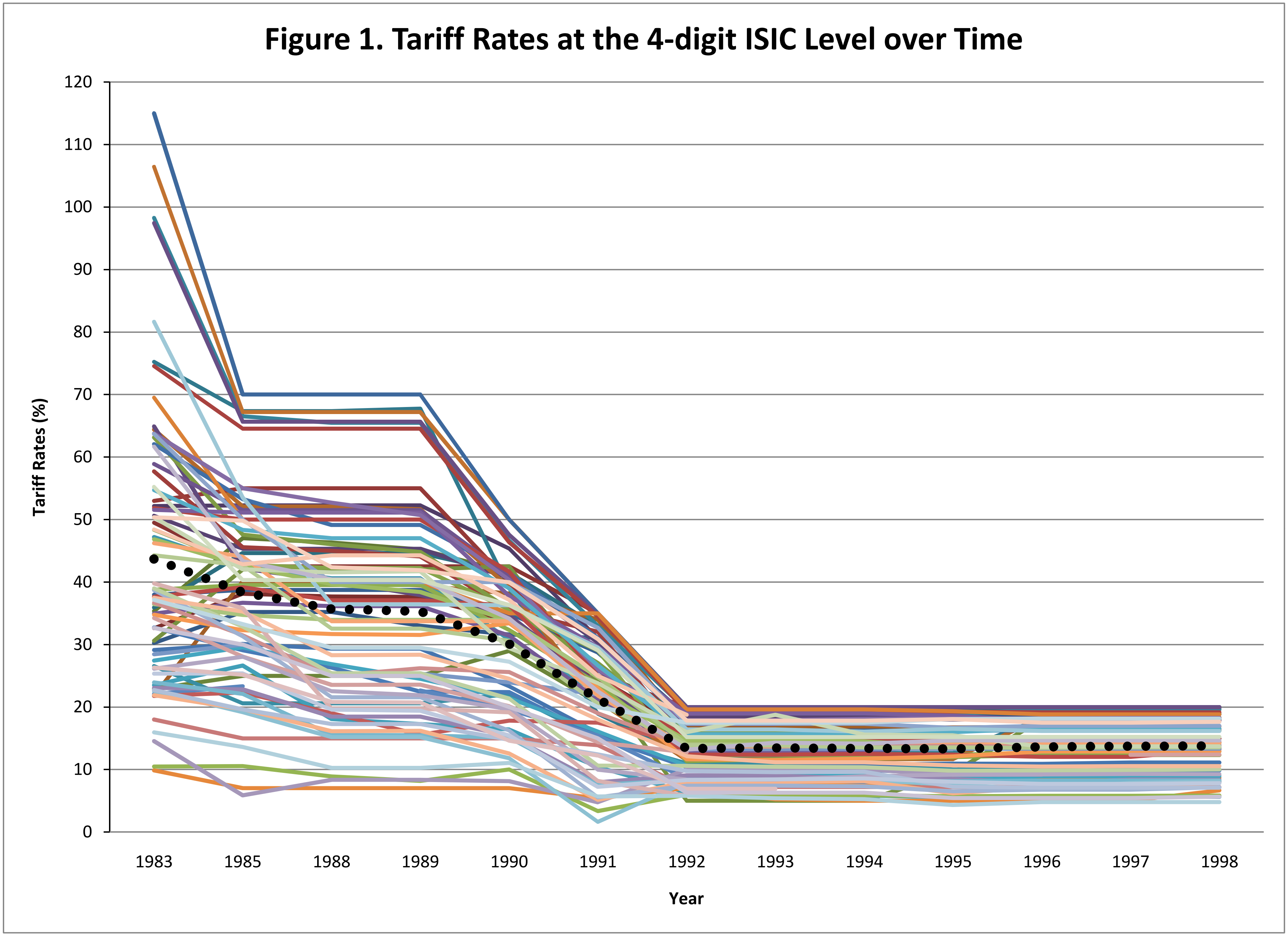
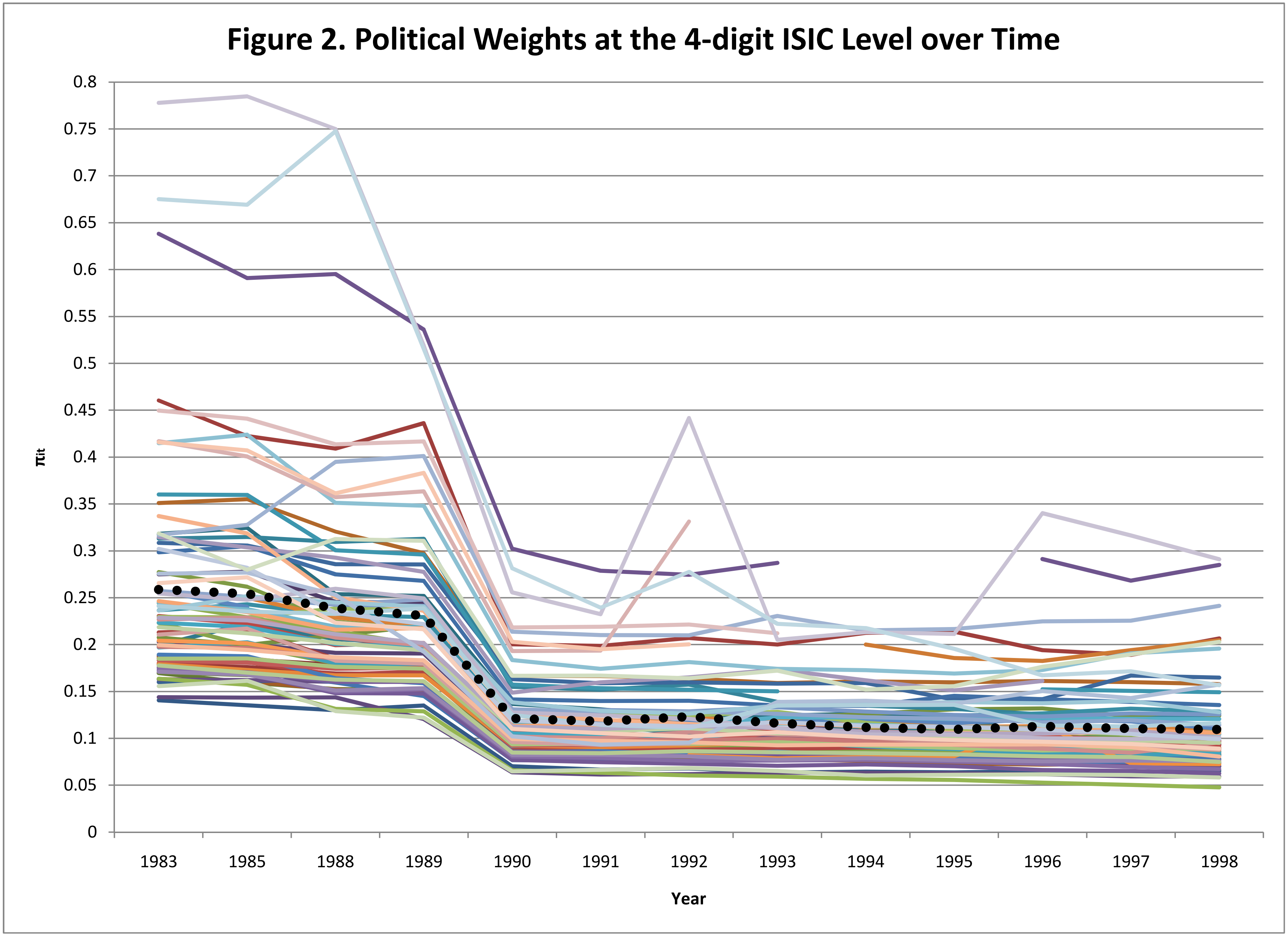


Note: The dashed line depicts average tariff rates over time.
Source: DNP, Colombia



Note: The dashed line depicts the average political weights over time.

Source: Author's own calculations

Table 1. Descriptive Statistics for 4-digit ISIC Level Tariffs over Time

Year	Observations	Mean	Standard Deviation	Coefficient of Variation	Minimum	Maximum
1983	75	43.698	21.814	0.499	9.865	115
1985	75	38.355	14.513	0.378	5.890	70
1988	74	35.691	15.312	0.429	7.017	70
1989	76	35.308	15.281	0.433	7.017	70
1990	75	30.129	11.261	0.374	7.017	50
1991	75	21.315	9.280	0.435	1.646	35
1992	77	13.371	4.496	0.336	5.000	20
1993	73	13.456	4.576	0.340	5.000	20
1994	71	13.396	4.553	0.340	5.000	20
1995	71	13.312	4.670	0.351	4.324	20
1996	73	13.649	4.657	0.341	4.783	20
1997	73	13.726	4.593	0.335	4.783	20
1998	73	13.789	4.535	0.329	4.783	20

Table 2. The Benchmark Model and Trade Reform

	(1) HOLS[a]	(2) HOLS[a]	(3) HOLS[a]	(4) HOLS[a]	(5) IV-GMM[b]	(6) IV-GMM[b]	(7) IV-GMM[b]	(8) IV-GMM[b]
Log($X/M*\varepsilon$)	0.342***		0.099***		1.013***		0.732***	
	(0.018)		(0.014)		(0.091)		(0.172)	
L.Log($X/M*\varepsilon$)		0.284***		0.095***		1.250***		0.461***
		(0.017)		(0.012)		(0.164)		(0.112)
REF			-0.830***	-0.855***			-0.320**	-0.661***
			(0.020)	(0.018)			(0.132)	(0.052)
Constant	-2.401***	-2.261***	-1.194***	-1.176***	-4.163***	-4.862***	-3.195***	-2.268***
	(0.097)	(0.104)	(0.068)	(0.067)	(0.248)	(0.454)	(0.546)	(0.336)
Observations	960	961	960	961	960	960	960	960
R-squared	0.539	0.469	0.814	0.816	n/a	n/a	n/a	n/a
Hansen's J p[c]	n/a	n/a	n/a	n/a	0.597	0.544	0.258	0.351

Notes:

(1) Robust standard errors in parentheses.

(2) *** p<0.01, ** p<0.05, * p<0.10.

(3) "L." stands for one-period (year) lag.

(4) All specifications include industry fixed effects that are jointly significant but not reported.

[a] "HOLS" stands for Cragg's heteroskedastic ordinary least squares estimator.

[b] "IV-GMM" stands for instrumental variable two-step efficient generalized method of moments estimator. The instruments are import unit values, log(scale), and upstream TFP in columns (5) and (7) and their one-period lags in columns (6) and (8).

[c] "Hansen's J p" row reports the p-value for the Hansen's (1982) *J* test of overidentifying restrictions for instrument validity.

Table 3. Political Weights

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
L.Log($X/M*\varepsilon$)	0.390***	0.720***	0.529***	0.390***	0.393***	0.271***	0.514***	0.273***
	(0.048)	(0.159)	(0.110)	(0.048)	(0.048)	(0.034)	(0.103)	(0.033)
REF	-0.676***	-0.548***	-0.520***	-0.676***	-0.674***	-0.628***	-0.530***	-0.627***
	(0.033)	(0.071)	(0.058)	(0.033)	(0.034)	(0.033)	(0.055)	(0.033)
L.Log(*ShEmp*)	-0.268***			-0.268***	-0.261***	-0.179***		-0.174***
	(0.064)			(0.064)	(0.067)	(0.050)		(0.052)
L.Log(*FirmCon*)		0.453***			0.045		0.317***	0.030
		(0.147)			(0.080)		(0.102)	(0.064)
L.Log(*LabInt*)			0.220***			0.205***	0.216***	0.205***
			(0.039)			(0.028)	(0.037)	(0.028)
Union				-0.475***				-0.238**
				(0.082)				(0.093)
Constant	-3.137***	-4.954***	-2.404***	-2.663***	-3.309***	-2.360***	-3.697***	-2.233***
	(0.347)	(0.991)	(0.329)	(0.332)	(0.454)	(0.273)	(0.661)	(0.373)
Observations	961	961	961	961	961	961	961	961
Hansen's J p[a]	0.011	0.698	0.300	0.011	0.012	0.003	0.297	0.003

Notes:
(1) Robust standard errors in parentheses.
(2) *** p<0.01, ** p<0.05, * p<0.10.
(3) "L." stands for one-period (year) lag.
(4) All specifications include industry fixed effects that are jointly significant but not reported.
(5) All specifications use IV-GMM (instrumental variable two-step efficient generalized method of moments estimator.) The instruments are one-period lags of import unit values, log(scale), and upstream TFP.
[a] "Hansen's J p" row reports the p-value for the Hansen's (1982) *J* test of overidentifying restrictions for instrument validity.

Table 4. Preferential Trade Agreements

	(1)	(2)	(3)	(4)	(5)	(6)
L.Log(*X*/*M***ε*)	0.796***	0.769***	0.756***	0.335***	0.578***	0.338***
	(0.222)	(0.207)	(0.200)	(0.040)	(0.133)	(0.039)
REF	-0.590***	-0.645***	-0.707***	-0.621***	-0.544***	-0.620***
	(0.079)	(0.070)	(0.060)	(0.034)	(0.058)	(0.034)
L.*ShM_ANDE*	1.666***	1.189***		0.681***	1.226***	0.691***
	(0.536)	(0.393)		(0.134)	(0.333)	(0.135)
L.*ShM_ANDE*x*REF*		0.515	1.401***			
		(0.314)	(0.466)			
L.Log(*ShEmp*)				-0.224***		-0.219***
				(0.054)		(0.056)
L.Log(*FirmCon*)					0.388***	0.034
					(0.134)	(0.067)
L.Log(*LabInt*)				0.221***	0.242***	0.221***
				(0.030)	(0.044)	(0.030)
Union						-0.348***
						(0.100)
Constant	-3.444***	-3.336***	-3.213***	-2.796***	-4.305***	-2.578***
	(0.711)	(0.656)	(0.616)	(0.305)	(0.915)	(0.406)
Observations	956	956	956	956	956	956
Hansen's J p[a]	0.468	0.258	0.222	0.049	0.387	0.054

Notes: (1) Robust standard errors in parentheses. (2) *** p<0.01, ** p<0.05, * p<0.10. (3) "L." stands for one-period (year) lag. (4) All specifications include industry fixed effects that are jointly significant but not reported. (5) All specifications use IV-GMM (instrumental variable two-step efficient generalized method of moments estimator.) The instruments are one-period lags of import unit values, log(scale), and upstream TFP. [a] "Hansen's J p" row reports the p-value for the Hansen's (1982) *J* test of overidentifying restrictions for instrument validity.

Table 5. Robustness Checks

	(1) ERP	(2) YearEff	(3) ε^{ALT}	(4) ε^{EIV}	(5) IV-TOBIT
L.Log(*X*/*M***ε*)	0.656***	0.090**	0.345***	0.345***	0.695***
	(0.165)	(0.035)	(0.071)	(0.071)	(0.152)
REF	-0.603***		-0.683***	-0.683***	-0.557***
	(0.078)		(0.041)	(0.041)	(0.076)
Constant	-1.408***	-0.989***	-1.865***	-1.861***	-2.973***
	(0.497)	(0.110)	(0.203)	(0.202)	(0.484)
Observations	949	961	912	912	961
Hansen's J p[a]	0.599	0.007	0.113	0.113	n/a

Notes:

(1) Robust standard errors in parentheses.

(2) *** p<0.01, ** p<0.05, * p<0.10.

(3) "L." stands for one-period (year) lag.

(4) All specifications include industry fixed effects that are jointly significant but not reported.

(5) IV-GMM (instrumental variable two-step efficient generalized method of moments estimator) is used in columns (1) through (4). The instruments are one-period lags of import unit values, log(scale), and upstream TFP.

(6) In column (1), log(effective rate of protection) is used in lieu of log(tariffs).

(7) In column (2), year effects are used in lieu of *REF*.

(8) In column (3), an alternative, time invariant import demand elasticity measure from Nicita and Olarreaga (2007) is used.

(9) In column (4), an errors-in-variables corrected import demand elasticity measure (following Gawande and Bandyopadhyay 2000) of the column (3) elasticity estimate is used.

(10) In column (5), IV-TOBIT (instrumental variable Newey's two-step tobit estimator with left censoring) with the same instruments is used.

[a] "Hansen's J p" row reports the p-value for the Hansen's (1982) *J* test of overidentifying restrictions for instrument validity.

Table A.1. Variable Definitions

Name	Definition [Source]
τ_{it}	Advalorem tariff rate (%), 4-dig [DNP]
X_{it}	Total output (000 USD), 4-dig [UNIDO]
M_{it}	Total imports (000 USD), 4-dig [COMTRADE]
ε_{it}	Import demand elasticity: structural estimates [Kee et al. (2004)] combined with GDP [WDI] and imports [COMTRADE] data, 3-dig
$IMPuv_{it}$	Average import unit value of goods (dollars per kilogram) entering the country, 3-dig [Nicita and Olarreaga 2007]
$Scale_{it}$	Value added [UNIDO] divided by the number of firms [Eslava et al. (2004)], 4-dig
$UpstrTFP_{it}$	Upstream total factor productivity (TFP): Weighted average of the TFP [Eslava et al. 2004)] of the upstream sectors excluding itself where weights for the share of inputs per upstream industry is obtained using input-output tables [Nicita and Olarreaga (2001)], 3-dig
REF_t	Trade Reform period dummy (intercept shifter) which is equal to 1 for 1990 and onwards, and 0 otherwise.
$ShEmp_{it}$	Share of employment: Share of the number of the employees in a sector to total number of employees in the country, 4-dig [UNIDO]
$FirmCon_{it}$	Firm Concentration: Ratio of the total number of firms in the country to the number of firms in an industry, 4-dig [Eslava et al. 2004]
$LabInt_{it}$	Industry level employment hours as a share of industry level physical output, 4-dig [Eslava et al. (2004)]
$Union_i$	A dummy variable which is equal to one if there exists labor union activity at the 3-digit industry level [Quintero (2006)]
ShM_ANDE_{it}	Share of imports from ANDEAN Group countries (Bolivia, Ecuador, Peru, and Venezuela) to total imports within a 4-digit ISIC industry for the sample period [COMTRADE]
τ_{it}^{ERP}	Effective rate of protection (%), 4-dig [DNP]
ε_i^{ALT}	Time invariant alternative import demand elasticity measure, 3-dig [Nicita and Olarreaga (2007)]
ε_i^{EIV}	Errors-in-variables corrected measure (following Gawande and Bandyopadhyay 2000) of the time invariant alternative import demand elasticity estimate of Nicita and Olarreaga (2007), 3-dig

Table A.2. Descriptive Statistics

Variable Name	Observations	Mean	Standard Deviation	Minimum	Maximum
Log(τ)	961	-1.668	0.644	-4.107	0.140
Log($X/M*\varepsilon$)	952	1.709	2.209	-4.811	12.450
L.Log($X/M*\varepsilon$)	961	1.754	2.239	-4.630	12.450
REF	961	0.688	0.464	0.000	1.000
IMPuv	961	4.832	4.951	0.183	33.031
L.*IMPuv*	961	4.812	4.802	0.183	31.470
Log(*Scale*)	954	7.368	1.311	3.424	11.965
L.Log(*Scale*)	961	7.286	1.322	1.599	11.965
UpstrTFP	960	1.524	0.137	1.094	2.052
L.*UpstrTFP*	961	1.512	0.133	1.094	2.052
L.Log(*ShEmp*)	961	-5.009	1.357	-10.905	-2.127
L.Log(*FirmCon*)	961	5.120	1.219	1.884	8.763
L.Log(*LabInt*)	961	-0.564	0.836	-3.325	1.606
Union	961	0.712	0.453	0.000	1.000
L.*ShM_ANDE*	956	0.118	0.187	0.000	1.000
L.*ShM_ANDE*x*REF*	956	0.094	0.171	0.000	1.000
Log(τ^{ERP})	949	-1.165	0.902	-5.266	1.556
L.Log($X/M*\varepsilon^{ALT}$)	913	1.770	2.323	-4.518	12.198
L.Log($X/M*\varepsilon^{EIV}$)	913	1.344	2.589	-4.871	12.188

Table A.3. First-Stage Regressions for the Benchmark Specification

	Table 1-Column (8) Specification
REF	-0.240***
	(0.067)
L.*IMPuv*	0.014
	(0.028)
L.Log(*Scale*)	-0.281**
	(0.112)
L.*UpstrTFP*	-0.766***
	(0.255)
Constant	6.128***
	(0.817)
Observations	960
R-squared	0.878

F test of excluded instruments: F(3,879) = 6.08, Prob > F = 0.000.
Kleibergen-Paap (2006) Statistic for underidentification: Chisq(3)=22.15, P-val=0.000.
Cragg-Donald (1993) Statistic for weak identification: F-stat=9.28, Stock-Yogo (2005) critical values": 5% = 13.91, 10%=9.08, 20%=6.46.
Andersen-Rubin (1949) test of endogenous regressor significance: F(3,879)=14.64, P-val=0.000.

Notes:
(1) OLS estimation.
(2) Robust standard errors in parentheses.
(3) *** p<0.01, ** p<0.05, * p<0.10.
(4) The dependent variable is L.Log($X/M*\varepsilon$).
(5) "L." stands for one-period (year) lag.
(6) All specifications include industry fixed effects that are jointly significant but not reported.

Table A.4. Cross-Industry over Time Variation in Estimated Political Weights for Selected Sectors

		Average Political Weight, π			Average Tariff Rate, τ		
ISIC 4	**Description**	**83 to 98**	**<1990**	**≥1990**	**83 to 98**	**<1990**	**≥1990**
3902	Manufacture of musical instruments	0.422	0.708	0.279	0.145	0.282	0.084
3312	Manfc. of made-up textile goods exc. wearing apparel	0.402	0.590	0.284	0.335	0.534	0.221
3903	Manufacture of sporting and athletic goods	0.360	0.652	0.214	0.232	0.324	0.191
3852	Manufacture of photographic and optical goods	0.324	0.430	0.218	0.167	0.233	0.102
3842	Manufacture of railroad equipment	0.320	0.385	0.239	0.212	0.439	0.111
3420	Printing publishing and allied industries	0.100	0.160	0.070	0.224	0.360	0.163
3843	Manufacture of motor vehicles	0.089	0.142	0.063	0.258	0.439	0.178
3116	Grain mill products	0.089	0.135	0.066	0.252	0.334	0.215
3211	Spinning weaving and finishing textiles	0.087	0.138	0.062	0.266	0.463	0.179
3522	Manufacture of drugs and medicines	0.086	0.145	0.057	0.071	0.095	0.060
	All	0.158	0.246	0.115	0.232	0.383	0.163