Varying Political Economy Weights of Protection: The Case of Colombia

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Abstract

In this paper, we examine trade policy determinants and trade reform in a developing country setting by using a political economy model. The government determines tariffs by balancing the political support from producers versus consumers, while placing a higher political weight on producers' welfare relative to average citizens. We then expand the model in several directions to guide our subsequent estimations at the three-digit industry level for Colombia between 1983 and 1998. We account for import substitution motives for protection but describe how the government's move away from these policies leads to unilateral trade liberalization. We innovatively allow the political weights to vary based on key industry variables beyond a common denominator. The sectors with higher employment, labor cost, and preferential trade agreement (PTA) import shares receive a larger political weight compared to otherwise similar sectors. The novelty of our approach is estimating the effect of sectoral characteristics on protection filtered through the political weights. We obtain more realistic estimates for these weights and provide some evidence for a slowing down effect of PTAs on trade liberalization.

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1 Introduction

According to trade theory the optimal trade policy for a small open economy is free trade, but we often observe small developing nations with more restrictive trade policies than larger, more developed nations. Political economy models explain this fact with the trade-off between the well-being of consumers versus producers in the policymaking process [see, for example, Findlay and Wellisz (1982), Mayer (1984), and Grossman and Helpman (1994), among others]. Higher tariffs increase domestic prices and industry profits but hurt consumer welfare. We rely on such a model where the government sets import tariffs balancing the political support from the producers against consumers and expand it in several directions to guide our estimations at the three-digit industry level for Colombia between 1983 and 1998.

In the 1980s and 1990s, several developing countries, such as Chile, Mexico, Turkey, and India, have experienced drastic trade liberalization. Colombia was no exception and in a matter of few years in the early 1990s, average applied tariff rates in the country came down from about 36% to 13%. Unlike some others, Colombia's shift in policy was not due to a conditional loan from an international organization and it was carried out swiftly on a *unilateral* basis. Moreover, the trade reform invariably affected all sectors so several studies used the Colombian experience to study the effects of trade reform by exploiting its widespread application in the economy [e.g. Eslava et al. (2004) and Fernandes (2007)].

In our model, the government places a higher political weight on producer welfare given that producers manage to get organized and press the government for protection while the average citizens remain unorganized in trade policy matters. As we discuss in the next section in detail, our model is related to the "Protection for Sale" model of Grossman and Helpman (1994), yet it provides a more general framework allowing us to account for the developing country experience of Colombia.

First, we introduce an import substitution motive into the government's objective function. We assume the government attempts to develop a national manufacturing base by shielding these industries from foreign competition. This provides an additional layer of protectionist motive and may explain the historically higher protection rates in developing countries. However, when the new government, which did not support such policies, took office in 1990, it swiftly reduced tariffs in all sectors on a unilateral basis in Colombia. This coincides with the experience of other developing countries within the context of the changing view across the globe about import substitution. The novelty of our approach is that as opposed to viewing tariff reduction as an exogenous shift, we model it with a decline in the political weights that affects all sectors and leads to widespread trade liberalization. Yet, cross-sectoral variation in protection persists in the face of the reform.

Second, we allow the political weights to vary based on two key industry characteristics beyond a common base: 1) The share of employment (ratio of employment in the sector to total employment in the economy); and 2) the labor cost share of an industry (measured as wages/value added) which serves as a proxy for skills in the industry and as a measure of its labor intensity. Intuitively, the government may mark up the weights for sectors that employ a large share of the labor force realizing their voting power and it may additionally favor sectors employing unskilled workers that are likely to be adversely affected by trade liberalization. The innovation here is that rather than arguing for a reduced set of variables that might directly affect protection, we propose using a plausible group of variables that might impact the *value* attached to the well-being of producers by the government. Therefore, the effect of these variables on protection is filtered through political weights.

Third, using the same framework, we consider the impact of the preferential/regional trade agreements on political weights by controlling for the sectoral share of imports from preferential trade agreement (PTA) partners as a determinant. Theoretically, PTAs may make it easier to lower external tariffs against non-members [c.f. Freund and Ornelas (2010)]; yet, as in Karacaovali and Limão (2008), we may expect countries to hold back reducing tariffs in sectors that are important for PTA partner countries. This is because each time external tariffs are liberalized, the preferential access is eroded.

We test our benchmark and expanded specifications along with the Grossman and Helpman (1994) model by using 3-digit ISIC level tariff, trade, and production data for Colombia between 1983 and 1998. Having carefully addressed endogeneity concerns, we find strong evidence for tariff rates being inversely related to the elasticity-adjusted import penetration ratio[1] and that there was a common decline in political economy weights as the consensus view about import substitution changed with the new government taking office in 1990. We see that the political weights are marked up for sectors with a high share of employment and labor cost. We also provide some evidence for a slowing down effect of PTAs on unilateral trade liberalization.[2]

We obtain more realistic estimates for political weights by allowing them to vary across sectors and controlling for their decrease due to a common shock which eventually leads to trade reform. 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. Gaining a better understanding of trade policy in developing countries is especially important in the face of the stifled multilateral trade negotiations at the Doha Round, also known semi-officially as the Doha Development Agenda.

The rest of the paper is organized as follows. In the next section, we present the basic theoretical framework that guides the estimations and then develop the econometric model. In Section 3, we describe the data, present the estimation results, and perform robustness checks. Section 4 concludes.

[1] A result which is consistent with the evidence in the empirical literature such as Goldberg and Maggi (1999) for the U.S., Mitra et al. (2002) for Turkey, and McCalman (2004) for Australia among others.

[2] This evidence supports the findings in Limão (2006) for the U.S. and in 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 we do in this paper becomes very important.

2 Theoretical Framework and Econometric Model

2.1 Basic Model

We employ a reduced form political economy model of trade policy in the spirit of the political support function approach of Hillman (1982) to guide the subsequent estimations at the 3-digit industry level for Colombia between 1983 and 1998. First, we start by presenting the basic model below and then in the next subsection, we discuss its connection to the "Protection for Sale" model of Grossman and Helpman (1994).[3] Finally, we develop the main model by accounting for trade reform and sectoral variation in political weights as will be detailed in sections 2.3 and 2.4.

Assume a small open economy where output and factor markets are perfectly competitive. The numeraire good $i = 0$ is produced with labor only: $Y_0(p_0) = L_0$. The other goods are produced with labor, L_i, and a sector specific factor, K_i (that is immobile across sectors), under constant returns to scale: $Y_i(p_i) = f_i(L_i, K_i)$ for $i = 1, ..., N$. 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 we assume there is enough labor for the numeraire good to be always produced in equilibrium.

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.[4] It is straightforward to extend the analysis to other forms of trade policy instruments. Yet, all industries are essentially import-competing and enjoy positive tariff rates in Colombia so focusing on import tariffs is actually reasonable.

The government determines tariffs by balancing the political support from the producers against consumers. Higher tariffs, hence higher domestic prices increase the industry profits/specific factor rents but reduce consumer surplus. The government's problem is

[3] The Grossman and Helpman (1994) model is later tested empirically with the same data set as well.

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

characterized by the maximization of the following weighted social welfare function

$$G(p) = L + CS(p) + \textstyle\sum_{i=1}^{N} \tau_i M_i(p_i) + (\omega + 1) \sum_{i=1}^{N} \pi_i(p_i) \tag{1}$$

L denotes both the aggregate labor supply and labor income. $CS(p) = \sum_{i=1}^{N} \left[\int_{1+\tau_i}^{\infty} D_i(p_i) dp_i \right]$ is the aggregate consumer surplus where $D_i(p_i)$ denotes demand.[5] $M_i(p_i) = D_i(p_i) - Y_i(p_i)$ is the aggregate import demand and assuming away wasteful government expenditures, the tariff revenue, $\sum_i^N \tau_i M_i(.)$, is rebated back to the public in its entirety. $\pi_i(p_i)$ measures the specific factor return for sector i with $Y_i(p_i) = \pi_i^{'}(p_i)$ due to envelope theorem. $\omega > 0$ measures the additional political weight the government places on the welfare of specific factor owners relative to average voters. In the absence of the political weight, i.e. $\omega = 0$, equation (1) boils down to a standard social welfare function. We can imagine specific factor owners obtaining a greater weight than the average citizens under the assumption that they get organized and exert pressure on the government for protection while consumers fail to overcome the collective action problem and cannot organize for free trade, and hence obtain a lower weight [Olson (1965)].

Maximizing equation (1) with respect to τ_i and using $p_i = 1 + \tau_i$ we obtain the following first order condition for an interior solution

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

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

$$\tau_i = -\omega \frac{Y_i(\tau_i)}{M_i'(\tau_i)} \equiv \omega \frac{Y_i(\tau_i)/M_i(\tau_i)}{\varepsilon_i(\tau_i)} \equiv \omega \frac{z_i(\tau_i)}{\varepsilon_i(\tau_i)} \tag{3}$$

where $\varepsilon_i(.)$ stands for the elasticity of import demand.[6] This expression is similar to those

[5] The underlying utility function is quasilinear, such that it is linear in the numeraire good $i = 0$ and concave in other goods.

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

obtained in various political economy models as shown in Helpman (1997). The tariff rate for a sector is directly related to the political weight placed on the well-being of producers (ω), while it is inversely related to the import demand elasticity (ε_i) and import penetration ratio ($M_i/Y_i \equiv 1/z_i$). 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 (M_i) creates a greater price distortion potential putting a downward pressure on tariffs. Moreover, the marginal benefit of a tariff to a producer is higher when it applies to more units (Y_i).

2.2 Grossman and Helpman (1994) Model

In an influential paper, Grossman and Helpman (1994) (GH model henceforth) provide microfoundations for the government's trade policy setting problem by focusing on the influence motive of campaign contributions and obtain the following expression for politically optimal tariffs

$$\tau_i = \frac{I_i - \alpha_L}{a + \alpha_L} \frac{z_i(\tau_i)}{\varepsilon_i(\tau_i)} \tag{4}$$

where $I_i = 1$ if i is an organized sector, and $I_i = 0$ otherwise. α_L denotes the fraction of the population that owns sector specific inputs and a stands for the weight the government places on social welfare relative to campaign contributions. The parameter a can be thought as the equivalent of the inverse of the political weight in our model, $a \equiv 1/\omega$, when we realize that truthful contributions are direct functions of producers' welfare by assumption in the GH model.

For organized sectors the tariffs are set similar to equation (3), whereas for unorganized sectors the GH model predicts negative protection since the lobbies are consumers of the goods produced by other sectors. It is actually reasonable to assume that only a negligible fraction of the population owns specific factors and lobbies for protection so that $\alpha_L \to 0$. In reality, tariff rates are never below zero and for Colombia, tariffs are always positive at the 3-digit industry level. Although this is not a critical assumption, it simplifies the analysis such that lobbies only care about the protection in their own sector and will not lobby against protection in other sectors on the grounds that it would increase their consumer surplus. We will test this empirically for Colombia in the next section.

More specifically, we use the optimal tariffs defined in equation (4) and assume the tariffs are set through the same political process each year to obtain the following specification in error form

$$\tau_{it} = \beta_1 \frac{z_{it}}{\varepsilon_{it}} + \beta_2 I_i \frac{z_{it}}{\varepsilon_{it}} + u_{it} \tag{5}$$

for sectors $i = 1, ..., N$ and years $t = 1, ..., T$. The model predicts negative tariffs for unorganized sectors ($I_i = 0$), i.e. we expect $\beta_1 < 0$, and positive tariffs for organized sectors ($I_i = 1$), i.e. we expect $(\beta_1 + \beta_2) > 0$, (with $\beta_2 > 0$). As we will show in Section 3.2.1, we find $\beta_1 = (-\alpha_L)/(a + \alpha_L) = 0$ and $\beta_2 = (1 - \alpha_L)/(a + \alpha_L) > 0$ which imply $\alpha_L = 0$. Therefore, there is empirical support for the assumption that specific factor owners (the lobbyists) constitute a negligible share of the population (i.e. $\alpha_L \to 0$).

Assuming $\alpha_L \to 0$ and substituting a with $1/\omega$, we can re-write equation (4) as follows

$$\tau_i = \omega I_i \frac{z_i(\tau_i)}{\varepsilon_i(\tau_i)} \tag{6}$$

which is equivalent to equation (3) for organized sectors. Therefore, another implicit assumption in our main model is that all manufacturing sectors at the 3-digit ISIC level are organized. This is a realistic assumption given the fact that in Colombia there is no formal process like the campaign contributions through political action committees as in the U.S. and lobbying by industrialists can be suspected to be especially prevalent in the absence of transparency as argued in Gawande et al. (2009). Yet, even in the U.S., at this level of aggregation all sectors report contributions and lobbying activity. In order to test the GH model with Colombian data, we will use a proxy measure for the organization dummy and later test the sensitivity of our main results to restricting the sample to "organized" sectors only.

2.3 Trade Reform and Benchmark Specification

In the late 1980s and early 1990s there was a change in the economic consensus such that old import substitution policies were abandoned for more liberal trade policies in Colombia. This was likely encouraged by World Bank research and policy dialogue [Edwards (1997)]. Edwards (2001) indicates that César Gaviria, who was the President of Colombia from 1990 to 1994, "developed from early on a critical view regarding CEPAL's [Economic Commission for Latin America] import substitution development strategies." After President Gaviria took office in 1990, his government swiftly reduced tariffs *unilaterally* in all sectors from an average of 36% to 13% in a matter of 3 years, and these rates have stayed about the same since then (see Figure 1).[7]

In order to account for the trade reform experience in the developing country of Colombia, we expand the government objective in the basic model in Section 2.1 to incorporate an import substitution motive. We model the government attaching an extra value to domestic production, and hence the producer surplus, on top of any political weight on producers' welfare because of industry pressure/lobbying. More specifically, $G(p)$ as defined in equation (1) now includes the additional term $\phi \sum_{i=1}^{N} \left[\int_0^{1+\tau_i} Y_i(p_i) dp_i \right]$. Maximizing the expanded government objective with respect to τ_i yields the following politically optimal tariffs:

$$\tau_i = (\omega + \phi) \frac{z_i(\tau_i)}{\varepsilon_i(\tau_i)} \tag{7}$$

Given the fact that trade reform was a common shock that hit all sectors, we can model it as a change in the view of the government moving away from import substitution. Therefore, we can conjecture that ϕ dropped down to zero after the new government took office in 1990 and we control for it in the estimations.

[7] Although Colombia is a founding member of the World Trade Organization since 1995, the Colombian trade liberalization that took place starting in 1990 was not in response to a multilateral process, and hence did not entail reciprocity [World Trade Organization (1996)]. Yet, this unilateral liberalization that occurred prior to joining the WTO was recognized as part of Colombia's tariff concessions.

Re-expressing equation (7) in log linear and error form we obtain

$$\log \tau_{it} = \alpha + \beta \log \frac{z_{it}}{\varepsilon_{it}} + \gamma D^{90s} + \theta_i + \upsilon_{it} \tag{8}$$

for sectors $i = 1, ..., N$ and years $t = 1, ..., T$, where $D^{90s} = 1$ for $t \geq 1990$ and zero otherwise. D^{90s} points to the common decline in political weights due to the shift away from import substitution view starting from 1990 and onwards. We employ industry fixed effects, θ_i to account for other factors that might make tariffs differ across sectors in a systematic way given the parsimonious nature of the model.

Based on theory, the expected sign for β is positive indicating that tariffs are directly related to the inverse of the elasticity-adjusted import penetration ratio, $z/\varepsilon = M/(Y\varepsilon)$. The specification allows us to estimate the additional political weight the government places on the well-being of the industry relative to average citizens. This weight declines in 1990 and onwards leading to a unilateral trade liberalization shock affecting all sectors. More specifically, the political weight estimates can be obtained by tying the regression coefficients to model parameters in equation (7) as follows: $\log \hat{\omega} = \hat{\alpha}$ and $\log(\hat{\omega} + \hat{\phi}) = \hat{\alpha} - \hat{\gamma}$. Regression results of this benchmark specification appear in Section 3.2.2.

2.4 Varying Political Weights

Next, we model the political weights to vary across sectors, and hence, in effect, replace ω with ω_i in equation (7). We allow a common element that measures the additional weight attached to the well-being of all producers relative to average citizens but then allow this common weight to be marked up or discounted based on a number of industry characteristics by expanding the benchmark regression equation (8) as

$$\log \tau_{it} = \alpha + \beta \log \frac{z_{it}}{\varepsilon_{it}} + \gamma D^{90s} + \sum_k \delta_k X_{kit} + \theta_i + \upsilon_{it} \tag{9}$$

where X_{kit} is the k^{th} factor measuring sectoral variation in political weights. Therefore, the varying weights are estimated as follows

$$\log \hat{\omega}_{it} = \hat{\alpha} + \sum_{k} \hat{\delta}_k X_{kit} \tag{10}$$

The important point to realize here is that rather than arguing for a reduced set of variables that might affect protection directly, we propose using a plausible group of variables that might impact the *value* attached to the well-being of producers by the government. In that respect, these variables affect tariffs through the political weights.

Keeping a parsimonious approach, we focus on two key industry level variables ($k = 2$): 1) The share of employment (the ratio of employment in the sector to total employment in the economy); and 2) the labor cost share of an industry (measured as wages/value added) which serves as a proxy for skills in the industry and as a measure of its labor intensity.

Intuitively, these two variables are expected to affect the political weights as follows: First, an industry with a higher share of employment commands more votes and may thus be more likely to be favored by politicians[Caves (1976)]. Second, labor intensive sectors that employ mostly unskilled workers may be favored based on a social justice motive as they may be impacted more adversely from import competition [Baldwin (1985)].

Since we expect industries with more employees and less skilled workers to have marked up political weights from the average, we predict $\delta_k > 0$ for $k = 1, 2$.

Finally, we use the same framework to account for the effect of preferential trade agreements (PTAs) on trade policy setting. 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. For example, Karacaovali (2014) shows that once an 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 in products imported duty-free from preferential partners less than non-PTA goods. However, the tariff reduction in products imported from new EU members weren't different from non-PTA goods. 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 effects of PTAs on global free trade.[8]

Although it is possible that a PTA may exert a downward pressure on external tariffs [c.f. Freund and Ornelas (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.

We will control for the influence of PTAs through the political weights. For instance, the government may mark up the protectionist weight allocated to a sector if it is an important export sector for the PTA partner under the stumbling block view. More specifically, we will consider the share of PTA imports relative to total imports in an industry as the third industry variable for X_{kit} in equation (9) to capture the so-called stumbling versus building block effects across industries. The results are presented in Section 3.2.3.

[8] The stumbling versus building block terminology refers to Bhagwati (1991).

2.5 Specification Issues

All estimations including the benchmark econometric model are potentially subject to endogeneity given the fact that elasticity-adjusted inverse import penetration, z/ε, which is the main right-hand-side variable, is a function of domestic prices, and hence tariffs. Therefore, OLS estimation is expected to produce biased results. As a way to get around the problem of endogeneity, we 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, that is tariffs, over time. Therefore, we consider an Instrumental Variables (IV) approach. While the validity and strength of instruments will be discussed in the next section, here we provide a brief intuition behind the choice of instruments.

First, we use import unit values as a proxy for world prices at the border which are correlated with domestic prices by definition but not Colombia's tariffs because it is a small country tradewise. Therefore, import unit values are useful to instrument for z/ε. Second, we employ a measure of scale defined by value added per firm as an instrument for import penetration given that scale is likely to be correlated with fixed costs of entry to an industry, and hence affect import penetration.[9] 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, we rely on the capital to output ratio of an industry as a measure of capital intensity, and hence comparative advantage, which affects Colombia's trade.

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, we use industry fixed effects in our main specifications while the instrumental variables approach is also expected to reduce such bias. Finally, other econometric concerns are addressed in

[9] The entry barriers may affect both domestic and foreign competitors and hence impact both domestic production and imports in an industry [Trefler (1993)].

Section 3.3 after estimation results are discussed.

3 Empirics

3.1 Data

The data for the estimations cover twenty-eight 3-digit International Standard Industrial Classification (ISIC) industries between 1983 and 1998, with the exception of 1986 and 1987. The ISIC codes are defined in Table A.1 and the descriptions of all variables used in the empirical analysis are provided in Table A.2. Here we present the data sources.

MFN applied 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 3-digit ISIC industry level by using simple averages.[10]

The main production data covering total output, value added, wages, number of firms are available through UNIDO's Industrial Statistics Database while bilateral and aggregate import data are from COMTRADE, UN Statistics Division. Import demand elasticity is obtained by combining the structural estimates in Kee et al. (2004) with GDP data from the World Bank's World Development Indicators (WDI) and import data from COMTRADE. An alternative time-invariant import demand elasticity measure is obtained from Nicita and Olarreaga (2007) as a robustness check. Import unit values, measured as dollars per kilogram, serve as a proxy for world prices at the border and are taken from Nicita and Olarreaga (2007) as well.

The political organization dummy, I_i, is from Quintero (2006) and it measures the indication of organization based on membership in economic associations and groups such as the National Association of Industries (ANDI). Finally, we rely on capital stock, labor, and output data from Eslava et al. (2004) to compute the capital to output ratio and the labor

[10] I thank Marcela Eslava for providing this data set. 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.

share of an industry.

Table 1 lists the average tariff rates and their dispersion across 3-digit industries for the main sample. There is a significant reduction in the average tariff rates–a process which starts in 1990–while the dispersion declines to a lesser extent as can be observed from the coefficients of variation.[11] The same trend can also be observed in Figure 1 which depicts the tariff rates at the 3-digit ISIC level over time. The trade reform affects all sectors, yet there is cross-industry variation which we conjecture to be attributed to political economy forces based on the model we developed in the previous section. Therefore, various specifications of the model are formally tested next. Table A.3 provides descriptive statistics for all the variables used in the estimations.

3.2 Estimation Results

3.2.1 Grossman and Helpman (1994) Model Specification

Table 2 presents the regressions based on equation (5) for politically optimal tariffs as introduced in Section 2.2. The GH model predicts a negative coefficient on the elasticity-adjusted inverse import penetration ratio, (z/ε), and a positive coefficient when (z/ε) is interacted with the organization indicator as we discussed. The tariff rates are actually positive in all 3-digit sectors and it is plausible to imagine that industries at this level of aggregation will not lobby against protection in other sectors. Our conjecture for Colombian data is that lobbying activity will be highly concentrated in a small fraction of the population constituting industrialists which implies $\alpha_L \to 0$, and hence the coefficient on (z/ε) will be equal to zero. More specifically, we predict $\beta_1 = (-\alpha_L)/(a + \alpha_L) = 0$ in equation (5).

In column 1, we directly follow the original model and estimate contemporaneous variables without a constant term using ordinary least squares (OLS). As is done in the literature, we also estimate equation (5) with a constant term in column 2 [e.g. Mitra et al. (2002) and

[11] 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.

Gawande and Bandyopadhyay (2000)]. In both cases, the coefficient on (z/ε), that is β_1, is not significantly different from zero which confirms our conjecture.

As discussed in Section 2.5, the endogeneity of the main right-hand-side (RHS) variable, (z/ε), is a valid concern. First, we check and find that endogeneity is present through a Durbin-Wu-Hausman test. Then, as an initial step to address this concern, we use one-year lags of the right-hand-side variables in columns 3 and 4. However, given the persistence in variables, this will be a weak method to address the endogeneity so we resort to an instrumental variables (IV)/two-stage least squares (2SLS) approach in columns 5 through 8 [like Mitra et al. (2002), and Gawande and Bandyopadhyay (2000), among others).

Under all specifications, the estimates indicate $\beta_1 = (-\alpha_L)/(a+\alpha_L) = 0$, hence lobbying is a minority activity ($\alpha_L \to 0$). As discussed in Section 2.2, when we have $\alpha_L \to 0$, the tariff equation under Grossman and Helpman becomes equivalent to the optimal tariffs for organized industries in our main reduced form model which is defined in equation (3).

The coefficient on the elasticity-adjusted inverse import penetration, (z/ε), interacted with the organization dummy (i.e. β_2) is positive and significant at the 1% level which is consistent with the findings in the literature.[12] The estimate for a (the weight on social welfare relative to contributions in the government objective) is 83.3[13] which is comparable to Mitra et al.'s (2002) estimates for Turkey ranging between 76.3 and 104.3.

We fail to reject the validity of the instruments defined in Section 2.5 based on Sargan's (1958) overidentifying restrictions test for which the p-values are reported in Table 2. Instruments will be further discussed in Section 3.3.

3.2.2 The Benchmark Specification

As discussed in Section 2.2 and the previous subsection, the Grossman and Helpman tariff equation becomes identical to the main expression of our basic model [equation (3)] for

[12] Accordingly, $(\beta_1 + \beta_2) > 0$ as well.

[13] This is obtained by using $\hat{\alpha}_L = 0$, and hence $\hat{\beta}_2 = (1 - \hat{\alpha}_L)/(\hat{a} + \hat{\alpha}_L) = 1/\hat{a} = 0.012$.

organized industries when lobbies make up a small fraction of the population, which appears to be the case in Colombia according to the data. Yet, we would like to move beyond the basic setup and expand our analysis by taking a more general approach to protection motives. The government may give in to pressures from the industry about tariff protection but also has an import substitution motive which is abandoned in 1990 when the new administration takes office and the consensus view changes from then on. This approach is captured by equation (8) which serves as the benchmark specification (Section 2.3). D^{90s} is a dummy variable which takes the value one for 1990 and onwards and *measures the common decline in political economy weights.*

Given the endogeneity of the elasticity-adjusted inverse import penetration ratio, (z/ε), we again use the one-period lags of the right-hand-side variables and an instrumental variables (IV) approach. More specifically, we employ 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)]. Heteroskedasticity is confirmed to be a problem using a Pagan and Hall (1983) test and warrants the use of an IV-GMM estimator.

The results from this benchmark specification appear in column 1 of Table 3. There is strong support for our model where the natural logarithm of (z/ε) is found to be inversely related to log tariffs and there is evidence for a common shock lowering government political economy weights in all sectors, both at the 1% significance level. As discussed above, the GH model under the assumption of all sectors being organized and lobbies making a negligible share of the population produces the same tariff protection expression in our basic model [see equation (6)]. In our benchmark model, we expand on this by accounting for the trade reform as well and ultimately estimate equation (8) in column 1. However, as a robustness check, we re-estimate the benchmark model restricting the sample to organized sectors only in column 2 and the results are highly similar both qualitatively and quantitatively. This confirms our expectation at this level of aggregation, especially for Colombia where, due to lack of transparency, all sectors are expected to be involved in lobbying the government in one form or another.[14]

3.2.3 Varying Political Economy Weights

In columns 3 through 7 of Table 3, we provide estimates for equation (9) where the political economy weights measuring the relative importance of producer welfare to consumers are specifically designed to vary across sectors beyond a common denominator. As discussed in Section 2.4, we focus on two key industry characteristics affecting these weights: 1) The share of employment in an industry relative to the whole economy; and 2) the labor cost share (wages/value added) of an industry as a proxy for skills. In columns 3 and 4 of Table 3, each variable is first considered one at a time and in column 5, both are included as regressors. We see that the political weights are marked up for sectors with a high share of employment and labor cost. First, sectors employing a larger share of the working age population receive a higher weight than the average given that they have a bigger voting power overall and obtain a favorable treatment from the government. Second, labor intensive sectors relying mostly on unskilled workers are more adversely affected from increasing import competition so they are given a higher weight and protected more.[15]

In columns 6 and 7 of Table 3, we investigate the role of preferential trade agreements (PTAs) on political weights. Our conjecture, as put forth in Section 2.4, is that governments may mark up the political weights for sectors that are important for PTA member nations, which would introduce a protective bias or friction in the face of trade liberalization. We 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. The Andean Group is the second biggest trade

[14] Even in the U.S., at the 3 digit industry level, all sectors provide political contributions [Gawande et al. (2009)].

[15] 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.

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 the 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 the Andean Group.

We account for the effect of PTAs on the political weights with the share of imports from the Andean Group to total imports in an industry.[16] In column 6, focusing on the PTA import share only, and in column 7, including PTA import share in addition to employment and labor cost shares of an industry, we find weak evidence, at the 10% significance level, of a higher weight in sectors important for PTA partners. This may be due to the fact that the Andean agreement was not initially deep. Then, it got strengthened whereby barriers to virtually all intra-regional trade were eliminated coinciding with the period of general trade reform in the country [World Trade Organization (1996)]. Therefore, the erosion in preferences was mostly avoided and we would expect only a weak stumbling block effect in the case of Colombia.

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

As indicated in Section 2.4, following the equations (7) and (8), the political weight estimates in the benchmark specification are related to model parameters as follows: $\log \hat{\omega} = \hat{\alpha}$ and $\log(\hat{\omega} + \hat{\phi}) = \hat{\alpha} - \hat{\gamma}$. Therefore, the constant term provides an estimate of the common

[16] One may suspect whether PTA import share could be endogenous to tariffs. The fact that we use the lag of it should alleviate such a potential problem. However, we 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.

political economy weight before the new government takes office, while the constant term plus the coefficient on D^{90s} is the political weight henceforth. Accordingly, the estimate for the political weight from column 1 of Table 3 is 0.203 before 1990 and 0.088 afterwards. This indicates that the government values producer welfare 20% more than an average citizen and this figure goes down to 9% after the trade reform. These estimates, although arguably small, are significantly higher than the comparable estimates testing the GH model [c.f. Gawande and Krishna (2003) and Imai et al. (2009)]. For example, for the U.S., Goldberg and Maggi (1999) estimate the political weight to be 0.014 and Gawande and Bandyopadhyay (2000) estimate it as 0.0003. Mitra et al.'s (2002) estimates for Turkey range between 0.010 and 0.013 while our GH model estimate was 0.012 (please see Table 2).

The innovation of our approach in estimating political weights is captured by the specification in column 7 of Table 3. Letting the weights vary based on the employment, labor cost, and PTA import shares of an industry, we obtain sector specific weights as indicated in equation (10). The average political weight estimate before 1990 is 0.145 which decreases to 0.062 afterwards. In Figure 2, the variation in political weights is illustrated over the sample period and the cross-industry variation is noteworthy. In Table 4, we present average tariffs and political weights before 1990 and afterwards along with the average key industry characteristics. As can be observed in Table 4 (and also Figure 2), the highest tariff rates are in the apparel (ISIC 322) and footwear industries (ISIC 324), whereas the lowest one is in petroleum refineries (ISIC 353). The highest political weights are in the apparel (ISIC 322) and food products (ISIC 311) industries while the lowest one is in tobacco (ISIC 314). These industries partially reflect the fact that political weights are designed to vary based on the employment, labor cost, and PTA import shares. For instance, the food products (ISIC 311) employ the highest share of workers in the country and also has a high labor cost and PTA import share. The tariffs and political weights are positively correlated with $r = 0.55$, yet there is substantial variation across industries. Therefore, we do indeed capture the effect of sectoral variables on tariffs filtered through the political weights.

Mitra et al. (2006), in their search for more realistic estimates for political weights, find that their estimates 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, our approach in this paper not only provides a novel way of estimating political weights but also produces more realistic estimates as compared to earlier related studies in the literature.

3.3 Specification Issues and Robustness

The Hansen (1982) test of overidentifying restrictions shows that the instruments–import unit values, scale measure, and capital to output ratio–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 row of table 3 with IV-GMM specifications. For instance, in column 7 of Table 3, under the main varying political economy weights specification, the p-value for Hansen's J test is 0.64 so we fail to reject the validity of instruments. The first-stage regressions from this specification are presented in Appendix Table A.4. The excluded instruments are jointly 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 and Donald (1993) statistic is 9.44 which lets us reject the presence of weak instruments at the 10% level using Stock and Yogo (2005) critical values. Finally, the Andersen and Rubin (1949) test, which is robust to the presence of weak instruments, indicates that the endogenous regressor, z/ε, is significant at the 1% level.

As a robustness check, an alternative time-invariant import demand elasticity measure from Nicita and Olarreaga (2007) was used. We also applied 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, our original time-varying import demand elasticity measure is the preferred one.

Tariff rates in general are censored from below given that they cannot be negative so we tested the robustness of the results to the IV-GMM procedure by considering Newey's (1987) two-step tobit estimator (IV-Tobit) instead. The results were 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, we do not expect the potential censoring from below to be a problem for our data set.

4 Concluding Remarks

Based on several political economy of trade policy models, tariff rates are expected to be inversely related to elasticity-adjusted import penetration ratios in a small open economy, which is also what we obtain in our basic model relying on the political support function approach of Hillman (1982). The government determines tariffs by balancing the political support from the producers against consumers and places a higher political weight on producers' welfare relative to average citizens. This is because the producers manage to get organized and press the government for protection while the consumers cannot overcome the collective action problem.

Our model is inherently connected to the seminal "Protection for Sale" model of Grossman and Helpman (1994), as we discuss in Section 2.2. However, our approach provides a more general setup that enables us to expand it in several directions to account for the developing country experience of Colombia. First, we introduce import substitution motives into the government's objective in setting tariffs and then control for the move away from these policies in the 1990s after the new government takes office by allowing a common drop in the political weights. Second, we allow the political weights to vary beyond a common denominator based on two key industry characteristics: the share of employment in an industry relative to the national total and the labor cost share of an industry as a proxy for skills and labor intensity. The government may mark up the weights for sectors that employ a large share of the labor force to garner their political support and it may display a protective bias for sectors employing unskilled workers likely to be adversely affected by trade liberalization. Third, we account for the effect of preferential trade agreements (PTAs) on political weights using the same framework. Theoretically, it is plausible that PTAs may make it easier to lower external tariffs against non-members. Yet, under the stumbling block rationale, erosion of preferential benefits may be slowed down because the elimination of preferences would mean the end of the PTA itself.

The novelty of our approach is not only allowing sectoral variation in political economy weights but also capturing the effect of key sectoral variables on tariffs mediated through these political weights which is different from the estimations in the earlier literature.

We test our benchmark and expanded specifications along with the Grossman and Helpman (1994) model by using 3-digit ISIC level tariff, trade, and production data for Colombia between 1983 and 1998. Having carefully addressed endogeneity concerns, we find strong evidence for tariff rates being inversely related to the elasticity-adjusted import penetration ratio and that there was a common decline in political economy weights as the consensus view about import substitution changed with the new government taking office in 1990. Furthermore, the sectors with higher employment, labor cost, and PTA import shares received a larger political weight compared to otherwise similar sectors.

In sum, we provide a general framework that accounts for alternative motives of trade policy formation and we document sectoral variation in the political weights on the well-being of producers relative to consumers. We obtain more realistic estimates for political weights by allowing them to vary across sectors and by controlling for their decrease due to a common shock which eventually leads to trade reform. We also provide some evidence of a slowing down effect of PTAs on unilateral trade liberalization.

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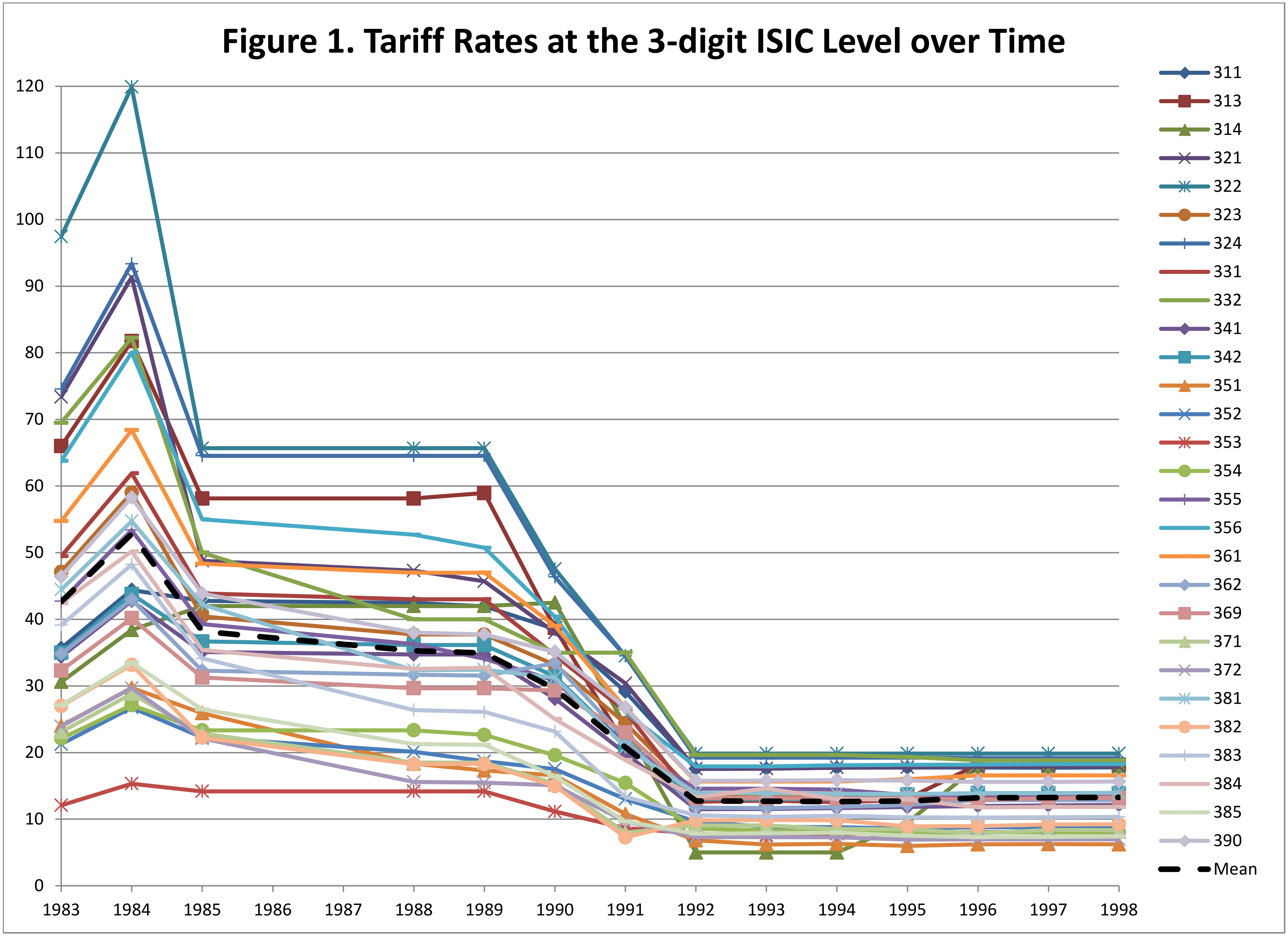
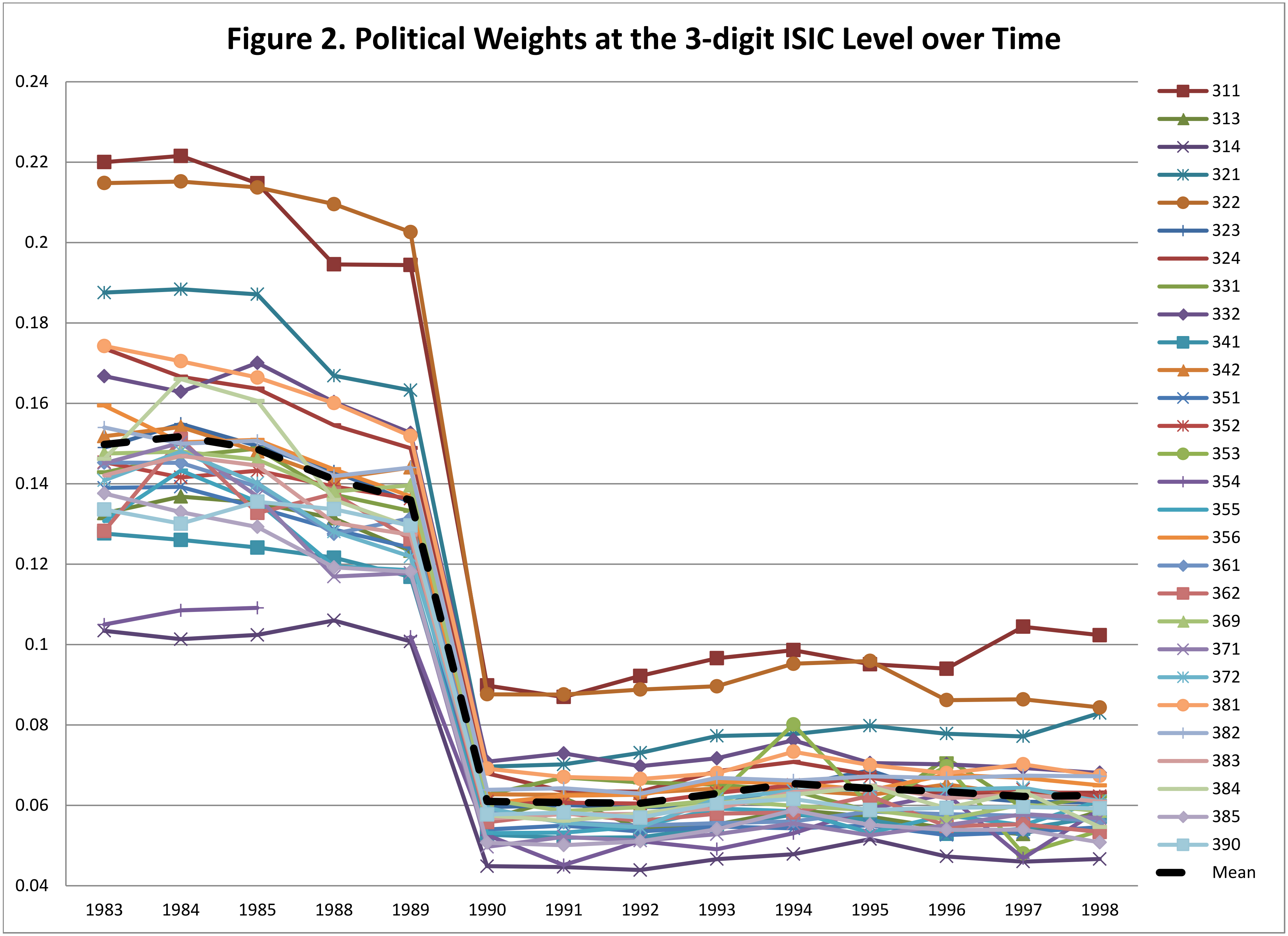


Note: The dashed line depicts average tariff rates.

Source: DNP, Colombia



Note: The dashed line depicts average political weights.
Source: Author's calculations

Table 1. Descriptive Statistics for 3-digit ISIC Level Advalorem Tariffs (%) by Year

Year	Observations	Mean	Standard Deviation	Coefficient of Variation	Minimum	Maximum
1983	27	43.788	19.434	0.444	21.264	97.444
1984	27	54.219	23.946	0.442	26.750	119.911
1985	27	39.075	12.735	0.326	22.097	65.667
1988	26	36.554	13.932	0.381	15.575	65.667
1989	27	35.719	13.965	0.391	15.482	65.667
1990	27	30.256	9.994	0.330	14.981	47.556
1991	27	21.218	8.172	0.385	7.265	35.000
1992	28	12.724	4.158	0.327	5.000	19.856
1993	28	12.694	4.254	0.335	5.000	19.859
1994	28	12.637	4.268	0.338	5.000	19.856
1995	28	12.750	4.127	0.324	5.981	19.856
1996	28	13.209	4.301	0.326	6.223	19.855
1997	28	13.253	4.266	0.322	6.246	19.855
1998	28	13.271	4.262	0.321	6.222	19.855

Table 2. Grossman and Helpman (1994) Model Specification

	(1) OLS[a]	(2) OLS[a]	(3) OLS[a]	(4) OLS[a]	(5) IV(2SLS)[b]	(6) IV(2SLS)[b]	(7) IV(2SLS)[b]	(8) IV(2SLS)[b]
$(z_{it}/\varepsilon_{it})$	0.000	0.000			0.000	0.000		
	(0.000)	(0.000)			(0.000)	(0.000)		
$I_i \times (z_{it}/\varepsilon_{it})$	0.003***	0.001***			0.012***	0.011**		
	(0.000)	(0.000)			(0.003)	(0.006)		
$(z_{it-1}/\varepsilon_{it-1})$			0.000	-0.000			0.000	0.000
			(0.000)	(0.000)			(0.000)	(0.000)
$I_i \times (z_{it-1}/\varepsilon_{it-1})$			0.003***	0.001***			0.012***	0.012*
			(0.000)	(0.000)			(0.002)	(0.006)
Constant		0.224***		0.228****		0.023		0.013
		(0.009)		(0.009)		(0.127)		(0.139)
Observations	386	386	384	384	386	386	384	384
R-squared	0.23	0.14	0.21	0.10	n/a	n/a	n/a	n/a
Sargan Overid p-val[c]	n/a	n/a	n/a	n/a	0.89	0.97	0.92	0.49
F statistic	58.49	31.16	51.28	21.27	16.03	2.08	22.66	1.85
F p-val	0.000	0.000	0.000	0.000	0.00	0.13	0.00	0.16

Notes:
(1) The dependent variable is the advalorem tariff rate at the 3-digit ISIC level, τ_{it}.
(2) *** p<0.01, ** p<0.05, * p<0.10.
(3) Robust standard errors in parentheses.
[a] "OLS" stands for ordinary least squares estimator.
[b] "IV (2SLS)" stands for instrumental variables (two-stage least squares) estimator.
[c] "Sargan Overid p-val" row reports the p-value for Sargan's (1958) overidentifying restrictions test for instrument validity.

Table 3. Benchmark Model and Varying Political Economy Weights

	(1) IV-GMM[a]	(2)[b] IV-GMM[a]	(3) IV-GMM[a]	(4) IV-GMM[a]	(5) IV-GMM[a]	(6) IV-GMM[a]	(7) IV-GMM[a]
$\log(z_{it\text{-}1}/\varepsilon_{it\text{-}1})$	0.251***	0.245***	0.255***	0.212***	0.220***	0.269***	0.232***
	(0.065)	(0.079)	(0.065)	(0.052)	(0.052)	(0.067)	(0.053)
D^{90s}	-0.831***	-0.835***	-0.828***	-0.830***	-0.823***	-0.825***	-0.818***
	(0.055)	(0.071)	(0.054)	(0.054)	(0.054)	(0.055)	(0.054)
Employment Share$_{it\text{-}1}$			3.573**		3.267**		3.222**
			(1.424)		(1.369)		(1.423)
Labor Cost Share$_{it\text{-}1}$				1.195**	1.291**		1.387**
				(0.525)	(0.530)		(0.543)
PTA Import Share$_{it\text{-}1}$						0.306*	0.272*
						(0.176)	(0.158)
Constant	-1.596***	-1.573***	-2.204***	-1.667***	-2.235***	-1.765***	-2.380***
	(0.236)	(0.288)	(0.331)	(0.261)	(0.356)	(0.264)	(0.376)
Observations	384	252	384	384	384	384	384
Hansen's J p-val[c]	0.55	0.91	0.49	0.70	0.63	0.55	0.64
F statistic	117.84	69.45	97.30	125.16	108.86	108.36	101.61
F p-val	0.000	0.000	0.000	0.000	0.00	0.00	0.00

Notes:

(1) The dependent variable is the natural logarithm of advalorem tariff rate at the 3-digit ISIC level, $log\tau_{it}$.

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

(3) Robust standard errors in parentheses.

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

[a] "IV-GMM" stands for instrumental variable two-step efficient generalized method of moments estimator.

[b] Column 2 uses the restricted sample of organized industries only, i.e. for which $I_i = 1$.

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

Table 4. Average Tariff Rates, Political Weights, and Industry Characteristics*

ISIC Code	Description	Tariffs		Political Weights		Emp. Share	Labor Cost Share	PTA Import Share
		<90	≥90	<90	≥90			
311	Food prod.	41.45	21.29	20.90	9.56	16.52	15.69	36.10
313	Beverages	64.60	18.47	13.19	5.66	5.68	8.71	12.22
314	Tobacco	39.00	16.11	10.28	4.66	0.49	5.75	18.01
321	Textiles	61.30	21.43	17.86	7.62	10.07	20.34	10.82
322	Apparel	82.87	24.56	21.12	8.91	10.74	30.97	1.92
323	Leather prod.	44.40	16.86	14.63	6.19	1.50	25.86	6.75
324	Footwear	72.31	23.93	16.15	6.56	2.94	27.53	12.52
331	Wood products	48.26	16.79	14.18	6.40	1.28	22.89	27.89
332	Furniture	56.35	22.75	16.25	7.11	1.71	34.60	8.70
341	Paper & prod.	36.34	14.53	12.33	5.44	2.02	15.68	2.97
342	Print. & publish.	37.54	16.35	14.79	6.32	3.57	23.09	3.72
351	Indust. chemicals	23.05	7.93	13.30	5.41	2.54	14.01	8.99
352	Other chemicals	21.83	10.31	14.12	6.30	6.11	15.84	6.07
353	Petrol. refineries	13.99	8.34		6.15	2.63	14.62	42.36
354	Miscel. petrol.&coal	23.71	10.23	10.61	5.32	0.27	8.33	26.73
355	Rubber products	41.15	16.69	12.95	5.62	0.92	20.30	7.86
356	Plastic products	60.45	21.47	14.82	6.45	4.84	20.51	8.78
361	Pottery china	53.09	19.82	13.77	5.67	1.25	21.55	7.31
362	Glass and prod.	34.65	15.71	13.52	5.69	1.17	19.22	15.42
369	Oth. non-metal. min. prod.	32.61	16.02	14.38	5.95	4.57	15.50	11.83
371	Iron and steel	22.34	9.23	13.34	5.37	1.52	15.13	16.63
372	Non-ferrous metals	21.35	8.28	13.58	6.09	0.38	16.71	52.98
381	Fabricated metal products	41.25	16.58	16.46	6.89	5.62	23.67	9.78
382	Machinery	23.78	9.79	14.81	6.59	3.79	25.39	0.73
383	Machinery electric	34.79	12.12	13.82	6.06	3.16	21.27	1.00
384	Transport equipment	38.66	14.81	14.76	5.97	3.11	20.86	8.42
385	Prof. & scientific equip.	25.90	8.82	12.74	5.32	0.52	18.29	0.55
390	Other manufac. products	44.85	19.10	13.25	5.91	1.45	22.28	2.88

*Note: All variables are expressed as percentages for illustrative purposes.

Table A.1. International Standard Industrial Classification (ISIC) 3-Digit Classification

ISIC Code	Description
311	Food products
313	Beverages
314	Tobacco
321	Textiles
322	Wearing apparel except footwear
323	Leather products
324	Footwear except rubber or plastic
331	Wood products except furniture
332	Furniture except metal
341	Paper and products
342	Printing and publishing
351	Industrial chemicals
352	Other chemicals
353	Petroleum refineries
354	Miscellaneous petroleum and coal products
355	Rubber products
356	Plastic products
361	Pottery china earthenware
362	Glass and products
369	Other non-metallic mineral products
371	Iron and steel
372	Non-ferrous metals
381	Fabricated metal products
382	Machinery except electrical
383	Machinery electric
384	Transport equipment
385	Professional and scientific equipment
390	Other manufactured products

Table A.2. Variable Definitions

Variable Name	Variable Definition [Source]
τ_{it}	Advalorem tariff rate (%) for ISIC 3-digit industry i and year t [DNP]
$z_{it} \equiv Y_{it}/M_{it}$	Y_{it}: Total output (000 USD) [UNIDO]; M_{it}: Total imports (000 USD) [COMTRADE] for ISIC 3-digit industry i and year t
ε_{it}	Structural estimates from Kee et al. (2004) combined with GDP data [WDI] and imports [COMTRADE] for ISIC 3-digit industry i and year t
I_i	Political organization dummy which is equal to 1 if there is indication of organization based on membership in economic associations and groups for ISIC 3-digit industry i [Quintero (2006)]
D^{90s}	Dummy variable which is equal to 1 for 1990 and onwards, and 0 otherwise.
$(Employment\ Share)_{it}$	Share of the number of employees in ISIC 3-digit sector i to total number of employees in the country in year t [Eslava et al. (2004)]
$(Labor\ Cost\ Share)_{it}$	Wages to value added ratio in ISIC 3-digit sector i, year t [UNIDO] which serves as a skills proxy for the industry and as a measure of its labor intensity.
$(PTA\ Import\ Share)_{it}$	Ratio of imports from ANDEAN Group countries (Bolivia, Ecuador, Peru, and Venezuela) to total imports in ISIC 3-digit industry i, year t [COMTRADE]
$(Import\ Unit\ Value)_{it}$	Average import unit value of goods (dollars per kilogram) entering the country in ISIC 3-digit sector i, year t [Nicita and Olarreaga (2007)] as a proxy for world prices at the border
$(Scale\ Measure)_{it}$	Value added divided by the number of firms in sector ISIC 3-digit i, year t [UNIDO] accounting for fixed costs of entry to an industry
$(Capital\ to\ Output\ Ratio)_{it}$	Industry level capital stock as a share of industry level physical output for ISIC 3-digit sector i, year t [Eslava et al. (2004)]

Table A.3. Descriptive Statistics

Variable Name	Count	Mean	Standard Deviation	Minimum	Maximum
τ_{it}	384	0.2485	0.1776	0.05	1.1991
$\log \tau_{it}$	384	-1.6110	0.6553	-2.9957	0.1816
$(z_{it}/\varepsilon_{it})$	384	1737.107	33,183.83	0.2269	650,248.9
$\mathrm{I}_i \times (z_{it}/\varepsilon_{it})$	384	17.9625	49.7199	0	654.183
$(z_{it-1}/\varepsilon_{it-1})$	384	2443.381	35,917.41	0.2269	650,248.9
$\mathrm{I}_i \times (z_{it-1}/\varepsilon_{it-1})$	384	18.6388	49.8297	0	654.183
$\log(z_{it-1}/\varepsilon_{it-1})$	384	1.9869	1.7507	-1.4830	13.3851
I_i	384	0.6563	0.4756	0	1
D^{90s}	384	0.6510	0.4773	0	1
$Employment\ Share_{it-1}$	384	0.0362	0.0363	0.0019	0.2002
$Labor\ Cost\ Share_{it-1}$	384	0.1977	0.0717	0.0312	0.4083
$PTA\ Import\ Share_{it-1}$	384	0.1154	0.1633	0	1
$(Import\ Unit\ Value)_{it}$	384	5.0353	5.4327	0.1833	33.0312
$(Import\ Unit\ Value)_{it-1}$	384	5.0750	5.3167	0.1833	31.4695
$\log(Import\ Unit\ Value)_{it-1}$	384	1.0312	1.2124	-1.6967	3.4490
$(Scale\ Measure)_{it}$	384	3687.809	12,502.47	145.8087	125,147.9
$(Scale\ Measure)_{it-1}$	384	3345.394	11,145.26	145.8087	125,147.9
$\log(Scale\ Measure)_{it-1}$	384	7.1966	1.1196	4.9826	11.7373
$(Capital\ to\ Output\ Ratio)_{it}$	384	0.2626	0.2475	0.0417	1.8187
$(Capital\ to\ Output\ Ratio)_{it-1}$	384	0.2452	0.2264	0.0351	1.7026
$\log(Capital\ to\ Output\ Ratio)_{it-1}$	384	-1.6480	0.6421	-3.3482	0.5321

Table A.4. First-Stage Regressions

	Table 3 - Column 7 Specification
D^{90s}	-0.594***
	(0.103)
Employment Share$_{it\text{-}1}$	3.753
	(4.017)
Labor Cost Share$_{it\text{-}1}$	-9.819***
	(1.525)
PTA Import Share$_{it\text{-}1}$	-0.176
	(0.930)
log(*Import Unit Value*)$_{it\text{-}1}$	0.154
	(0.430)
log(*Scale Measure*)$_{it\text{-}1}$	-0.666***
	(0.162)
log(*Capital to Output Ratio*)$_{it\text{-}1}$	-0.433***
	(0.154)
Constant	8.538***
	(1.358)
Observations	384
R-squared	0.906

F test of excluded instruments: F(3,349) = 9.44, Prob > F = 0.000.
Kleibergen and Paap (2006) statistic for underidentification: Chisq(3)=43.99, P-val=0.000.
Cragg and Donald (1993) Wald F-statistic for weak identification: F-stat=9.44, Stock and Yogo (2005) critical values: 5% = 13.91, 10%=9.08, 20%=6.46.
Andersen and Rubin (1949) Wald test of joint significance of endogenous regressors: F(3,327)=9.11, P-val=0.000.

Notes:
(1) The dependent variable is $\log(z_{it\text{-}1}/\varepsilon_{it\text{-}1})$.
(2) *** p<0.01, ** p<0.05, * p<0.10.
(3) Robust standard errors in parentheses.
(4) All specifications include industry fixed effects that are jointly significant but not reported.
(5) OLS estimation.