

Substitutability and protectionism: Latin America's trade policy and imports from China and India*

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Abstract

This paper examines the trade policy response of Latin American governments to the rapid growth of China's and India's exports in world markets. To explain higher protection in sectors where a large share of imports originates in these countries, we extend the 'protection for sale' model to allow for different – region specific – degrees of substitutability between domestically produced and imported varieties. The extension suggests that higher levels of protection towards China's and India's goods can be explained by the higher substitutability of Chinese and Indian goods between domestically produced varieties. The data supports our model, which performs better than the original 'protection for sale' framework in explaining Latin America's structure of protection.

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China's and India's fast economic growth during the past decade is paralleled by their increased presence in policy discussions throughout Latin America. The success of the two Asian economies is not only looked upon with admiration, but is often accompanied by concerns about the effects that growing trade integration with China (and India to a lesser extent) has on the manufacturing sector throughout the region. Textiles, apparel, shoe manufacturing and toys are amongst the sectors worst hit by international competition.

As imports from China and India grow (see Figure 1),¹ requests for explicit protection are becoming more and more common in many Latin American countries. When at the end of 2005 Brazilian imports of textiles from China surged, local manufacturers officially asked the Brazilian government to limit imports of Chinese silk, velvet and polyester thread by imposing import quotas and/or increasing tariffs. At the same time, they suggested that imports of an additional 70 Chinese products were being closely examined, to determine whether similar measures were to be requested also on them. Around the same time, a communiqué by Argentina's Confederation of Medium Enterprises (CAME) called for not repeating the "mistakes of the nineties, when an 'invasion' of Chinese products destroyed entire sectors of the manufacturing sector."²

Local politicians have not left these calls for help unanswered. After a recent meeting with his Chinese counterpart, the Brazilian Minister for Industry, Development and Commerce Luiz Furlan was quick to highlight that "I made it very clear to Minister Bo Xilai that we will take the legal steps to give the Brazilian industry the right to protect itself".³ In early 2006, and following earlier demands of Brazilian textile manufacturers, Brazil and China signed

an agreement under which China was to limit the export growth of 70 textile products. Notwithstanding their country's privileged access to the US market, Mexican politicians have shown similar feelings and are growing more and more nervous about Mexico's burgeoning trade deficit with China. It is not surprising then that after a recent meeting with Chinese leaders, president Fox was very happy to report that "Today we heard from President Hu his enthusiasm, his help, his support in closing the commercial gap..."⁴

While GATT-WTO bounds in principle do not allow countries to increase protection *vis à vis* China and India's products, most developing countries have bound tariffs well above their applied levels, a situation that de facto enables them to significantly increase protection without violating their GATT obligations.⁵

Similarly, antidumping and safeguard rules are quite lax and these instruments have been often used by both developed and developing countries against Chinese imports.⁶ Even more interestingly, Bown (2010) has shown that Latin American countries impose antidumping measures more intensively on China than on any other import source. In particular, China was the top target of antidumping authorities in Argentina, Brazil and Mexico in the 2002-2004 period (see Table 8.3 in Bown 2010). Moreover, there is evidence that the intensity of the antidumping usage against Chinese imports has increased after the country's accession to the WTO.

Given the substantial degree of flexibility enjoyed by domestic policy makers in implementing trade policies within the WTO rules, we are interested in exploring whether the characterization of China and India as sources of "cheap" and "unfair" imports has led to

increased protectionism on goods that are heavily imported from the two Asian economies.⁷

Our initial analysis indicates that this is indeed the case for Latin American imports. Controlling for time, country, and industry fixed effects and instrumenting the import share of China and India to account for potential reverse causality, we find that on average, tariffs and non-tariff barriers tend to be higher for goods that are heavily imported from China and India. The empirical evidence also suggests that this result is mainly driven by imports from China. This result holds for all sub-regions within Latin America, the only exception being Central America, for which there is evidence of lower levels of protection on goods imported from both China and India.

Motivated by these results, we turn to a more structural explanation of the differences in the levels of protection observed in goods imported heavily from China and India. Taking the ‘protection for sale’ model of Grossman and Helpman (1994) as a starting point, we extend it and incorporate the Armington assumption by allowing for imperfect substitution between domestic and imported varieties of a good. In such a setup, trade policy applies only to the imported variety. However, via the degree of substitutability in consumption between the domestic and the imported variety, the level of protection also affects the equilibrium price of the domestic variety. Explicitly taking this into account, all the relevant payoffs can be expressed in terms of the tariff. Solving the model, the degree of pass-through of trade policy into domestic prices – which in turn depends on the degree of substitutability between domestic and imported varieties – enters multiplicatively in the tariff equation of the extended model.

Estimating the extended model on our sample of Latin American countries, we find that it performs better than the traditional ‘protection for sale’ framework along two important dimensions: first, it explains better the tariff structure of the economies we are considering (in terms of R-squared and a non-nested specification J test); and second, the estimates we obtain for the structural parameters are more realistic than the ones obtained by the existing literature. In fact, most existing applications of the ‘protection for sale’ approach estimate the weight governments put on industry lobbying at levels representing less than 1 percent of the weight attached to social welfare, and this is a well-known problem which has been pointed out for instance by Gawande and Krishna (2004) in their thorough review of the empirical literature on the ‘protection for sale’ model.⁸ The extended model that allows for imperfect substitution between domestically produced and imported goods indicates instead that in our sample of Latin American countries, governments’ weight on industry lobbying is on average 32 percent of the weight governments attaches to social welfare, and it is as high as 89 percent for Central American countries.

Our extension of the protection for sale model suggests that higher substitutability between domestic and imported varieties leads – *ceteris paribus* – to higher trade barriers. Are imports from China and India closer substitutes for Latin American domestic output than imports from the rest of the world? To answer this question we introduce in the second half of our empirical analysis a two-tier utility function, which allows for different elasticities of substitution between the domestic variety and imports from different regions of the world. Interestingly, we find that China’s imports are closer substitutes to domestically produced

goods than imports from the rest of the world, whereas for India the evidence is less clear cut.

This is indeed consistent with the fact that Latin America's private sector and policy makers are relatively more concerned about China's growing presence than India's. Recent estimates by Calderón (2008) also suggest that the correlation of output between China and LAC is generally higher than for India and LAC. Moreover, 60 percent of the explained variation in output correlation is attributed to time effects, suggesting that China and LAC tend to be affected by similar exogenous shocks. This provides indirect evidence that China produces goods that are closer substitutes to LAC goods than the ones produced by India.

The imperfect substitutability of imported and domestic varieties in the context of the protection-for-sale model has been introduced first by Chang (2005). In that paper, the author develops a framework featuring Dixit-Stiglitz like differentiated goods sectors and analyzes the effects that this market structure has on the trade policy outcome of the lobbying game. As Chang (2005) points out, her framework is ideally suited to study the intra-industry trade flows that dominate North-North trade. In our theoretical model we instead stop short of such a change in the market structure, because we are interested in South-South and North-South trade. Furthermore, and more importantly, we want to allow for different elasticities of substitution vis-à-vis different source countries, a generalization that cannot be easily introduced in a Dixit-Stiglitz framework. For these reasons, we use a simpler, perfectly competitive setup that, while foregoing the rent shifting effects of Chang's model, allows us to establish unambiguously the effect of the elasticity of substitution on trade policy.

The rest of the paper is organized as follows. Section 2 provides some *prima-facie* evidence regarding Latin American tariffs on goods heavily imported from China and India. In section 3 we develop the extension to Grossman and Helpman's (1994) "protection for sale" model. Section 4 presents the empirical methodology and results. Section 5 concludes.

1 Do imports from China and India face higher average trade barriers in LAC?

To answer this question we start by exploring the correlation between Latin America's structure of protection and the relative importance of China and India as a source of imports. This exercise is undertaken at the highest level of disaggregation that is possible for trade data to be internationally comparable: the six digit level of the Harmonized System (HS). We consider the 1992-2004 period and the country coverage and data sources are discussed in the Appendix (Section S1.2 in the supplemental material, available at <http://wber.oxfordjournals.org/>).

During this period, Latin America's *average* import-weighted MFN tariffs applied on world's imports is 13 percent. The import-weighted applied MFN tariff on goods imported from China and India is 9 percent higher.⁹ The largest protectionist bias against China and India is to be found among Central American and Andean countries, with average levels of protection that are 66 and 26 percent higher, respectively, on imports from China and India than on imports from the rest of the world. But tariffs are only part of the

story. Anti-dumping duties, quantitative restrictions and technical regulations have become an important and often arbitrarily used instrument for trade protection. Latin America's import-weighted overall level of protection (i.e., including ad-valorem equivalents of non tariff barriers) on overall world's imports is 27 percent (Kee, Nicita, and Olarreaga 2009), and on imports from China and India 10 percent higher. The largest protectionist bias against China and India, once we include ad-valorem equivalents of non tariff barriers, is to be found in the Southern Cone, with average levels of protection 20 percent higher on imports from China and India than on imports from the rest of the world.

It is important to note that the differences in average applied MFN tariffs are due to an import-bundle composition effect, and not to higher tariffs applied specifically to imports from China and India. Indeed, all Latin American countries were WTO members in the period considered in our analysis and as a result were required to apply MFN tariffs to India, another WTO member. The same is true also for China, even if the country formally entered the WTO only in 2001, as Latin American countries' tariff policy at the time did not distinguish between WTO members and non-members.¹⁰

But one has to be careful before interpreting these averages as evidence that imports from China and India lead to higher MFN tariffs in Latin America. There are two important issues that need to be addressed before we can reach such a conclusion. First, the causal relation could well go in the opposite direction. In other words, higher tariffs may hit harder the less competitive trading partners, and this may lead to a growing share of imports from China and India. This would happen in a Melitz type model for example if Chinese firms are more

productive and better able to overcome the fixed costs of exporting. Higher tariffs in Latin America may lead to higher imports from China relative to other non-Chinese firms because the Chinese firms are more productive.¹¹ Secondly, our correlations might be affected by endogeneity due to omitted variable bias, as the products in which China and India have a comparative advantage might be those in which Latin American countries have the highest protection because of internal political economy forces, that have little to do with imports from either China or India. For example, China and India are likely to have a comparative advantage in unskilled labor intensive industries, and these are the sectors which have the strongest political clout in Latin America.¹²

We address these two problems by instrumenting the share of imports from China and India using these countries' share in world trade by product – a measure of their success in the global marketplace – as well as the capital-labor ratio of the United States in each industry – a measure of their source of comparative advantage. We also introduce country, year and sector fixed effects to further address the omitted variable bias.

The logic is as follows. First of all, we expect measures of the success and of the drivers of China and India exports in the global market to be correlated with the share of the two countries in Latin American imports by sector. Secondly, as individual Latin American countries are 'small' – none of them represents more than 2 percent of world trade – we expect their trade policy not to affect China and India's competitive position in the world market or its drivers (exclusion restriction). Still we are concerned that the same determinants of China's and India's comparative advantage might be driving the political economy of

trade policy in Latin America. For example, it is well known that unskilled-labor intensive sectors tend to enjoy higher levels of protection. To deal with this issue, we introduce sector fixed effects in the tariff equation. First, we consider two-digit HS fixed effects that vary by country and year to capture sector specific protectionist forces that are common across sectors but that vary by country and year.¹³ Additionally, we are concerned that both the sources of comparative advantage and of protectionism might be operating at a more disaggregated level, and for this reason we also introduce a full set of six digit fixed effects.

Thus, the equation to be estimated takes the following form:

$$t_{k,c,t} = \beta_0 + \beta_k + \beta_I I_{k \in 2 \text{ digit},c,t} + \beta_m m_{k,c,t} + \beta_S s_{k,c,t} + \mu_{k,c,t} \quad (1)$$

where $t_{k,c,t}$ is the level of protection on good k (at the six digit level of the HS) in country c at time t , β_k are six digit HS fixed effects, $I_{k \in 2 \text{ digit},c,t}$ are a full set of product fixed effects (at the level of the 2 digit HS) that vary by country and year, $m_{k,c,t}$ are imports and $s_{k,c,t}$ is the share of imports that comes from China and India in sector k of country c at time t ; $\mu_{k,c,t}$ is a mean zero error term when we assume that there are no six-digit time and country invariant determinants of tariffs.¹⁴ We use two specifications. In the first we include the overall share of imports from China and India, while in the second we introduce the share of imports from China separately.

The instrumental variable results are reported in Tables 1 and 2 for a pool of 10 Latin American countries, and four sub-regions: Andean countries (Bolivia, Colombia, Peru and Venezuela), Central America (Costa Rica and Guatemala), Mexico, and the Southern Cone

(Argentina, Brazil and Uruguay). Tables 1 reports results using tariffs as the left-hand-side variable, and Table 2 reports results using ad-valorem equivalents of non tariff barriers as well. Because the ad-valorem equivalents are only available for the year 2001, there is no time variation in the results reported in Table 2.¹⁵

For Latin America as a whole, our estimates in the first column of Table 1 suggest that an increase of 10 percent in the share of a China and India in the imports in a given sector translates in a 2.6 percent increase in the average ad valorem applied MFN tariff. Recalling that by end of our period the average import share from China and India is 10 percent and that the average MFN tariff on imports from all countries is 13 percent (Table S1.3 in the Appendix), our estimation suggests that doubling up the importance of China and India in Latin America's trade would lead to an increase of approximately 20 percent in the average MFN tariff.¹⁶

The same results continue to hold when we look also at different subgroups of Latin American countries, with the exception of Central America. The latter finding is somewhat surprising as Central American countries were among those for which we found a large bias against Chinese and Indian imports (see Table S1.3 in the Appendix), and highlights the importance of introducing different layers of fixed effects in our estimation. In fact, the strong average bias against Chinese and Indian imports observed among central American countries may be simply due to the fact that politically strong sectors in this region are those in which China has a strong comparative advantage, but this does not imply that Central America has responded to the rapid growth in imports from China and India with stronger

levels of protection.

As it turns out, our results are mainly driven by China. Indeed, when we run a separate regression with imports from China only, we obtain parameter estimates which are not statistically different from the ones reported when we include the sum of imports from both China and India. This result highlights the relative importance of these two countries as a source of imports for Latin America, which may explain why the protectionist bias towards goods imported from China is larger than the protectionist bias towards goods imported from India.¹⁷

Does the pattern of higher protection applied to Chinese and Indian goods hold for non tariff barriers as well? The answer is positive and the results are reported in Table 2, where we use the same specification as in Table 1, but add to the six digit Harmonized System tariffs the ad-valorem equivalents of non tariff barriers obtained by Kee, Nicita, and Olarreaga (2009).¹⁸ The statistical significance remains as high as for those in Table 1 and the same pattern is present. Note that we do not have results for Central American countries, because there are no estimates available for the trade restrictiveness of their non-tariff barriers.¹⁹

In sum, sectors characterized by a larger share of imports from China and India tend to receive higher protection, and this is mainly driven by imports from China, which are several times larger than imports from India.²⁰ Given these results, we turn to a more structural explanation of the differences in the level of protection observed in goods imported heavily from China and India. To that end we extend the Grossman and Helpman (1994) ‘protection for sale’ model to allow for imperfect substitution between domestically produced goods and

imported goods. We then take the extended model to the data.

2 Introducing imperfect substitution in the protection for sale model

To analyze the political economy consequences of increased commercial ties with emerging economies such as China and India, we consider a model in which a small open economy sets trade policy vis-a-vis imports from the rest of the world (ROW). The key hypothesis in our model is that goods are differentiated by location of origin, that is, we adopt the Armington assumption and regard imports and domestically produced varieties as imperfect substitutes. Our model features $n + 1$ different types of goods, and we allow each type to be either domestically produced or imported from abroad.²¹

Indicating by subscript k the type of good, consumers in the home country maximize the following quasi-linear utility function:

$$U = X_0 + \sum_{k=1}^n U_k(X_k) \quad (2)$$

where $U_k(\cdot)$ are strictly concave sub-utility functions ($U_k = E_k \ln X_k$, that is, an upper tier Cobb-Douglas) that depend on a CES aggregate of the imported and domestic variety of the good, denoted with subscripts d and i respectively, i.e.

$$X_k = [x_{k,d}^{\rho_k} + x_{k,i}^{\rho_k}]^{\frac{1}{\rho_k}} \quad 0 < \rho_k < 1 \quad (3)$$

where $x_{k,d}$ stands for the consumption of the domestic variety of good $k \in \{1, \dots, n\}$, $x_{k,i}$ is the consumption of the imported variety, $\sigma_k = \frac{1}{1-\rho_k} > 1$ is the elasticity of substitution between the two varieties, and good zero is the numéraire. Note that quasi-linearity implies that there is neither an income nor a substitution effect for non-numéraire goods, as it is standard in the protection for sale model.

The supply side is a specific-factor model where the primary inputs are sector-specific capital and mobile labor. Each individual in this economy is endowed with labor and at most one sector-specific input. The specifics of supply in each sector are summarized by profit functions $\pi_k(p_{k,d})$, where $p_{k,d}$ is the price of the domestic variety. To keep the analysis tractable, we are going to work with linear supply schedules, i.e. we will assume that the profit functions are quadratic. Production of good zero uses only labor under constant returns to scale, and by appropriate choice of unit its price as well as the wage rate are normalized to one.

For an individual with income E the maximization of utility function (2) subject to the budget constraint $E = X_0 + \sum_{k=1}^n (p_{k,d}x_{k,d} + p_{k,i}x_{k,i})$ yields the following demands for the domestic and imported varieties of each product:

$$x_{k,d}(p_{k,d}, p_{k,i}, E_k; \rho_k) = \frac{E_k p_{k,d}^{\frac{1}{\rho_k-1}}}{p_{k,d}^{\frac{\rho_k}{\rho_k-1}} + p_{k,i}^{\frac{\rho_k}{\rho_k-1}}} \quad (4)$$

$$x_{k,i}(p_{k,d}, p_{k,i}, E_k; \rho_k) = \frac{E_k p_{k,i}^{\frac{1}{\rho_k-1}}}{p_{k,d}^{\frac{\rho_k}{\rho_k-1}} + p_{k,i}^{\frac{\rho_k}{\rho_k-1}}} \quad (5)$$

where $p_{k,i} = p_k^* + t_k$ is the price of the imported variety that results as the sum of the exogenous world market price and the import tariff, and E_k is the expenditure on good k (see the parameter of the Cobb-Douglas above). Note that — in line with a substantial part of the literature and in view of the goal of this paper — we do not explicitly consider export policies.

The price of the domestic variety results from the interplay of domestic supply and domestic demand, where the latter varies not only with the price of the domestic variety but also with the price of the imported variety, and this relation depends on the degree of substitutability. In particular, setting demand equal supply in the market for the domestic variety, i.e. $x_{k,d}(p_{k,d}, p_{k,i}, E_k; \rho_k) = \pi'(p_{k,d})$, implicitly defines the equilibrium price of the domestic variety

$$p_{k,d} \equiv p_{k,d}(p_k^* + t_k; \rho_k) \tag{6}$$

as a function of the price of the imported variety, where the relationship depends on the elasticity of substitution. To obtain further insights into the relationship between the price of the domestic variety and the price of the imported variety, and on how it is influenced by the elasticity of substitution, we assume that the supply of the domestic variety takes the following linear form:

$$y_{k,d} = p_{k,d} \tag{7}$$

Setting supply equal demand in the market for the domestic variety then results in the following equilibrium condition:

$$p_{k,d} = \frac{E_k p_{k,d}^{\frac{1}{\rho_k - 1}}}{p_{k,d}^{\frac{\rho_k}{\rho_k - 1}} + p_{k,i}^{\frac{\rho_k}{\rho_k - 1}}} \quad (8)$$

Since we are unable to explicitly solve for $p_{k,d}$, we proceed by totally differentiating the equilibrium condition. Keeping in mind that the demand function is given by equation (4) above, we obtain

$$dp_{k,d} - \frac{\partial x_{k,d}}{\partial p_{k,d}} dp_{k,d} - \frac{\partial x_{k,d}}{\partial p_{k,i}} dp_{k,i} - \frac{\partial x_{k,d}}{\partial \rho_k} d\rho_k = 0 \quad (9)$$

We are interested in analyzing the relationship between the price of the domestic and foreign varieties, i.e. $\frac{dp_{k,d}}{dp_{k,i}}$. Holding ρ constant, equation (9) implies:

$$\frac{dp_{k,d}}{dp_{k,i}} = \frac{\frac{\partial x_{k,d}}{\partial p_{k,i}}}{1 - \frac{\partial x_{k,d}}{\partial p_{k,d}}} \quad (10)$$

where

$$\frac{\partial x_{k,d}}{\partial p_{k,i}} = -\frac{\frac{\rho_k}{\rho_k - 1} E_k (p_{k,d} p_{k,i})^{\frac{1}{\rho_k - 1}}}{[p_{k,d}^{\frac{\rho_k}{\rho_k - 1}} + p_{k,i}^{\frac{\rho_k}{\rho_k - 1}}]^2}, \quad (11)$$

and

$$\frac{\partial x_{k,d}}{\partial p_{k,d}} = \frac{E_k p_{k,d}^{\frac{2}{\rho_k - 1}} (1 - \rho_k + (p_{k,d}/p_{k,i})^{\frac{-\rho_k}{\rho_k - 1}})}{[p_{k,d}^{\frac{\rho_k}{\rho_k - 1}} + p_{k,i}^{\frac{\rho_k}{\rho_k - 1}}]^2 (\rho_k - 1)}, \quad (12)$$

It is easy to show that both the numerator and the denominator are positive, since $0 < \rho_k < 1$.

We have thus established that $\frac{dp_{k,d}}{dp_{k,i}} \geq 0$, i.e. that the price of the domestic variety increases if the price of the imported variety does, for example because of an increase in the tariff.

How does a change in the substitutability between the two varieties affect the relationship between the price of the domestic and the imported varieties? First, consider two extreme cases at each end of the spectrum. If the elasticity of substitution between the domestic and the imported variety equals one ($\rho_k = 0$), we are in the case of a Cobb-Douglas aggregator. In this case, the price of the domestic variety is unaffected, in other words $dp_{k,d}/dp_{k,i} = 0$. On the other hand, if the domestic and the imported varieties are perfect substitutes ($\rho_k = 1$), a change in the price of the imported variety translates one-for-one ($dp_{k,d}/dp_{k,i} = 1$) into the price of the domestic variety. This puts us back in the standard framework of the Grossman-Helpman protection for sale model.²²

To analyze intermediate cases, and to show more formally that $dp_{k,d}/dp_{k,i}$ is increasing in ρ , we need to differentiate equation (10) with respect to ρ :

$$\frac{\partial\left(\frac{dp_{k,d}}{dp_{k,i}}\right)}{\partial\rho_k} = \frac{\frac{\partial\left(\frac{\partial x_{k,d}}{\partial p_{k,i}}\right)}{\partial\rho_k}\left(1 - \frac{\partial x_{k,d}}{\partial p_{k,d}}\right) + \frac{\partial x_{k,d}}{\partial p_{k,i}} \frac{\partial\left(\frac{\partial x_{k,d}}{\partial p_{k,d}}\right)}{\partial\rho_k}}{\left(1 - \frac{\partial x_{k,d}}{\partial p_{k,d}}\right)^2} \quad (13)$$

Solving explicitly, this can be shown to be positive (see section S1.1 of the Appendix), as long as demand and supply of the domestic variety do not diverge too much.²³ We can therefore conclude that $p'_k \equiv dp_{k,d}/dp_{k,i}$ is a positive function of ρ_k .

2.1 Lobbying Game

We model the lobbying game along the lines of Grossman and Helpman’s familiar “protection for sale” model. In the first stage, owners of sector specific capital in the home country lobby the government for advantageous trade policies on imported substitutes. In particular, they offer contribution schedules $C_k(\mathbf{t})$ that depend on the full vector of import tariffs. Each individual’s consumer surplus is $CS(\mathbf{t}) = \sum_k [U_k(X_k(p_{k,d}, p_{k,i})) - p_{k,d}(p_k^* + t_k)x_{k,d} - (p_k^* + t_k)x_{k,i}]$ and receives also a lump sum transfer from the government representing a share of the total tariff revenues $TR(\mathbf{t}) = \sum_k t_k x_{k,i}(p_{k,d}(p_k^* + t_k), p_k^* + t_k)$, that are rebated to the public on an equal, per capita basis. Note that both these components depend on the price of the domestic variety, and we have made use of expression (6) to express them in terms of tariffs.

Assuming that the ownership of the specific factor is highly concentrated in the population, and in particular that the factor owners represent a negligible fraction of the total population, the objective function of each organized group can be approximated by

$$W_k(\mathbf{t}) = l_k + \pi_k(\mathbf{t}) \tag{14}$$

where l_k is the total labor supply (and also labor income since $w = 1$) of the owners of the specific input used in industry k . In the second stage, each government chooses trade policy and collects the contributions that were offered. Formally, it seeks to maximize the following objective function:

$$G(\mathbf{t}) = \sum_k C_k(\mathbf{t}) + aW(\mathbf{t}) \quad (15)$$

where \mathbf{t} is the vector of tariffs applied by the Home country and a is the weight the government puts on social welfare in its objective function. $W(\mathbf{t})$ denotes the aggregate social welfare function, which is defined as follows:

$$W(\mathbf{t}) = L + \sum_k \pi_k(\mathbf{t}) + CS(\mathbf{t}) + TR(\mathbf{t}) \quad (16)$$

where L denotes the labor force.

We use two additional assumptions to solve for the optimal tariff. First, we follow most of the literature and focus on contribution schedules which are differentiable (i.e. truthful) around the equilibrium point. Second, we assume that all sectors are politically organized.²⁴ As Mitra, Thomakos, and Ulubasoglu (2006) have argued, at the 4 digit SIC level, all sectors of the US economy can be considered politically organized. Thus, although we do not have information on political organization by sector in Latin America (because there is no legal requirement for public disclosure of industries' political contributions), since our structural analysis is carried out at the more aggregate 3 digit ISIC level, this assumption is likely to be harmless.²⁵

Taking the first order condition of the government's maximization problem in (15), and rearranging we obtain:

$$\frac{t_k^0}{1 + t_k^0} = \frac{1}{a} \times \frac{z_k}{\epsilon_k} \times p'_k \quad (17)$$

where $t_k^0 \equiv t_k/p_{k,i}$ is the ad valorem tariff, $z_k \equiv x_{k,d}/x_{k,i}$ is the inverse import penetration ratio, and ϵ_k is the total price elasticity of import demand that consists of the direct price effect and the cross-price effect due to the tariff's impact on the domestic price.²⁶ The last term is the main innovation vis-à-vis the standard model (p'_k is given by equation (10)). We have shown above that it depends positively on the elasticity of substitution. Thus in the presence of high substitution between domestically produced goods and imported goods, tariffs are likely to be higher.

3 Empirical Analysis

To assess the ability of our model to explain the patterns of protection towards Chinese and Indian imports we have highlighted in our preliminary data analysis, we proceed in two steps. First, we compare the performance of our model against the standard Grossman and Helpman benchmark. If product heterogeneity is important, we expect our model to fit better the data than the standard benchmark. We focus on this in section 4.1.

Second, if the model performs better, in order to help us explain higher levels of protection towards imports from China and India, we need to investigate whether the elasticity of substitution between domestic products and Chinese and Indian imports is higher than the elasticity of substitution between domestic products and imports from the rest of the world.

This analysis is carried out in section 4.2. Due to constraints on data availability, all of our structural empirical investigations are carried out at the three-digit SIC level.²⁷

3.1 Does the extended model perform better?

To analyze the importance of allowing for imperfect substitutability between domestic and imported goods, we will compare the performance of the traditional ‘protection for sale’ model with homogeneous goods, with our extended model on the sample of Latin American countries: the extended model is provided by equation (17) and the traditional model is the one where the last term in equation (17) is not present. We will then explore which of the two models better explains Latin America’s tariff structure by comparing the R-squared of the two regressions and by applying Davidson and Mackinnon (1981) non-nested J-test.²⁸

We will also assess the two specifications in terms of their economic significance. One problem with the empirical literature on the ‘protection for sale’ model is that the estimates obtained for a , the parameter describing the weight attached by the government to aggregate welfare, are unreasonably high (see Gawande and Krishna 2004 for a survey of the empirical literature). According to the existing estimates of the traditional ‘protection for sale’ model with homogeneous goods, the weight attached by the government on industry lobbying when setting trade policy represents between 1 and 2 percent of the weight the government puts on social welfare. This is hardly consistent with observed behavior and tariff structures. If we were to obtain a lower estimate for a when bringing our extended model to the data, we would have evidence suggesting that our framework provides more realistic estimates for

this key parameter.

Note that in order to estimate the extended model, we need an estimate for p'_k , the derivative of domestic prices with respect to the price of the composite imported good, which we have defined in (10). Substituting in (10) the derivative of domestic demand with respect to the price of the domestic (equation 11) and of the foreign variety (equation 12), we obtain the following

$$p'_k = \frac{dp_{k,d}}{dp_{k,i}} = \frac{-\frac{\frac{\rho_k}{\rho_k-1} E_k(p_{k,d} p_{k,i})^{\frac{1}{\rho_k-1}}}{[p_{k,d}^{\frac{\rho_k-1}{\rho_k}} + p_{k,i}^{\frac{\rho_k-1}{\rho_k}}]^2}}{1 - \frac{E p_{k,d}^{\frac{2}{\rho_k-1}} (1 - \rho_k + (p_{k,d}/p_{k,i})^{\frac{-\rho_k}{\rho_k-1}})}{[p_{k,d}^{\frac{\rho_k-1}{\rho_k}} + p_{k,i}^{\frac{\rho_k-1}{\rho_k}}]^2 (\rho_k - 1)}} \quad (18)$$

To be able to estimate p'_k we first need data on the prices of the domestic and the composite imported good, as well as consumer expenditure in sector k . The relative price between the domestic good and the composite imported good is obtained using the two first order conditions of the consumer maximization problem. More precisely, we take the ratio between (4) and (5) and solve for the relative price. The quasi-linear structure of the theoretical framework implies that there are neither income nor substitution effects for non-numeraire goods. Thus, the model allows us to assume that the price of the imported good in every sector is equal to one. Consumption is readily obtained from trade and production data.²⁹ We use then equations (4) and (5) to calculate the price of the domestic varieties.

To compute p'_k we will also need an estimate of ρ_k . This can be easily obtained by noting that the absolute value of the price elasticity of import demand is equal to $1/(1 - \rho_k)$.³⁰ Using recent estimates of price elasticities of import demand (Kee, Nicita, and Olarreaga

2008), we solve the equality above to obtain $\rho_k = (\epsilon_{k,i} - 1)/\epsilon_{k,i}$.

One concern we have when constructing p'_k in this fashion is measurement error. Moreover, a well-known problem with the estimation of the ‘protection for sale’ model is the endogeneity of the right-hand-side variables. In order to correct for these issues we decide to use as instruments China and India’s share in world trade by product, and the capital-labor ratio of the United States in each industry. Our strategy consists in instrumenting the term p'_k jointly with the rest of the right hand side term to address these problems.³¹

Table 3 provides the results of the estimation of equation 17 and of the traditional ‘protection for sale model’ for the entire pooled sample and the four sub-regions we are considering. Results for the pooled sample, and for the Southern Cone, always have the expected positive sign on the coefficient of the GH term and of the extended GH term. For Andean countries, Central America, and Mexico the coefficient is negative when using the traditional GH specification, which is at odds with theory. This coefficient is instead positive for the extended GH specification, a result that is consistent with our theoretical predictions. In fact, in all cases the extended GH coefficient has the expected sign and it is statistically different from zero.

Our results also suggest that the extended GH model performs better in terms of R-squared, indicating that the extended model – that allows for imperfect substitutability between domestic and imported varieties – fits better the data. The Davidson-McKinnon non-nested J test for model specification indicates that the extended GH model dominates the model with homogenous goods using either the pooled sample or the data by sub-regions.

As can be seen from (17) the coefficient of the GH term (both in its traditional and extended form) is represented by $1/a$, i.e., the inverse of the weight the government puts on social welfare relative to industry lobbying when setting trade policy. In the case of the extended GH model we obtain estimates for this parameter that are all positive (as expected) and statistically different from zero.

More interestingly, the estimates of the weight governments put on welfare relative to industry lobbying (a) are more realistic than the figures obtained in the existing literature. In fact, for the extended GH model, the estimated parameter a ranges between 1 and 5, rather than between 900 and 1600 (or even negative) as in the traditional GH model. Allowing for imperfect substitution between domestically produced goods and imported goods thus provides one possible solution to the puzzle of large estimates of a .³² In fact, the results from the traditional GH model would suggest that the relative weight the government puts on industry lobbying is around 0.1 percent of the weight put on social welfare for the pooled LAC sample ($0.001 = 1/[a = 918]$). With such a weight, assuming no other market imperfections, it would be very difficult to explain the high levels of trade protection that can be observed in Latin America. On the other hand, the estimates from the extended GH model suggest that the weight attached by the government on industry lobbying is 32 percent of the weight it puts on social welfare ($0.32 = 1/[a = 3.16]$), suggesting a much larger scope for lobbies' influence. Governments with the lowest concern for social welfare are to be found in Central America (where a is estimated at 1.12), whereas Andean countries have governments with the highest concern for social welfare, with an average a estimated at 4.73.

We undertook several robustness checks. First, we estimate (17) and the traditional ‘protection for sale model’ using the overall level of protection that includes ad-valorem equivalents of non-tariff barriers as the left hand side variable. Results for the pooled sample suggest that the parameter a equals 898 in the case of the traditional GH model, and is equal to 2.17 in the case of the extended GH model.³³ Second, one could be concerned with the assumption that all sectors are politically organized, and therefore we decided to check for the robustness of our results using a specification where we allow for the existence of sectors that are politically not organized. To do this we have decided to follow a simple rule where we classify sectors as not organized whenever the levels of protection fall below what we predicted with our theoretical model assuming all sectors to be organized (i.e., sectors that are not politically organized are those where the error term in the estimates reported in Table 3 was negative).³⁴ Results are remarkably robust. The estimates for the average a in Latin America when using the basic GH specification is 333, whereas our extended model provides an estimate of a around 2. When we run these regressions at the regional level, we found that in the case of the Southern Cone the estimates a is reduced from 714 in the basic GH model to 5 in the extended GH model. In the case of Mexico the estimate of a falls from 50 to 1. In Central America from 77 to 1, and among Andean countries from 200 to 3.

Table 4 shows the first-stage regression results which were used to obtain the results displayed in table 3. The F-statistics indicate that the instrumental variables used to estimate the traditional GH model are statistically significant for the pooled sample, for the Andean and for the Mercosur countries, and for Mexico. There is no evidence that the instrumental

variables are suitable in estimating the traditional GH model for Central American countries. The F-test indicates that the instrumental variables used to estimate the extended GH model are jointly significant for the pooled sample and for Andean, Central American, and Mercosur countries. In most cases, the F-statistics is greater than 10 in the extended GH model's first stage regressions. Since we have a single endogenous regressor, these results reinforce our belief in the appropriateness of the instrumental variables used to estimate the extended GH model.³⁵

3.2 Estimating the substitutability between domestically produced goods and imports from China and India

If the better performance of the extended model were to explain higher levels of protection towards Chinese and Indian products, one would need to observe a higher degree of substitutability between domestically produced goods and goods imported from China and India than between domestic and foreign varieties imported from the rest of the world.

To measure the degree of substitutability between imports from different sources and LAC's domestically produced goods we follow the strategy described in Sato (1967) and disaggregate the imported variety using a two-tier CES utility structure. In other words, starting from the utility function in equation (2), remember that the aggregate good k was represented by

$$X_k = [x_{k,d}^{\rho_k} + x_{k,i}^{\rho_k}]^{\frac{1}{\rho_k}} \quad 0 < \rho_k < 1 \quad (19)$$

where $x_{k,d}$ is the domestic variety, while $x_{k,i}$ is the imported variety and $\sigma_k = \frac{1}{1-\rho_k}$.

Denote now imports of good k from country j by $x_{k,i,j}$, where j is an element of the set $\{\text{China, India, Rest of the World}\}$. Let

$$x_{k,i} = \left[\sum_j \phi_j x_{k,i,j}^{\gamma_k} \right]^{\frac{1}{\gamma_k}} \quad (20)$$

where $\phi_j > 0$, and $\sigma_{k,i} = \frac{1}{1-\gamma_k}$ denotes the elasticity of substitution among imported varieties of good k . Denote by $\sigma_{k,j}$ the elasticity of substitution between the domestically produced variety of good k and the variety of good k imported from region j . With this nested CES preference structure Sato (1967) shows that the relationship between $\sigma_{k,i}$, σ_k , and $\sigma_{k,j}$ is given by:

$$\frac{\left(\frac{1}{\theta_{k,j}} - \frac{1}{\theta_{k,i}} \right) + \left(\frac{1}{\theta_{k,d}} - \frac{1}{\theta_{k,i}} \right)}{\sigma_{k,j}} = \frac{\left(\frac{1}{\theta_{k,j}} - \frac{1}{\theta_{k,i}} \right)}{\sigma_{k,i}} + \frac{\left(\frac{1}{\theta_{k,d}} + \frac{1}{\theta_{k,i}} \right)}{\sigma_k} \quad \text{for } j \in \{C, I, ROW\} \quad (21)$$

where $\theta_{k,j}$ is the share of total expenditure on the imported variety of good k from region j , $\theta_{k,i}$ is the share of total expenditure on imports of that good (i.e., $\theta_{k,i} = \sum_j \theta_{k,j}$), and $\theta_{k,d}$ is the share of total expenditure on the domestic variety of good k .

Using equation (5) we can derive the price elasticity of the composite of imported goods, ϵ_k . Solving for σ_k we have $\sigma_k = -\epsilon_k$. Thus, with an estimate of the price elasticity of the imported composite good k that we can borrow from the existing literature, we can derive an estimate for σ_k . With data on σ_k and on the share of expenditure on domestic and on

imported varieties we can use the relationship described in equation (21) to obtain estimates for the degree of substitutability between domestically produced goods and respectively imports from China, India and the rest of the world.³⁶

Before bringing equation (21) to the data, notice that the shares of expenditure on domestic and imported varieties appear in the left and right-hand sides of equation (21). Therefore, we need to rearrange the expression to be able to estimate the parameters of interest. As a result, the equation to be estimated becomes:

$$\frac{1}{\sigma_k} = \alpha_{1,j} \frac{\left(\frac{1}{\theta_{k,d}} - \frac{1}{\theta_{k,i}}\right)}{\left(\frac{1}{\theta_{k,d}} + \frac{1}{\theta_{k,i}}\right)} + \alpha_{2,j} \frac{\left(\frac{1}{\theta_{k,j}} - \frac{1}{\theta_{k,i}}\right)}{\left(\frac{1}{\theta_{k,d}} + \frac{1}{\theta_{k,i}}\right)} + \varepsilon_k \text{ for } j \in \{C, I, ROW\} \quad (22)$$

where $\alpha_{1,j} = \frac{1}{\sigma_j}$ and $\alpha_{2,j} = \left(\frac{1}{\sigma_j} - \frac{1}{\sigma_i}\right)$ are the parameters of interest, while ε_k is a zero mean error term that captures measurement errors in the dependent variable. The parameter σ_j can be estimated by calculating $1/\alpha_{1,j}$. is the basis for our estimation of the relative degrees of substitutability between the domestically produced and the goods imported from different regions.

We estimate equation (27) using data on imports from China, India, and rest of the world separately. Parameter estimates obtained for equation (27), as well as the implied σ 's are reported in Table 5. A quick look suggests that the degree of substitution between Chinese goods and goods produced in Latin America – as measured by σ_C – is higher than the degree of substitution between goods imported from the rest of the world and Latin American goods – as measured by σ_{ROW} . Indeed, the estimates for σ_C are numerically larger than the estimates for σ_{ROW} in all cases, even if they are not always statistically

different from each other.³⁷ The differences are statistically different from zero at the 10 percent level in the case of the entire sample and for Andean countries. The estimates of the elasticity of substitution between imports from India and domestically produced goods (σ_I) are larger than the estimated σ_{ROW} for all specifications but for the Andean countries and Mercosur, but the differences are never statistically different from zero.

Given that these estimates represent averages across different products, it is interesting to disentangle whether the higher estimates for the elasticity of substitution between domestically produced goods and Chinese imports is due to a composition effect or within product effect. Indeed, the higher elasticity may simply be explained by the fact that China's export bundle is closer to Latin America's production bundle than the export bundle from the rest of the world. We address this issue by introducing product fixed effects in the specifications we run to estimate equation (27) to control for any product specific effect (some products having larger elasticities of substitution than others), and we restrict the sample to those observations where imports from China represent more than 2 percent of total imports. Results are provided in Table S1.6 of the Appendix, and are qualitatively very similar to those obtained in Table 5. This suggests that the higher substitutability between Chinese goods and domestically produced goods is not only due to a composition effect.

To further verify the importance of within industry differences in the elasticities of substitution, we run equation (27) by ISIC 2-digit industries. We find that in those industries where imports from China are relatively important (ISIC 32: textiles & apparel, ISIC 36: pottery and earthenware and ISIC 38: machinery, see Figure 1) the elasticity of substitution

between Chinese and domestically produced goods is always estimated to be larger than the elasticity of substitution between domestic varieties and goods imported from the rest of the world. The same does not apply to imports from India. ³⁸

Finally, turning back to the first stage regressions, the results reported in Table 4 tend to suggest that the larger is the share of China's in world markets the larger is the extended GH-ext term since the coefficient of the share of imports from China is always positive and in most cases is statistically significant. And the GH-ext term contains p'_k which as shown in Section S1.1 of the appendix tends to increase with the elasticity of substitution between domestically produced goods and imported goods. Thus, these first stage regressions also provide indirect evidence that the larger is the share of China in world markets, the larger is the degree of substitution between Latin America's import and domestically produced bundles, which in turn increases protectionist pressures.

4 Conclusion

The growing presence of China and India in world markets has caught the attention of many observers. This paper has explored the response of Latin American policy-makers to the growth of imports from China and India in their markets. We have found that sectors in which the share of imports from China is larger tend to have higher tariffs controlling for reverse causality and industry, year and country fixed effects. In the case of India the evidence is more mixed.

In order to explain this evidence, we have extended the Grossman and Helpman (1994)

‘protection for sale’ model to allow for imperfect substitution between domestically produced and imported goods. The model suggests that as the elasticity of substitution between domestically produced and imported goods increases, the incentives to lobby also increase, and the resulting equilibrium tariff is higher.

Our analysis has been carried out in two steps. First, we have brought our extended framework to the data and have shown that it outperforms the traditional Grossman Helpman framework in two respects: it fits better the data – and even more interestingly – explicitly modeling imperfect substitutability allows us to obtain substantially more realistic structural parameter estimates than those resulting from an homogenous good specification.

Second, given that the extended model performs better, we have measured the substitutability of domestically produced goods with imports from either China, India or the Rest of the World. We have found that Chinese imports are on average closer substitutes to goods domestically produced in Latin America than goods originating in the rest of the world, while the evidence for India is more mixed. This suggests that it is indeed closer substitutability with imports that drives higher tariffs in those sectors where Chinese imports are more relevant.

Our analysis has important implications for the existing empirical literature on the ‘protection for sale model’. By ignoring imperfect substitutability between domestically produced and imported goods, the existing studies have obtained estimates for the weight attached by the government to social welfare that are substantially upward biased. Our analysis suggests that for our entire sample, under imperfect substitutability, the weight attached to industry

lobbying is around 32 percent of the weight attached by the government to aggregate welfare. Under perfect substitutability, it is only 0.1 percent. If the same type of bias were to exist in the estimates of the protection for sale model carried out for the United States, the ‘true’ value of a would be around 10 for the model estimated by Gawande and Bandhopadhyay (2000) (instead of 3000 like in their paper), and 0.23 for the benchmark specification in Goldberg and Maggi (1999). In particular, the latter would imply that the US government puts five times more weight on political contributions than social welfare when determining trade policy. Of course, these are just back of the envelope calculations, and a more careful analysis on US data needs to be undertaken before reaching any definitive conclusion. However, these results do suggest that allowing for imperfect substitutability within the Grossman and Helpman (1994) framework can lead to much more credible parameter estimates.

A possible explanation for this result is that by assuming an infinite elasticity of substitution between domestic and imported varieties, the existing literature has both overstated the benefits and understated the costs of a tariff. Concerning the former, if substitution between domestic and imported varieties is less than perfect, the domestic producer is likely to observe a smaller price increase in his product than the one he was hoping for. As for the latter, under imperfect substitutability, domestic consumers will have a harder time substituting away from the more expensive domestic varieties. Since in the political equilibrium trade policy is determined by the equality between the marginal increase in producer surplus associated with the tariff and the marginal decline in social welfare weighted by the parameter a , it is clear that if the left hand side is upward biased, and the estimate of the change

in social welfare is downward biased, the only way to explain the observed tariff is with a relatively high estimate of a . The absence of perfect substitutability between imported and domestic varieties could therefore explain one of the important puzzles in the empirical political economy literature.³⁹

Notes

¹Imports from China and India represented around 1 percent of total imports in 1992 whereas they reached more than 10 percent by 2004. The increase was particularly impressive in sectors like textile and apparel, in which the share of imports from China and India has grown from 3 to 20 percent over this period.

²See CAME's communiqué of November 16th 2004 at <http://procom.org.ar/comunicado.php3?id=335>.

³As reported by Yahoo! on October 4 2005. See <http://sg.biz.yahoo.com/051004/1/3veny.htm>.

⁴See "Mexico builds trade ties with China" by J.C. McKinley in The New York Times of September 19, 2005.

⁵For example, Brazil's bound tariff in textiles, apparel and footwear are bound at 35% in the WTO, and applied tariffs on these products have varied between 16 and 30 percent during the 1990s.

⁶See Hoekman and Kostecki (2001). See also the recent 35% tariff imposed by the Obama administration on the imports of Chinese tyres.

⁷Below are some common characterizations of China as a source of "cheap" and "unfair" imports: "Countries around the world are bracing for a surge of cheap imports from China, which benefits from cheap, union-free labor and rising productivity" Taipei Times, January 2nd 2005. "And a villain always helps. Our polling indicates that 31% of Americans see China as the country that ignores agreements and breaks rules the most often." Frank Luntz in *Republican Playbook*.

⁸Gawande and Bandhopadhyay (2000), Goldberg and Maggi (1999) and Facchini, Van Biesebroeck, and Willmann (2006) estimated this model for the United States, Mitra, Thomakos, and Ulubasoglu (2002) for Turkey. McCalman (2004) found for Australia that industry lobbying accounts for 2 percent of the Australian government objective function. For a different approach on how to deal with this issue, see Mitra, Thomakos, and Ulubasoglu (2006).

⁹Throughout the paper we work with applied MFN tariffs as reported by each country. A problem with this measure is that it does not capture exemptions due to investment or export regimes (e.g., duty drawback or rebate systems that reduce or eliminate duties paid on intermediate inputs). The alternative consisting of

using effectively collected tariffs is unfortunately not feasible, because we do not have access to customs data on tariff collection by product for the Latin American countries in our sample. Thus, the assumption we are making is that if these tariff exemptions exist, the rent is fully captured by the user of the intermediate good, and this will not affect the lobbying equilibrium. For a full treatment of tariff exemptions, see Cadot, de Melo, and Olarreaga (2003).

¹⁰As discussed before, the MFN bound rates agreed by Latin American countries were substantially higher than the actual applied rates, leading to substantial room for more restrictive policies to be implemented. In particular, Foletti, Fugazza, Nicita, and Olarreaga (2009) find that bound tariffs were on average three times larger than applied tariffs.

¹¹We are grateful to Chad Bown for providing us with this example.

¹²Bown (2010), who finds a positive correlation between Chinese import tariffs and the likelihood of being targeted by an antidumping procedure in the rest of the world, suggests a similar interpretation of his finding.

¹³This will ensure that there is no correlation between the error term in the tariff equation and the instrumented share of imports from China and India that could be due to common drivers of protection at the sectoral level across Latin American countries.

¹⁴The instrumental variables used in this paper vary by product (6 digit of the HS) and year but they do not vary by country. For this reason, we estimate equation (1) using robust standard errors with clustering at the product-year level. Moreover, we estimate the regressions for sub-regions, and compare the results at the regional level with the results at the country level. We could not find significant differences between the estimates at the regional level and at the country level.

¹⁵Most first-stage regressions are highly statistically significant with p-values lower than 0.01. The results of first-stage regressions are available upon request. Notice also that when we use overall measures of protection in Table 2 our we only have data for 2001, and using sector fixed effects at the 6-digit level would prevent us from using our instrumental variables since they are organized at the same level of aggregation.

¹⁶To account for the potentially high degree of inertia in the tariff structure, we have also run the specifications in Table 1 using a dynamic panel estimator (Arellano-Bond). The results are qualitatively not

affected and are available upon request from the authors.

¹⁷We have also experimented estimating equation (1) controlling for the presence of preferential trade agreements and for the accession of China to the WTO. In the former case, we have introduced dummies to control for the presence of preferential agreements. In the latter case, we have also used as an explanatory variable the interaction between the share of imports from China and a dummy variable which is equal to one after the year of 2000. The results described above remain intact and are available upon request. We also run a specification that explains pre-1995 tariffs in Latin America using the share of imports from China and India in 2004. Interestingly, we find that there is no correlation between the tariff observed before 1995 and imports from China in 2004, thus suggesting that our findings do not suffer from reverse causality. The results are reported in Table S1.4 in the appendix.

¹⁸Note that these estimates only exist for the year 2001, so we lost the time dimension in our sample. Results reported in Table 2 are for the year 2001 only. One should keep in mind that the ad-valorem equivalents of NTBs constructed by Kee, Nicita, and Olarreaga (2009) and used in Table 2 do not vary by source country. This is appropriate for most NTBs which are applied on a non discriminatory basis (e.g. sanitary and phytosanitary measures), but is more problematic for other NTBs, like antidumping duties, which are instead inherently discriminatory. Interestingly, Bown (2010) has shown that for Latin America heavy users of antidumping (Argentina, Brazil and Mexico), tend to disproportionately target China with their antidumping investigations.

¹⁹The first-stage regressions used to compute the results in table 2 are in general highly statistically significant. The first-stage results are available upon request.

²⁰Imports from China account on average for 85 percent of total imports from China and India from 1992 to 2004. The share of imports from China ranges between 76 percent and 91 percent of total imports from these two countries over the years under consideration.

²¹Different varieties in a protection for sale model under monopolistic competition have been analyzed by Chang (2005).

²²If both varieties were complementary ($\rho_k < 0$) then we would obtain a negative correlation between both

prices, a case we do not consider as we are modeling two varieties of the same good.

²³A sufficient condition is that $|\ln p_{k,d} - \ln p_{k,i}| < (1 - \rho)/\rho$.

²⁴For a similar approach, see Grossman and Helpman (1995) and Mitra, Thomakos, and Ulubasoglu (2006).

²⁵As discussed at the end of section 4.1 we have also carried out a robustness check in which we have identified unorganized sectors through a simple ‘rule of thumb’. The results are not affected.

²⁶Formally, $\epsilon_k = \epsilon_{k,i} + \epsilon_{k,d}\epsilon_{p,k} = \frac{\partial x_{k,i}}{\partial p_{k,i}} \frac{p_{k,i}}{x_{k,i}} + \frac{\partial x_{k,i}}{\partial p_{k,d}} \frac{p_{k,d}}{x_{k,i}} \times \frac{\partial p_{k,d}}{\partial p_{k,i}} \frac{p_{k,i}}{p_{k,d}}$.

²⁷In particular, to estimate equation we need to construct the share of domestically produced goods in total consumption, and data on production is only available at the 3-digit ISIC level.

²⁸This consists of running the two specifications, taking the predicted value of each specification and adding the predicted of the alternative specification to the null specification. If the predicted value is statistically significant, then we cannot reject that the alternative is the right specification. The problem with this test is that we may not be able to reject either of the alternatives, or we may be able to reject both, i.e., the test may be inconclusive.

²⁹Or rather apparent consumption which equals imports plus domestic production minus exports.

³⁰This assumes that a change in the import price does not affect the overall price index of good k (i.e., the denominator of equation (5)), which leads to the often obtained result with CES preferences that the elasticity of substitution equals the price elasticity of demand.

³¹For a discussion of the use of these instruments, see section 2.

³²Standard errors for the estimated weight attached to social welfare (a) were calculated using the Delta-Method. Table 3 shows estimation results using robust standard errors that are clustered at the product-year level. We follow this strategy since the instrumental variables do not vary by country. Notice that we also estimate the tariff equations that emerge from the Traditional and Extended ‘protection for sale models’ for sub-regions in Latin America, and these results do not differ significantly from the results obtained at the country level.

³³However none of the estimates of the traditional GH model are statistically different from zero (although they are different from each other). In the case of the extended GH model, the estimates for the pooled

sample as well as for Mercosur countries are statistically significant and have the expected sign. The results for the other sub-regions are not statistically significant.

³⁴This rule follows the main line of argument found in Cadot, Dutoit, Grether, and Olarreaga (2007).

³⁵This observation follows from the “rule of thumb” suggested by Staiger and Stock (1997).

³⁶Note though that in our empirical analysis we will only be able to capture the average degree of substitutability σ_C , σ_I , and σ_{ROW} , as we have too few observations to estimate sector specific elasticities.

³⁷Standard errors for the estimated elasticities of substitution were calculated using the delta method.

³⁸Notice however, that none of the differences at the 2-digit ISIC level are statistically significant.

³⁹We would like to thank one of the referees for suggesting this explanation.

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Figure 1: Share of LAC imports from China and India, 1992-2004

(Total and by main product category)

Source: Author's calculations based on the United Nations COMTRADE database.

Table 1: Tariffs in LAC and imports from China and India^a

	Latin America	Andean countries	Central America	Mexico	Southern Cone
Total imports	0.67 ** (0.20)	-0.69 (0.46)	0.27 (0.71)	0.12 (0.14)	0.99 ** (0.38)
$(m_{k,c,t})$					
Share of imports from China & India	0.26 ** (0.06)	0.05 * (0.03)	-0.12 ** (0.05)	0.35 ** (0.14)	0.21 * (0.11)
$(s_{k,c\text{from } C\&I,t})$					
Share of imports from China	0.21 ** (0.05)	0.06 * (0.03)	-0.13 ** (0.06)	0.28 ** (0.11)	0.19 ** (0.09)
$(s_{k,c\text{from } C,t})$					
R-squared	0.54	0.67	0.63	0.54	0.64
F-stat. – first stage	9.49	9.79	5.97	2.06	12.5
F-stat. – orthog.	1.00	0.00	0.01	0.02	0.07
# observations	247822	101010	34215	22676	89921
# countries	10	4	2	1	3

^a *Source:* Author’s analysis based on data described in the online Appendix.

Regressions are estimated using an instrumental variable approach, where all variables are instrumented (OLS results are reported in Table S1.5 of the appendix). Instruments are the shares of China and India in world markets, and the US capital-labor ratio. The F-statistics of the first stage is for a regression of the share of imports from China (or China and India) on the instrumental variables. The F-statistics related to the regression of the error term on the instrumental variables is denoted by "F-stat. – orthog". All regressions include dummies that vary by 2 digit HS sector, year and country, as well as six digit HS fixed effects. White robust standard errors clustered at the product (HS6)-year level are reported in parenthesis. Total imports are measured in \$ 100 million. ** stands for statistical significance at the 5 percent level, and * for statistical significance at the 10 percent level.

Table 2: Overall levels of protection in LAC and imports from China and India^a

	Latin America	Andean countries	Central America	Mexico	Southern Cone
Total imports	1.03 ** (0.27)	1.81 (1.13)	NA	0.67 ** (0.20)	0.57 (0.40)
($m_{k,c,t}$)					
Share of imports from China & India	0.46 ** (0.10)	0.16* (0.09)	NA	2.87 ** (0.65)	0.21* (0.11)
($s_{k,c}$ from C&I,t)					
Share of imports from China	0.41 ** (0.10)	0.12* (0.07)	NA	2.06 ** (0.59)	0.14 ** (0.06)
($s_{k,c}$ from C,t)					
R-squared	0.16	0.25		0.18	0.14
F-stat. – first stage	27.09	65.29		10.27	47.12
F-stat. – orthog.	0.68	0.20		12.11	8.58
# observations	22542	7514		3757	11271
# countries	8	4	0	1	3

^a *Source:* Author’s analysis based on data described in the online Appendix.

Regressions are estimated using an instrumental variable approach, where all variables are instrumented (OLS results are reported in Table S1.5 of the appendix). Instruments are the US capital-labor ratio and the shares of China and India in world markets. The F-statistics of the first stage is for a regression of the share of imports from China (or China and India) on the instrumental variables. The F-statistics related to the regression of the error term on the instrumental variables is denoted by "F-stat. – orthog". All regressions include dummies that vary by 2 digit HS sector and country. White robust standard errors clustered at the 6-digit of the HS are reported in parenthesis. Total imports are measured in \$ 100 million. ** stands for statistical significance at the 5 percent level, and * for statistical significance at the 10 percent level.

Table 3: Estimating the classic and extended Grossman-Helpman (GH) model.^a

	Latin America	Andean countries	Central America	Mexico	Southern Cone
GH term	0.001 **	-0.003**	-0.002	0.002	0.0006**
($z_{k,c,t}/\epsilon_{k,c,t}$)	(0.00)	(0.001)	(0.002)	(0.003)	(0.0001)
Implied a	918 **	-293**	-363	337	1639**
(weight on welfare)	(169.33)	(58.59)	(210.28)	(405.70)	(334.07)
Extended GH term	0.32**	0.21**	0.89**	0.28**	0.32**
($z_{k,c,t}p'/\epsilon_{k,c,t}$)	(0.03)	(0.02)	(0.11)	(0.10)	(0.03)
Extended implied a	3.16**	4.73**	1.12**	3.59**	3.14**
(weight on welfare)	(0.32)	(0.51)	(0.134)	(1.27)	(0.38)
R-squared	0.21	0.29	0.09	0.10	0.20
# observations	1374	532	155	100	587
# countries	10	4	2	1	3
J -non nested test ^b	Ext-GH	Ext-GH	Ext-GH	Ext-GH	Ext-GH

^a *Source*: Author's analysis based on data described in the online Appendix.

Regressions are estimated using an instrumental variable approach, where the GH terms are instrumented using the United States capital labor ratio in that industry, and China and India's shares in world trade by product. All regressions include country and year dummies. White robust standard errors clustered at the product (ISIC3)-year level are reported in parenthesis. Standard errors for the structural parameter 'a' were calculated using the Delta Method. ** stands for statistical significance at the 5 percent level, and * for statistical significance at the 10 percent level.

^b Davidson-McKinnon (1981) J non-nested test for model specification. "GH" stands for the traditional GH model dominating the extended model at the 5 percent level, and "Ext-GH" stands for the extended GH model dominating the traditional GH model at the 5 percent level.

Table 4: First-Stage Results.^a

	Latin America		Andean countries		Central America		Mexico		Southern Cone	
	GH	Ext-GH	GH	Ext-GH	GH	Ext-GH	GH	Ext-GH	GH	Ext-GH
US K/L	3278.49 ** (1195.87)	-8.33 (8.74)	1231.11** (463.61)	-4.22 (8.02)	991.34 (1411.88)	-2.61 (8.50)	-141.94 (585.00)	-47.6 (37.1)	6515.69** (2349.58)	-7.24 (8.56)
China's world share	218.67 ** (66.38)	1.02** (0.27)	17.61 (15.18)	1.17** (0.3)	-25.42 (39.50)	0.51* (0.29)	-2.00 (23.30)	0.77 (0.62)	600.74** (136.92)	1.08** (0.28)
India's world share	-739.48** (171.27)	0.96 (0.92)	-104.97** (38.56)	1.20 (0.97)	-211.40 (150.35)	0.80 (1.03)	-206.90** (80.11)	-2.20 (1.84)	-1570.25** (367.42)	1.04 (1.01)
Constant	4.52* (2.38)	0.04** (0.01)	7.00** (1.43)	0.02 (0.01)	4.72 ** (1.55)	0.02 (0.01)	7.94 ** (2.08)	0.15** (0.05)	3.02 (4.50)	0.06** (0.02)
F	6.69**	16.00**	3.24*	11.03**	2.25	3.08 *	2.98*	2.14	6.76**	16.04**
# observations	1374	1374	532	532	155	155	100	100	587	587
# countries	10	10	4	4	2	2	1	1	3	3

^a *Source:* Author's analysis based on data described in the online Appendix.

The dependent variable is the explanatory variable in Table 3 (i.e., either the GH or the extended GH term). US capital levels (K) are measured in \$ 100 million. White robust standard errors are reported in parenthesis. ** stands for statistical significance at the 5 percent level, and * for statistical significance at the 10 percent level.

Table 5: Estimating the degree of substitutability with domestically produced goods^a

	Latin America	Andean Countries	Central America	Mexico	Southern Cone
$\alpha_{1,ROW}$ (Rest of the world)	0.091** (0.009)	0.087** (0.014)	0.031* (0.023)	0.165** (0.055)	0.115** (0.016)
σ_{ROW} (Rest of the world)	10.989** (1.079)	11.507** (1.842)	32.573 (24.346)	6.064** (2.033)	8.673** (1.179)
$\alpha_{1,C}$ (China)	0.073** (0.007)	0.066** (0.010)	0.021 (0.014)	0.143** (0.047)	0.104** (0.013)
σ_C (China)	13.698** (1.353)	15.174** (2.33)	47.847 (31.212)	6.978** (2.319)	9.588** (1.193)
$\alpha_{1,I}$ (India)	0.088** (0.007)	0.104** (0.011)	0.016 (0.015)	0.068* (0.034)	0.125** (0.014)
σ_I (India)	11.363** (0.899)	9.569** (0.995)	64.102 (63.471)	14.706* (7.307)	7.974** (0.888)
R-squares	[0.14, 0.19]	[0.08, 0.21]	[0.01,0.03]	[0.10,0.15]	[0.16,0.18]
Number of observations	[1499,1899]	[419,603]	[114,162]	[116,130]	[637,753]
Number of countries	10	4	2	1	3

^aSource: Author's analysis based on data described in the online Appendix.

All regressions use country-fixed effects. White robust standard errors are provided in parentheses, both for the coefficients and the implied substitution parameters. ** stands for statistical significance at the 5 percent level and * stands for statistical significance at the 10 percent level. The table provides the range of R-squares and number of observations of the regressions using data on imports from China, India and the rest of the of the world.