Substitutability and Protectionism: Latin America's Trade Policy and Imports from China and India

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This article examines the trade policy response of Latin American governments to the rapid growth of Chinese and Indian exports in world markets. To explain more protection in sectors where a large share of imports originates in China and India, the "protection for sale" model is extended to allow for region-specific degrees of substitutability between domestic and imported varieties of a good. The results suggest that more protection toward Chinese and Indian goods can be explained by the higher substitutability of Chinese and Indian goods with domestic varieties. The data support the model, which performs better than the original protection for sale framework in explaining Latin America's structure of protection. JEL classification numbers: F10, F11, F13

China's and India's fast economic growth during the past decade is paralleled by their increased presence in policy discussions throughout Latin America. Their success is looked on with admiration, accompanied by concerns about the effects that growing trade integration with China (and India to a lesser extent) has on Latin America's manufacturing sector. Textiles, apparel, shoe manufacturing, and toys are among the sectors most affected by international competition.
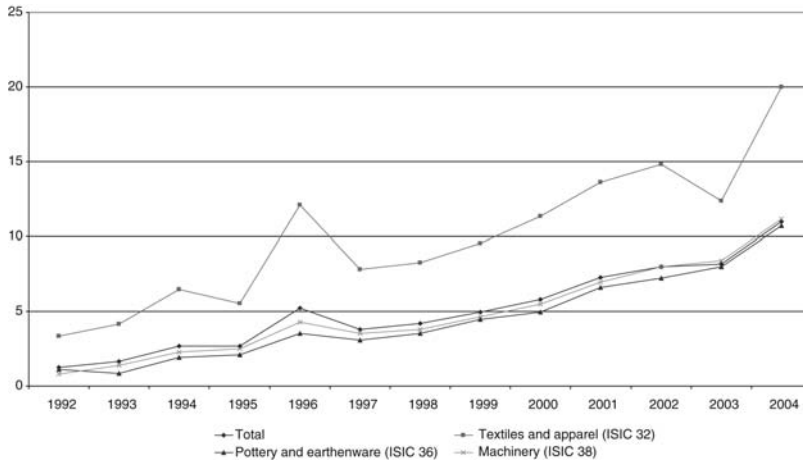
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FIGURE 1. Share of Imports from China and India to Latin America, 1992–2004

Source: Authors' calculations based on the United Nations Statistics Division's Comtrade database.

As Chinese and Indian imports have grown in many Latin American countries (figure 1),[1] requests for explicit protection have become more common. When Brazilian imports of Chinese textiles surged at the end of 2004, local manufacturers asked the Brazilian government to limit imports of Chinese silk, velvet, and polyester thread through import quotas or higher tariffs. The Brazilian government also examined imports of 70 other Chinese products to determine whether similar measures were needed. A 2004 comuniqué by Argentina's Confederation of Medium Enterprises called for not repeating the "mistakes of the nineties, when an 'invasion' of Chinese products destroyed entire sectors of the manufacturing sector" (CAME 2004).

Local politicians have not left these calls for help unanswered. Brazilian Minister for Industry, Development, and Commerce Luiz Furlan was quick to highlight that "I made it very clear to [my Chinese counterpart] Minister Bo Xilai that we will take the legal steps to give the Brazilian industry the right to protect itself" (*Iran Daily* 2005). In early 2006 Brazil and China signed an agreement for China to limit the export growth of 70 textile products. Mexican politicians have shown similar feelings—notwithstanding their country's privileged access to the U.S. market—and are growing more nervous about Mexico's burgeoning trade deficit with China. It is thus unsurprising that after a 2005 meeting with Chinese leaders, Mexican President Vicente Fox was very happy to report that "Today we heard from President Hu his

1. Imports from China and India accounted for about 1 percent of total imports in 1992 and more than 10 percent by 2004. The increase was particularly impressive in textiles and apparel, which grew from 3 percent to 20 percent.

enthusiasm, his help, his support in closing the commercial gap ..." (McKinley 2005).

While General Agreement on Tariffs and Trade/World Trade Organization (GATT/WTO) bound tariffs in principle do not allow countries to increase protection against Chinese and Indian products, most developing countries have bound tariffs well above their applied levels, which enables them to substantially increase protection without violating GATT/WTO obligations.[2]

Similarly, lax rules on antidumping and safeguarding have allowed both developed and developing countries to use these instruments against Chinese imports.[3] Bown (2010) shows 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 over 2002–04 (see table 8.3 in Bown 2010). Moreover, there is evidence that the intensity of antidumping measures against Chinese imports has increased since the country's accession to the WTO.

The substantial flexibility enjoyed by domestic policymakers in implementing trade policies under WTO rules raises the question of whether characterizing China and India as sources of "cheap" and "unfair" imports has increased protectionism on products that are heavily imported from the two economies.[4]

The initial analysis here indicates that protectionism is indeed greater for Chinese and Indian imports to Latin America. Controlling for country, year, and sector fixed effects and instrumenting China's and India's import shares to account for potential reverse causality show that, on average, tariffs and non-tariff barriers are higher for products that are heavily imported from China and India. The empirical evidence also suggests that this result is driven mainly by Chinese imports. This result holds for all subregions within Latin America except Central America, for which there is evidence of less protection on both Chinese and Indian imports.

Motivated by these results, the discussion turns to a more structural explanation of the differences in protection in products heavily imported from China and India. Extending Grossman and Helpman's (1994) "protection for sale" model and incorporating the Armington assumption by allowing for imperfect substitution between domestic and imported varieties of a good yield a setup where trade policy applies only to the imported variety. However, the level of

2. For example, Brazil's bound tariff in textiles, apparel, and footwear are fixed at 35 percent in the World Trade Organization, and applied tariffs on these products have varied between 16 percent and 30 percent during the 1990s.

3. See Hoekman and Kostecki (2001). See also the recent 35 percent tariff imposed by the United States on the imports of Chinese tires in 2009 (*The Economist* 2009).

4. 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* 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" (Luntz 2005).

protection also affects the equilibrium price of the domestic variety through the degree of substitutability in consumption between the domestic and the imported varieties. Explicitly taking this into account, all the relevant payoffs can be expressed in terms of the tariff. In 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 the sample of Latin American countries shows that it performs better than the traditional protection for sale framework along two important dimensions. First, the model better explains the tariff structure of the economies being considered (in terms of R-squared and a non-nested specification J test). Second, the estimates obtained for the structural parameters are more realistic than the ones obtained by the existing literature while assessing the traditional model. In fact, most applications of the protection for sale approach estimate the weight governments put on sector lobbying at 1–2 percent of the weight attached to social welfare, and this is a well known problem (see, for example, Gawande and Krishna's 2004 review of the empirical literature on the protection for sale model).[5] The extended model that allows for imperfect substitution between domestic and imported varieties indicates instead that in the sample of Latin American countries, government weight on sector lobbying averages 32 percent of the weight on social welfare, peaking at 89 percent in Central America.

Extending the protection for sale model suggests that greater substitutability between domestic and imported varieties leads to higher trade barriers, all other things being equal. 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, the second half of the empirical analysis introduces a two-tier utility function that allows for different elasticities of substitution between the domestic variety and varieties imported from different regions of the world. Interestingly, Chinese imports are closer substitutes for domestic goods than are imports from the rest of the world, whereas the evidence is less clear-cut for Indian imports.

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

5. Gawande and Bandhopadhyay (2000), Goldberg and Maggi (1999), and Facchini, Van Biesebroeck, and Willmann (2006) estimated this model for the United States, and Mitra, Thomakos, and Ulubasoglu (2002) estimated it for Turkey. McCalman (2004) found 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).

evidence that China produces goods that are closer substitutes to Latin American goods than those produced by India.

The imperfect substitutability of domestic and imported varieties was first introduced in the protection for sale model by Chang (2005), who developed a framework featuring Dixit-Stiglitz-like differentiated goods sectors and analyzed the effects of this market structure on the trade policy outcome of the lobbying game. As Chang points out, the framework is ideally suited to study the intra-industry trade flows that dominate trade between developed countries. The theoretical model here instead stops short of such a change in the market structure because trade between developing countries and trade between developed and developing countries are of interest. More important, different elasticities of substitution should be allowed through different source countries, a generalization that cannot be easily introduced in a Dixit-Stiglitz framework. For these reasons, a simpler, perfectly competitive setup is used that forgoes the rent-shifting effects of Chang's model but allows the effect of the elasticity of substitution on trade policy to be unambiguously established.

The article is organized as follows. Section I provides some prima facie evidence of Latin American tariffs on goods heavily imported from China and India. Section II develops the extension to Grossman and Helpman's protection for sale model. Section III presents the empirical methodology and results. Section IV offers some implications of the findings.

I. Do Chinese and Indian Imports Face Higher Average Trade Barriers in Latin America?

Answering this question requires 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 possible with internationally comparable data—the six-digit level of the Harmonized System (HS)—using data for 1992–2004. Country coverage and data sources are discussed in the supplemental appendix available at http://wber.oxfordjournals.org/.

During 1992–2004, Latin America's average import-weighted most favored nation tariff applied to world imports was 13 percent. The import-weighted most favored nation tariff applied to Chinese and Indian imports was 9 percent higher.[6] The largest protectionist bias against China and India was found

6. Applied most favored nation tariffs as reported by each country are used throughout the article. One problem with this measure is that it does not capture exemptions due to investment or export regimes (such as duty drawback or rebate systems that reduce or eliminate duties paid on intermediate inputs). The alternative—using effectively collected tariffs—is not feasible because customs data on tariff collection by product for the Latin American countries in the sample are unavailable. Thus, the assumption 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).

among Central America (66 percent higher) and Andean countries (26 percent higher). But tariffs are only part of the story. Antidumping 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 (that is, including ad valorem equivalents of nontariff barriers) on overall world imports was 27 percent (Kee, Nicita, and Olarreaga 2009), but it was 10 percent higher on Chinese and Indian imports. The largest protectionist bias against China and India, including ad valorem equivalents of nontariff barriers, was in the Southern Cone countries (20 percent higher).

The differences in average applied most favored nation tariffs are due to an import-bundle composition effect and not to higher tariffs applied specifically to Chinese and Indian imports. Indeed, all Latin American countries were WTO members during the period of analysis and were thus required to apply most favored nation tariff rates to India, another WTO member. The same is true for China, even though the country formally entered the WTO only in 2001, because Latin American countries' tariff policies at the time did not distinguish between WTO members and nonmembers.[7]

But care should be used before interpreting these averages as evidence that Chinese and Indian imports led to higher most favored nation tariffs in Latin America. Two important issues need to be addressed. First, the causal relationship could go in the opposite direction. In other words, higher tariffs may affect less competitive trading partners more, which 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 more imports from China than from other countries because Chinese firms are more productive.[8] Second, the correlations might be affected by endogeneity due to omitted variable bias, because the goods in which China and India have a comparative advantage might be those in which Latin American countries have the most protection due to 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, the sectors with the strongest political clout in Latin America.[9]

These two issues are addressed by instrumenting the share of imports from China and India using their countries' share in world trade by good—a

7. As discussed, the most favored nation 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 and others (2009) find that bound tariffs were on average three times larger than applied tariffs.

8. The authors are grateful to Chad Bown for providing this example.

9. Bown (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.

measure of their success in the global marketplace—as well as the U.S. capital-labor ratio in each sector—a measure of their source of comparative advantage. Country, year, and sector fixed effects are also introduced to further address omitted variable bias.

The logic is as follows. First, measures of the success and of the drivers of Chinese and Indian exports in the global market should be correlated with the share of imports from the two countries to Latin America by sector. Second, because individual Latin American countries are small (none accounts for more than 2 percent of world trade), their trade policy should not affect China's or India's competitive position in the world market or its drivers (exclusion restriction). Still, 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 more protection. To deal with this issue, sector fixed effects are introduced in the tariff equation. First, two-digit HS fixed effects that vary by country and year are considered to capture sector-specific protectionist forces that are common across sectors but that vary by country and year.[10] And the sources of both comparative advantage and protectionism might be operating at a more disaggregated level, so a full set of six-digit fixed effects is also introduced.

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

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

where $t_{k,c,t}$ is the level of protection on good k (at the six-digit HS level) in country c at time t, β_k is six-digit HS fixed effects, $I_{k \in 2 \text{ digit},c,t}$ is a full set of product fixed effects (at the two-digit HS level) that vary by country and year, $m_{k,c,t}$ is imports and $s_{k,c,t}$ is the share of imports from China and India in sector k of country c at time t, and $\mu_{k,c,t}$ is a mean-zero error term when it is assumed that there are no six-digit time and country invariant determinants of tariffs.[11] Two specifications are used. The first includes the overall share of imports from China and India; the second introduces the share of imports from China separately.

Tables 1 and 2 report the instrumental variable results for a pool of 10 Latin American countries and four subregions: Andean countries (Bolivia, Colombia, Peru, and Venezuela), Central America (Costa Rica and Guatemala), Mexico, and the Southern Cone countries (Argentina, Brazil, and Uruguay). Table 1 reports results using tariffs as the left-side variable, and table 2 reports results using ad valorem equivalents of nontariff barriers.

10. This 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.

11. The instrumental variables used in this article vary by product (six-digit HS level) and year but not by country. For this reason equation (1) is estimated using robust standard errors with clustering at the product-year level. Moreover, the regressions are estimated for subregions, and the results at the regional level are compared with the results at the country level. Differences between the estimates at the regional level and at the country level were not significant.

T ABLE 1. Tariffs in Latin America and Imports from China and India

	Latin America		Andean countries		Central America		Mexico		Southern Cone	
Total imports ($ hundred millions)	0.67**	0.61**	−0.69	−0.62	0.27	0.31	0.12	0.03	0.99**	1.00**
($m_{k,c,t}$)	(0.20)	(0.20)	(0.46)	(0.49)	(0.71)	(0.71)	(0.14)	(0.11)	(0.38)	(0.38)
Share of imports from China and India (%)	0.26**		0.05*		−0.12**		0.35**		0.21*	
($s_{k,c}$ from C&b t)	(0.06)		(0.03)		(0.05)		(0.14)		(0.11)	
Share of imports from China (%)		0.21**		0.06*		−0.13**		0.28**		0.19**
($s_{k,c}$ from, t)		(0.05)		(0.03)		(0.06)		(0.11)		(0.09)
R-squared	0.54	0.54	0.67	0.67	0.63	0.63	0.54	0.53	0.64	0.63
F-statistic, first stage	9.49	12.4	9.79	10.90	5.97	6.43	2.06	2.27	12.5	17.17
F-statistic, orthogonal	1.00	1.80	0.00	0.03	0.01	0.01	0.02	0.07	0.07	0.18
Number of observations	247,822	247,822	101,010	101,010	34,215	34,215	22,676	22,676	89,921	89,921
Number of countries	10	10	4	4	2	2	1	1	3	3

*Statistically significant at the 10 percent level; ** statistically significant at the 5 percent level.

Note: Regressions are estimated using an instrumental variable approach, where all variables are instrumented (ordinary least squares results are reported in table S3 of the supplemental appendix). Instruments are the shares of goods from China and India in world markets and the U.S. 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*-statistic, orthogonal."All regressions include dummy variables that vary by two-digit Harmonized System (HS) sector, country, and year, as well as six-digit HS fixed effects. Numbers in parentheses are White robust standard errors clustered at the six-digit HS product-year level.

Source: Authors' analysis based on data described in the supplemental appendix, available at http://wber.oxfordjournals.org.

TABLE 2. Overall Protection in Latin America and Imports from China and India

	Latin America		Andean countries		Central America		Mexico		Southern Cone	
Total imports ($ hundred millions) ($m_{k,c,t}$)	1.03**	0.84**	1.81	1.13	na	na	0.67**	0.36**	0.57	0.23
	(0.27)	(0.29)	(1.13)	(0.82)			(0.20)	(0.18)	(0.40)	(0.20)
Share of imports from China and India (%) ($s_{k,c}$ from C&I,t)	0.46**		0.16*		na		2.87**		0.21*	
	(0.10)		(0.09)				(0.65)		(0.11)	
Share of imports from China (%) ($s_{k,c}$ from C,t)		0.41**		0.12*		na		2.06**		0.14**
		(0.10)		(0.07)				(0.59)		(0.06)
R-squared	0.16	0.17	0.25	0.25			0.18	0.18	0.14	0.15
F-statistic, first stage	27.09	26.67	65.29	65.29			10.27	11.41	47.12	54.11
F-statistic, orthogonal	0.68	3.14	0.20	0.13			12.11	8.58	0.08	0.00
Number of observations	22,542	22,542	7,514	7,514	0	0	3,757	3,757	11,271	11,271
Number of countries	8	8	4	4	0	0	1	1	3	3

* Statistically significant at the 10 percent level; ** statistically significant at the 5 percent level. na indicates that the data is not available.

Note: Regressions are estimated using an instrumental variable approach, where all variables are instrumented (ordinary least squares results are reported in table S3 of the supplemental appendix). Instruments are the shares of goods from China and India in world markets and the U.S. 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*-statistic, orthogonal." All regressions include dummy variables that vary by two-digit Harmonized System (HS) sector and country. Numbers in parentheses are White robust standard errors clustered at the six-digit HS level.

Source: Authors' analysis based on data described in the supplemental appendix, available at http://wber.oxfordjournals.org.

Because the ad valorem equivalents are available only for 2001, table 2 does not report time variation in the results.[12]

For the pooled Latin American group, the results suggest that an increase of 10 percentage points in the share of imports from China and India in a given sector translates into a 2.6 percentage points increase in the average ad valorem applied most favored nation tariff. Recalling that by end of the period the average import share from China and India was 10 percent and that the average most favored nation tariff on imports from all countries was 13 percent (see table S1 in the supplemental appendix), the estimation suggests that doubling the importance of China and India in Latin America's trade would lead to an increase of approximately 20 percentage points in the average most favored nation tariff.[13]

The results hold for subgroups of Latin American countries, except Central America. This finding is somewhat surprising, because Central American countries had a large bias against Chinese and Indian imports (see table S1 in the supplemental appendix), and highlights the importance of introducing different layers of fixed effects in the estimation. In fact, the strong average bias against Chinese and Indian imports observed among Central American countries may be due simply to the fact that politically strong sectors in the region are those in which China has a strong comparative advantage, which does not imply that Central America has responded to the rapid growth in imports from China and India with more protection.

As it turns out, the results are driven mainly by China. A separate regression with imports from only China yields parameter estimates that are not statistically different from the ones reported when the sum of imports from both China and India is included. This highlights the relative importance of these two countries as sources of imports for Latin America and may explain why the protectionist bias toward goods from China is greather than that toward goods from India.[14]

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12. Most first-stage regressions are highly statistically significant with *p*-values less than 0.01. The results of first-stage regressions are available from the authors upon request. When overall measures of protection are used in table 2, data are available only for 2001, and using sector fixed effects at the six-digit level would prevent the use of instrumental variables because they are organized at the same level of aggregation.

13. To account for the potentially high degree of inertia in the tariff structure, the specifications in table 1 were run using a dynamic panel estimator (Arellano-Bond). The results are qualitatively not affected and are available from the authors upon request.

14. Equation (1) was also estimated controlling for the presence of preferential trade agreements and for the accession of China to the WTO. The first exercise introduces dummy variables to control for the presence of preferential agreements. The second one uses instead the interaction between the share of imports from China and a dummy variable equal to 1 after 2000 as an explanatory variable. The results from above remain intact and are available from the authors upon request. A specification that explains pre-1995 tariffs in Latin America was also run using the share of imports from China and India in 2004. It showed no correlation between the tariff observed before 1995 and imports from China in 2004, suggesting that the findings do not suffer from reverse causality. The results are reported in table S2 in the supplemental appendix.

Does the pattern of more protection applied to Chinese and Indian goods hold for nontariff barriers as well? The answer is yes, according to the results in table 2, which is based on the same specification as table 1 but includes the ad valorem equivalents of nontariff barriers obtained by Kee, Nicita, and Olarreaga (2009) at the six-digit HS level.[15] The statistical significance remains as high as in table 1 and with the same pattern. Results are not available for Central America because there are no estimates available for the trade restrictiveness of the subregion's nontariff barriers.[16]

In sum, sectors with a larger share of Chinese and Indian imports tend to receive higher protection, driven mainly by Chinese imports, which are several times larger than Indian imports.[17] Given these results, a more structural explanation of the differences in protection observed in goods heavily imported from China and India is in order. To that end, the Grossman-Helpman protection for sale model is extended to allow for imperfect substitution between domestic and imported varieties of goods. The extended model can then be taken to the data.

II. 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, the model features a small open economy that sets trade policy on imports from the rest of the world (ROW). The key hypothesis is that goods are differentiated by location of origin—this uses the Armington assumption regarding imported and domestic varieties of a good as imperfect substitutes. The model has $n + 1$ types of goods, and each type can be either produced domestically or imported.[18]

Consumers in the home country maximize the following quasi-linear utility function:

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

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15. These estimates exist only for 2001, so the time dimension in the sample is lost. The ad valorem equivalents of nontariff barriers constructed by Kee, Nicita, and Olarreaga (2009) and used in table 2 do not vary by source country. This is appropriate for most nonttariff barriers, which are applied on a nondiscriminatory basis (such as sanitary and phytosanitary measures), but is more problematic for other nontariff barriers that are inherently discriminatory (such as antidumping duties). Bown (2010) has shown that for Latin America heavy users of antidumping measures (Argentina, Brazil, and Mexico) tend to disproportionately target China.

16. The first-stage regressions used to compute the results in table 2 are generally highly statistically significant and are available from the authors upon request.

17. The share of imports from China ranges from 76 percent to 91 percent of total imports from China and India combined for 1992–2004, with an average of 85 percent.

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

where k is the type of good and U_k (·) is a strictly concave subutility function ($U_k = E_k \ln X_k$ —that is, an upper tier Cobb-Douglas) that depends on a constant elasticity of substitution aggregate of the imported (i) and domestic (d) variety of the goods:

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

where $x_{k,d}$ is consumption of the domestic variety of good $k \in \{ 1, \ldots, n\}$, $x_{k,i}$ is consumption of the imported variety, $\sigma_k = \frac{1}{1-\rho_k} > 1$ is the elasticity of substitution between the two varieties, and good 0 is the numéraire. 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, linear supply schedules are used—that is, profit functions are assumed to be quadratic. Production of good 0 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 1.

For an individual with income E, maximizing 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 good:

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

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

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). In line with a substantial part of the literature and in view of this article's goal, export policies are not explicitly considered.

The price of the domestic variety results from the interplay of domestic supply and demand, which varies with the prices of both the domestic and the imported varieties. This relationship depends on the degree of substitutability. In particular, setting demand equal to supply in the market for the domestic variety—or $x_{k,d}(p_{k,d}, \ p_{k,i}, \ E_k; \ \rho_k) = \pi'(p_{k,d})$ —implicitly defines the

equilibrium price of the domestic variety

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

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 prices of both the domestic and imported varieties and on how it is influenced by the elasticity of substitution, the supply of the domestic variety is assumed to take the following linear form:

$$(7) \qquad y_{k,d} = p_{k,d}.$$

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

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

Since $p_{k,d}$ cannot be explicitly solved for, the next step is to totally differentiate the equilibrium condition. Using the demand function given by equation (4) above yields

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

To analyze the relationship between the price of the domestic and foreign varieties, $\frac{dp_{k,d}}{dp_{k,i}}$, ρ is held constant in equation (9), implying

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

where

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

and

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

It is easy to show that both the numerator and the denominator are positive, since $0 < \rho_k < 1$. Thus, it has been established that $\frac{dp_{k,d}}{dp_{k,i}} \geq 0$ or 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 their prices? First, consider two extreme cases at each end of the spectrum. If the elasticity of substitution between the domestic and the imported varieties equals 1 ($\rho_k = 0$), we are in the case of a Cobb-Douglas aggregator. In this case, the price of the domestic variety is unaffected: $dp_{k,d}/dp_{k,i} = 0$. 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 is the assumption made in the Grossman-Helpman protection for sale model.[19]

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

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

Solving explicitly, this can be shown to be positive (see the supplemental appendix), as long as demand and supply of the domestic variety do not diverge too much.[20] It can therefore be concluded that $p'_k \equiv dp_{k,d}/dp_{k,i}$ is a positive function of ρ_k.

Lobbying Game

The lobbying game is modeled along the lines of Grossman and Helpman's (1994) 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 consumer has surplus $CS(\mathbf{t}) = \sum_k [U_k(X_k(p_{k,d}, p_{k,i})) - p_{k,d}(p_k^* + t_k)x_{k,d} - (p_k^* + t_k)x_{k,i}]$ and receives 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. Both components depend on the price of the domestic variety, and expression (6) is used to convey them in terms of tariffs.

Assuming that ownership of the specific factor is highly concentrated in the population and in particular that the factor owners account for a negligible

19. If both varieties were complementary ($\rho_k < 0$), there would be a negative correlation between both prices. Such a case is not considered here because two varieties of the same good are being modeled.

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

share of the total population, the objective function of each organized group can be approximated by

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

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

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

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

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

where L denotes the labor force.

Two additional assumptions are used to solve for the optimal tariff. First, as in most of the literature, the focus is on contribution schedules, which are differentiable (truthful) around the equilibrium point. Second, all sectors are assumed to be politically organized.[21] As Mitra, Thomakos, and Ulubasoglu (2006) argue, at the four-digit International Standard Industrial Classification (ISIC) level, all sectors of the U.S. economy can be considered politically organized. Thus, although information on political organization by sector in Latin America is unavailable (because there is no legal requirement for public disclosure of sectors' political contributions), this assumption is likely to be harmless because the structural analysis is carried out at the more aggregate three-digit ISIC level.[22]

Rearranging the first-order condition of the government's maximization problem in equation (15) yields

$$(17) \qquad\qquad \frac{t_k^0}{1 + t_k^0} = \frac{1}{a} \times \frac{z_k}{\varepsilon_k} \times p_k'$$

where $t_k^0 = t_k/p_{k,i}$ is the ad valorem tariff, $z_k = x_{k,d}/x_{k,i}$ 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

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

22. As discussed in section III, a robustness check was carried out in which unorganized sectors are identified through a simple "rule of thumb." The results are not affected.

impact on the domestic price.[23] The last term is the main innovation with respect to the standard model (p'_k is given by equation [10]), which is shown above to depend positively on the elasticity of substitution. Thus in the presence of high substitutability between domestic and imported varieties of a good, tariffs are likely to be higher.

III. Empirical Analysis

Assessing the ability of the model to explain the patterns of protection toward Chinese and Indian imports highlighted in the preliminary data analysis involves two steps. First, the performance of the model is compared with that of the standard Grossman-Helpman benchmark. If product heterogeneity is important, the model should fit the data better than the standard benchmark does (see section IV).

Second, if the model performs better, explaining more protection toward Chinese and Indian imports requires investigating whether the elasticity of substitution between domestic goods and Chinese and Indian goods is higher than the elasticity of substitution between domestic goods and goods imported from the rest of the world (see section IV). Due to constraints in data availability, all the structural empirical investigations are carried out at the three-digit ISIC level.[24]

Does the Extended Model Perform Better?

To analyze the importance of allowing for imperfect substitutability between domestic and imported goods, the performance of the traditional protection for sale model with homogeneous goods is compared with the extended model on the sample of Latin American countries. The extended model is provided by equation (17), and the traditional model by equation (17) but without the last term. The R-squared of the two regressions is compared, and the Davidson and Mackinnon (1981) nonnested *J*-test is applied to determine which model better explains Latin America's tariff structure.[25]

The two specifications are also assessed for 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 Kirshna 2004 for a survey of the empirical literature). According to existing

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23. Formally, $\varepsilon_k = \varepsilon_{k,i} + \varepsilon_{k,d}\varepsilon_{p,k} = \frac{\partial x_{k,i}}{\partial p_{k,i}} \frac{p_{k,i}}{x_{k,i}} + \frac{\partial x_{k,i}}{\partial p_{k,d}} \frac{p_{k,d}}{x_{k,i}} \times \frac{\partial p_{k,d}}{\partial p_{k,i}} \frac{p_{k,i}}{p_{k,d}}$.

24. In particular, to estimate equation (22), the share of domestic goods in total consumption is needed, but data on production are available only at the three-digit ISIC level.

25. This 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, that the alternative is the right specification cannot be rejected. The problem with this test is that either of the alternatives may not be rejected or both may be rejected, meaning that the test is inconclusive.

estimates of the traditional protection for sale model with homogeneous goods, the weight attached by the government to sector lobbying when setting trade policy is just 1–2 percent of the weight the government puts on social welfare. This is hardly consistent with observed behavior and tariff structures. A lower estimate for a when bringing the extended model to the data would suggest that it provides more realistic estimates for this key parameter.

The extended model requires an estimate for p'_k, the derivative of domestic prices with respect to the price of the composite imported good, which is defined in equation (10). Substituting in equation (10) the derivative of domestic demand with respect to the prices of the domestic (equation [11]) and the foreign varieties (equation [12]) yields

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

To estimate p'_k, data on the prices of the domestic and the composite imported good are needed, as is consumer expenditure in sector k. The relative price of the domestic good and the composite imported good is obtained using the two first-order conditions of the consumer maximization problem—more precisely, the ratio between equations (4) and (5). The quasi-linear structure of the theoretical framework implies that there are neither income nor substitution effects for nonnuméraire goods. Thus, the model allows the assumption that the price of the imported good in every sector equals 1. Figures for consumption are readily obtained from trade and production data.[26] Equations (4) and (5) can then be used to calculate the price of the domestic varieties.

To compute p'_k, an estimate of ρ_k is also needed. This can be easily obtained by noting that the absolute value of the price elasticity of import demand equals $1/(1 - \rho_k)$.[27] Using recent estimates of price elasticities of import demand (Kee, Nicita, and Olarreaga 2008), the equality above can be solved to obtain $\rho_k = (\varepsilon_{k,i} - 1)/\varepsilon_{k,i}$.

One concern 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-side variables. To correct for these issues,

26. These data are actually apparent consumption, which equals imports plus domestic production minus exports.

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

China and India's share in world trade by product and the U.S. capital-labor ratio in each sector are used as instruments, and p'_k is instrumented jointly with the rest of the right-side term.[28]

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 subregions. Results for the pooled sample, the Southern Cone countries and Mexico always have the expected positive sign on the coefficient of the Grossman-Helpman term and the extended Grossman-Helpman term. For Andean countries and Central America the coefficient is negative when using the traditional Grossman-Helpman specification, which is at odds with theory, but positive for the extended Grossman-Helpman specification, which is consistent with theory. In fact, in all cases the extended Grossman-Helpman coefficient has the expected sign and is statistically different from 0.

The results also suggest that the extended Grossman-Helpman model performs better in terms of R-squared, indicating that it fits the data better. The Davidson-McKinnon nonnested *J* test for model specification indicates that the extended Grossman-Helpman model dominates the model with homogenous goods using either the pooled sample or the data by subregion.

As seen in equation (17), the coefficient of the Grossman-Helpman term (in both its traditional and extended forms) is represented by $1/a$—that is, the inverse of the weight the government attaches to social welfare relative to sector lobbying when setting trade policy. The extended Grossman-Helpman model yields estimates for this parameter that are all positive (as expected) and statistically different from 0.

More interestingly, the estimates of the weight governments put on welfare relative to sector lobbying (*a*) are more realistic than the figures obtained in the literature. For the extended Grossman-Helpman model, the estimated parameter *a* ranges from 1.12 to 4.73, rather than from 900 to 1,600 (or even negative) as in the traditional Grossman-Helpman model. Allowing for imperfect substitution between domestic and imported goods thus provides one solution to the puzzle of large estimates of *a*.[29] The results from the traditional Grossman-Helpman model suggest that in our sample the relative weight the government puts on sector lobbying is about 0.1 percent of the weight put on social welfare for the pooled Latin America sample $(0.001 = 1/[a = 918])$. With such a weight, assuming no other market imperfections, it would be very difficult to explain the high trade protection observed in Latin America. On the other hand, the estimates from the extended Grossman-Helpman model

28. For a discussion of the use of these instruments, see section II.

29. Standard 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. This strategy is used because the instrumental variables do not vary by country. The tariff equations that emerge from the traditional and extended protection for sale models for subregions in Latin America are also estimated, and the results do not differ significantly from the results obtained at the country level.

TABLE 3. Estimating the Classic and Extended Grossman-Helpman Model

	Latin America	Andean countries	Central America	Mexico	Southern Cone
Grossman-Helpman term	0.001**	−0.003**	−0.002	0.002	0.0006**
$(z_{k,c,t}/\varepsilon_{k,c,t})$	(0.00)	(0.001)	(0.002)	(0.003)	(0.0001)
Implied a	918**	−293**	−363	337	1,639**
(weight on welfare)	(169.33)	(58.59)	(210.28)	(405.70)	(334.07)
Extended Grossman-Helpman term	0.32**	0.21**	0.89**	0.28**	0.32**
$(z_{k,c,t} p'/\varepsilon_{k,c,t})$	(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	0.29 0.40	0.09 0.34	0.10 0.17	0.20 0.28
Number of observations	1,374 1,374	532 532	155 155	100 100	587 587
Number of countries	10 10	4 4	2 2	1 1	3 3
J-nonnested test[a]	Ext-GH	Ext-GH	Ext-GH	Ext-GH	Ext-GH

* Statistically significant at the 10 percent level; ** statistically significant at the 5 percent level.

Note: Regressions are estimated using an instrumental variable approach, where all variables are instrumented. Instruments are the shares of goods from China and India in world markets and the U.S. capital-labor ratio. All regressions include country and year dummy variables. Numbers in parentheses are White robust standard errors clustered at the three-digit International Standard Industrial Classification level.

a. Davidson-McKinnon (1981) *J*-nonnested test for model specification. "Ext-GH" indicates that the extended Grossman-Helpman model dominated the tradition Grossman-Helpman model at the 5 percent level.

Source: Authors' analysis based on data described in the supplemental appendix, available at http://wber.oxfordjournals.org.

suggest that the weight the government puts on sector lobbying is 32 percent of the weight it puts on social welfare $(0.32 = 1/(a = 3.16))$, suggesting a much larger scope for lobbying influence. Central American governments have the lowest concern for social welfare $(a = 1.12)$, whereas Andean governments have the highest concern $(a = 4.73)$.

Several robustness checks were undertaken. First, equation (17) and the traditional protection for sale model were estimated using the overall level of protection that includes ad valorem equivalents of nontariff barriers as the left-side variable. Results for the pooled Latin American sample suggest that parameter *a* equals 898 for the traditional Grossman-Helpman model and 2.17 for the extended Grossman-Helpman model.[30] Second, a specification that allows for the existence of politically unorganized sectors was used to check the robustness of the baseline results, which are instead based on the assumption that all sectors are politically organized. Sectors were classified as not organized when protection fell below that predicted by the theoretical model assuming all sectors to be organized (that is, sectors that are not politically organized are those where the error term in the estimates reported in table 3 was negative).[31] Results are remarkably robust. The estimates for the average *a* in Latin America when using the basic Grossman-Helpman specification is 333, whereas the estimate using the extended model is about 2. At the regional level, the estimate is reduced from 200 to 3 for the Andean countries, from 77 to 1 for Central America, from 50 to 1 for Mexico, and from 714 to 5 for the Southern Cone countries.

Table 4 shows the first stage regressions results used to obtain the results in table 3. The *F*-statistics indicate that the instrumental variables used to estimate the traditional Grossman-Helpman model are statistically significant for the pooled Latin America sample, the Andean countries, Mexico, and the Southern Cone countries but not for Central America. The *F*-test indicates that the instrumental variables used to estimate the extended Grossman-Helpman model are jointly significant for the pooled Latin America sample, the Andean countries, Central America, and the Southern Cone countries. In most cases the *F*-statistic is greater than 10 in the extended Grossman-Helpman model's first-stage regressions. Since there is a single endogenous regressor, these results reinforce the appropriateness of the instrumental variables used to estimate the extended Grossman-Helpman model.[32]

30. However, none of the estimates of the traditional Grossman-Helpman model is statistically different from 0 (although they are different from each other). For the extended Grossman-Helpman model, the estimates for the pooled sample as well as for the Southern Cone are statistically significant and have the expected sign. The results for the other subregions are not statistically significant.

31. This rule follows the main line of argument found in Cadot and others (2007).

32. This observation follows from the "rule of thumb" suggested by Stager and Stock (1997).

TABLE 4. First-Stage Results

	Latin America		Andean countries		Central America		Mexico		Southern Cone	
	GH	Ext-GH	GH	Ext-GH	GH	Ext-GH	GH	Ext-GH	GH	Ext-GH
U.S. capital to labor ratio ($ hundred millions)	3,278.49**	−8.33	1,231.11**	−4.22	991.34	−2.61	−141.94	−47.6	6,515.69**	−7.24
	(1,195.87)	(8.74)	(463.61)	(8.02)	(1,411.88)	(8.50)	(585.00)	(37.1)	(2,349.58)	(8.56)
China's world share	218.67**	1.02**	17.61	1.17**	−25.42	0.51*	−2.00	0.77	600.74**	1.08**
	(66.38)	(0.27)	(15.18)	(0.3)	(39.50)	(0.29)	(23.30)	(0.62)	(136.92)	(0.28)
India's world share	−739.48**	0.96	−104.97**	1.20	−211.40	0.80	−206.90**	−2.20	−1,570.25**	1.04
	(171.27)	(0.92)	(38.56)	(0.97)	(150.35)	(1.03)	(80.11)	(1.84)	(367.42)	(1.01)
Constant	4.52*	0.04**	7.00**	0.02	4.72**	0.02	7.94**	0.15**	3.02	0.06**
	(2.38)	(0.01)	(1.43)	(0.01)	(1.55)	(0.01)	(2.08)	(0.05)	(4.50)	(0.02)
F-test	6.69**	16.00**	3.24*	11.03**	2.25	3.08*	2.98*	2.14	6.76**	16.04**
Number of observations	1,374	1,374	532	532	155	155	100	100	587	587
Number of countries	10	10	4	4	2	2	1	1	3	3

* Statistically significant at the 10 percent level; ** statistically significant at the 5 percent level.

Note: The dependent variable is the explanatory variable in table 3 (either the traditional Grossman-Helpman term or the extended Grossman-Helpman term). Numbers in parentheses are White robust standard errors clustered at the six-digit Harmonized System level.

Source: Authors' analysis based on data described in the supplemental appendix, available at http://wber.oxfordjournals.org.

Estimating the Substitutability between Domestic Goods and Goods Imported from China and India

For the better performance of the extended model to explain the higher level of protection towards goods imported from China and India, there must be a higher degree of substitutability between domestic goods and goods imported from China and India than between domestic goods and goods imported from the rest of the world.

To measure the substitutability between domestic and imported varieties, Sato's (1967) strategy is followed: the imported variety is disaggregated using a two-tier constant elasticity of substitution utility structure. In other words, starting from the utility function in equation (2), the aggregate good k is represented by

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

where $x_{k,d}$ is the domestic variety, $x_{k,i}$ is the imported variety, and $\sigma_k = \frac{1}{1-\rho_k}$. Denote the imports of good k by country j are denoted $x_{k,i,j}$, where j is an element of the set {China, India, rest of the world}. Let

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

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

$$\frac{\left(\frac{1}{\theta_{k,j}} - \frac{1}{\theta_{k,i}}\right) + \left(\frac{1}{\theta_{k,d}} - \frac{1}{\theta_{k,i}}\right)}{\sigma_{k,j}} = \frac{\left(\frac{1}{\theta_{k,j}} - \frac{1}{\theta_{k,i}}\right)}{\sigma_{k,i}}$$

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

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 (that is, $\theta_{k,i} = \sum_j \theta_{k,j}$), and $\theta_{k,d}$ is the share of total expenditure on the domestic variety of good k. Equation (5) can be used to derive the price elasticity of the composite of imported goods, ε_k. Solving for σ_k yields $\sigma_k = -\varepsilon_k$. Thus, with an estimate of the price elasticity of the imported composite good k from the existing literature, an estimate for σ_k can be derived. With data on σ_k and on the share of expenditure on domestic and on imported varieties, the relationship

described in equation (21) can be used to obtain estimates for the degree of substitutability between domestically produced goods and imports from China, India, and the rest of the world.[33]

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

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

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 $1/\alpha_{1,j}$. Equation (22) is the basis for the estimate of the relative substitutability between domestic goods and goods imported from different regions.

Equation (22) is estimated using data on imports from China, India, and the rest of the world separately. Parameter estimates obtained for equation (22), as well as the implied σ's, are reported in table 5. The degree of substitutability between goods imported from China and goods produced in Latin America—as measured by σ_C—is higher than the degree of substitutability between goods imported from the rest of the world and goods produced in Latin America—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.[34] The differences are statistically different from 0 at the 10 percent level for the pooled sample and for Andean countries. The estimates of the elasticity of substitution between goods imported from India and domestic goods (σ_I) are larger than the estimated σ_{ROW} for all specifications except the Andean countries and the Southern Cone countries, but the differences are never statistically different from 0.

Given that these estimates represent averages across different products, it is interesting to disentangle whether the higher values for the elasticity of substitution between domestic goods and Chinese imports is due to a composition effect or a within-product effect. Indeed, the higher elasticity may be explained simply by the fact that China's export bundle is closer to Latin America's production bundle than to the export bundle from the rest of the world. This issue is addressed by introducing product fixed effects in the specifications to estimate equation (22) to control for any product-specific effect (some goods having larger elasticities of substitution than others), and the sample is

33. The empirical analysis captures only the average degree of substitutability, σ_C, σ_I, and σ_{ROW}, as there are too few observations to estimate sector-specific elasticities.

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

TABLE 5. Estimating the Degree of Substitutability with Domestic Goods

	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-squared	[0.14, 0.19]	[0.08, 0.21]	[0.01, 0.03]	[0.10, 0.15]	[0.16, 0.18]
Number of observations	[1,499, 1,899]	[419, 603]	[114, 162]	[116, 130]	[637, 753]
Number of countries	10	4	2	1	3

* Statistically significant at the 10 percent level; ** statistically significant at the 5 percent level.

Note: All regressions use country fixed effects. Numbers in parentheses are White robust standard errors both for the coefficients and the implied substitution parameters. The range of R-squares and number of observations of the regressions are based on data on imports from China, India, and the rest of the of the world.

Source: Authors' analysis based on data described in the supplemental appendix, available at http://wber.oxfordjournals.org.

restricted to observations where imports from China accounts for more than 2 percent of total imports. Results are provided in table S4 of the supplemental appendix and are qualitatively similar to those in table 5. This suggests that the higher substitutability between Chinese goods and domestic goods is not due only to a composition effect.

To further verify the importance of within-sector differences in the elasticities of substitution, equation (22) was re-run using ISIC two-digit industries. In industries where goods imported from China are relatively important—ISIC 32, textiles and apparel; ISIC 36, pottery and earthware; and ISIC 38, machinery (see figure 1)—the elasticity of substitution between domestic goods and goods imported from China is always estimated to be larger than the elasticity of substitution between domestic goods and goods imported from the rest of the world. The same does not apply to goods imported from India.[35]

Finally, turning back to the first-stage regressions, the results in table 4 suggest that the larger the share of goods from China in world markets, the larger the extended Grossman-Helpman term, since the coefficient of the share of imports from China is always positive and usually statistically significant. And the extended Grossman-Helpman term contains p'_k, which as shown in the supplemental appendix increases with the elasticity of substitution between domestic goods and imported goods. Thus, these first-stage regressions also provide indirect evidence that the larger the share of goods from China in world markets, the larger the substitutability between Latin America's import and domestic bundles, which in turn increases protectionist pressures.

IV. Conclusion

The growing presence of Chinese and Indian goods in world markets has caught the attention of many observers. This article explores the response of Latin American policymakers to the growth of imports from China and India in their markets. Sectors in which the share of Chinese imports is larger tend to have higher tariffs, controlling for reverse causality and country, year, and sector fixed effects. The evidence is mixed for Indian imports.

To explain these results, the Grossman and Helpman (1994) protection for sale model is extended to allow for imperfect substitution between domestic and imported goods. The model suggests that as the elasticity of substitution between domestic and imported goods increases, the incentives to lobby also increase, and the resulting equilibrium tariff is higher.

The analysis is carried out in two steps. First, the extended framework developed in the paper is brought to the data and is shown to outperform the traditional Grossman-Helpman framework in two respects: it fits the data better and—more interesting—explicitly modeling imperfect substitutability yields

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35. None of the differences at the two-digit ISIC level is statistically significant.

substantially more realistic structural parameter estimates than those resulting from a homogeneous good specification.

Second, given that the extended model performs better, the substitutability of domestic goods with goods imported from either China, India, or the rest of the world is measured. Goods imported from China are on average closer substitutes to domestic goods than are goods imported from the rest of the world, while the evidence on goods imported from India is mixed. This suggests that closer substitutability with imports drives higher tariffs in sectors where Chinese imports are more relevant.

The analysis has important implications for the empirical literature on the protection for sale model. By ignoring imperfect substitutability between domestic and imported goods, previous studies have obtained estimates for the weight attached by the government to social welfare that are substantially upward biased. The analysis here suggests for the entire sample, under imperfect substitutability, that the weight attached to sector lobbying is about 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 existed in estimates of the protection for sale model carried out for the United States, the true value of a would be about 10 for the model estimated by Gawande and Bandhopadhyay (2000), instead of 3,000, as they report, and 0.23 for the benchmark specification in Goldberg and Maggi (1999). In particular, the latter would imply that the U.S. government puts five times more weight on political contributions than on social welfare when determining trade policy. Of course, these are just back of the envelope calculations, and a more careful analysis on U.S. data is needed before reaching a definitive conclusion. However, these results do suggest that allowing for imperfect substitutability within the Grossman-Helpman 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 literature has both overstated the benefits and understated the costs of a tariff. If substitutability between domestic and imported varieties is less than perfect, the domestic producer is likely to observe a smaller price increase than hoped for. And under imperfect substitutability, domestic consumers will have a harder time substituting away from 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 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.[36]

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REFERENCES

Bown, C.P. 2010. "China's WTO Entry: Antidumping, Safeguards, and Dispute Settlement." In S.-J. Wei and R. Feenstra, eds., *China's Growing Role in World Trade*. Chicago, Ill.: University of Chicago Press.

Cadot, O., J. de Melo, and M. Olarreaga. 2003. "The Protectionist Bias of Duty Drawbacks: Evidence from Mercosur." *Journal of International Economics* 59(1): 161–82.

Cadot, O., L. Dutoit, J.-M. Grether, and M. Olarreaga. 2007. "Endogenous Tariffs in a Common-Agency Model: A New Empirical Approach Applied to India." University of Geneva.

Calderón, C. 2008. "Trade, Specialization and Cycle Synchronization: Explaining Output Comovement between Latin America." In D. Lederman, M. Olarreaga, and G. Perry, eds., *China's and India's Challenge to Latin America: Opportunity or Threat?*. Washington, D.C.: World Bank.

CAME (Confederation of Medium Enterprises). 2004. "Las importaciones desde China crecieron 78.2% en el primer bimestre." Communiqué, November 16. [www.redcame.org.ar/index.php/came/comunicado/383].

Chang, P.L. 2005. "Protection for Sale under Monopolistic Competition." *Journal of International Economics* 66(2): 509–26.

Davidson, R. and J.G. Mackinnon. 1981. "Several Tests for Model Specification in the Presence of Alternative Hypotheses." *Econometrica* 49(3): 781–93.

The Economist. 2009. "Barak Obama Does a Bad Thing." September 14.

Facchini, G., J. Van Biesebroeck, and G. Willmann. 2006. "Protection for Sale with Imperfect Rent Capturing." *Canadian Journal of Economics* 39(3): 845–73.

Foletti, L., M. Fugazza, A. Nicita, and M. Olarreaga. 2009. "Smoke in the (Tariff) Water." University of Geneva, Department of Economics.

Gawande, K. and U. Bandhopadhyay. 2000. "Is Protection for Sale? A Test of the Grossman-Helpman Theory of Endogenous Protection." *Review of Economics and Statistics* 89(1): 139–52.

Gawande, K. and P. Krishna. 2004. "The Political Economy of Trade Policy: Empirical Approaches." In K. Choi and J. Harrigan, eds., *Handbook of International Trade*. New York: Basil Blackwell.

Goldberg, P. and G. Maggi. 1999. "Protection for Sale: An Empirical Investigation." *American Economic Review* 89(5): 1135–55.

Grossman, G. and E. Helpman. 1994. "Protection for Sale." *American Economic Review* 84(4): 833–50.

———. 1995. "The Politics of Free Trade Agreements." *American Economic Review* 85(4): 667–90.

Hoekman, B. and M. Kostecki. 2001. *The Political Economy of the World Trading System*. New York: Oxford University Press.

Iran Daily. 2005. "Brazil Imposing Limits on Chinese Goods." October 4. [/www.nitc.co.ir/iran-daily/1384/2393/html/ieconomy.htm#s96435].

Kee, H.L., A. Nicita, and M. Olarreaga. 2008. "Import Demand Elasticities and Trade Distortions." *Review of Economics and Statistics* 90(4): 666–82.

———. 2009. "Estimating Trade Restrictiveness Indices." *Economic Journal* 119(1): 172–99.

Luntz, Frank. 2005. *Republican Playbook*. New York: Hyperion.

McCalman, P. 2004. "Protection for Sale and Trade Liberalization: An Empirical Investigation." *Review of International Economics* 12(1): 81–94.

McKinley, J.C. 2005. "Mexico Builds Trade Ties with China." *New York Times*, September 19.

Mitra, D., D. Thomakos, and M. Ulubasoglu. 2002. "Protection for Sale in a Developing Country: Democracy versus Dictatorship." *Review of Economics and Statistics* 84(3): 497–508.

———. 2006. "Can We Obtain Realistic Estimates for the 'Protection for Sale' Model?" *Canadian Journal of Economics* 39(1): 187–210.

Sato, K. 1967. "A Two-Level Constant-Elasticity-of-Substitution Production Function." *Review of Economic Studies* 34(2): 201–18.

Staiger, D. and J. Stock. 1997. "Instrumental Variables Regression with Weak Instruments." *Econometrica* 65(3): 557–86.

Taipei Times. 2005. "Court Bars Bush from Imposing China Textile Limits." January 2.