Chapter IX

Armington elasticities for Brazil[1]

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In this article we estimate substitution elasticities for goods distinguished by place of production, specifying whether they are imported or produced domestically. These are known as the Armington (1969) elasticities and are widely used to assess the impact on the domestic economy of policy changes in countries' tariff structures; and, in particular, to evaluate the costs and benefits of signing free trade agreements. The sample period for our study is 1986-2002, and the estimation is done separately for each of the 28 industrial sectors specified in the Brazilian input-output table. Special consideration is given to the fact that the data is affected by import restrictions for part of that period, and that foreign trade liberalization

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occurred in Brazil in 1990. The estimation procedure is automated and takes into consideration the stochastic dynamic properties of the quantity and price series, using the appropriate estimation approach in each case. The Armington elasticities we estimate have the correct sign, and are significant at the 5% level for 20 sectors, at 10% for two sectors, and at 20% for two others. In one sector the estimated value is significant, but has the incorrect sign (negative). For three sectors the estimated elasticity is not significantly different from zero; but these represent only 12% of the average value of total import value in the period 1997-2002. The point estimate of the elasticity of substitution, for the sectors where it is positive and statistically different from zero, varies from 0.16 to 3.6; and its weighted average value is 0.93.

A. Introduction

Regional free trade agreements are hard to evaluate economically because they affect multiple productive sectors in several different ways, and they impact the performance of the national economy in a complex manner.[5] Nonetheless, it is possible to assess their main consequences using either a partial-equilibrium approach or, more broadly, computable general equilibrium models (CGEs). In either case, the agreement is represented by the tariff reductions it imposes on the countries signing it. These, in turn, alter the domestic price of the imported good relative to the domestically produced good, and this change affects the proportion of domestic demand that is supplied by imports. To analyse this effect and attempt to forecast its intensity, requires estimating the elasticity of substitution between goods distinguished by place of origin. These are known as Armington elasticities, in honour of the economist who first drew attention to their importance. Moreover, as the substitutability between imports and domestic production varies widely from one product to another, these estimates are needed at a disaggregated product level.

Estimates of Armington elasticities by sector are not readily available for most countries, despite their crucial importance for evaluating the impact of changes in trade policy on foreign trade flows. To deal with this lack of empirical data, studies in this area often make use of elasticity values obtained for other countries, in many cases completely disregarding important differences that may exist between

5 Brazil is engaged in the negotiation of several trade agreements, of varying scope: multilateral – the World Trade Organization (WTO); free trade agreements with the countries of MERCOSUR and the European Union (UE); and bilateral agreements, with South Africa and India, among others.

the production and consumption structures of different countries.[6] As these elasticity values are also not available for Brazil, the aim of this study is to estimate them for our country, using the longest data series available, which runs from 1986 to 2002. We have adopted the same level of aggregation as in the Brazilian input-output matrix, to make it easy to use the estimates in empirical studies of the country's import basket.

The approach proposed here makes a methodological innovation with respect to the literature in this area, by advocating extensive use of the time-series properties of the series in question, to select the most suitable estimation method and the corresponding equation. Depending on the order of integration of the relative-price and quantity series, we employ one of four approaches: simple regression of the levels of variables, regression of first differences, linear model of mixed equations, or a vector error correction model (VEC).[7] We also consider the possibility of a structural break occurring in the data series caused by the foreign-trade liberalization which began in Brazil in 1990, together with the possibility of seasonal factors and a time trend. This is done both in the tests to determine the order of integration of the series, and in the estimation itself. Lastly, we also consider the possibility that the demand for imports is affected not only by the level of the relative price of imported goods, but also by its uncertainty .

This painstaking approach to model specification and estimation has clear empirical advantages in the case of Brazil for the period considered, since an attempt to employ simpler methods had led to poor or incorrect estimates. We believe that the approach proposed here will also prove useful in the case of other countries, especially if the sample period includes a trade liberalization episode.

This paper is divided into five sections including the introduction. In section B we briefly review the concept of elasticity of substitution, and introduce the approach used to deal with the impact on the observed data of the foreign-trade liberalization initiative that began in 1990 and lasted through 1993. Section C discusses the tests used to determine the existence of unit roots in the price and quantity series, and the models used in the estimation process. Section D reports the estimates obtained for the 28 industrial sectors of the Brazilian input-output table; and section E summarizes the main conclusions.

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[6] For example, Sánchez (2001) evaluates the costs and benefits for Mercosur of joining FTAA using the applied general equilibrium model of the Global Trade Analysis Project (GETAP), but it arbitrarily multiplies the original elasticities by six. Harrison et al (2002) analyse the impacts of regional and multilateral trade agreements for Brazil while using elasticities estimated for Hong-Kong.

[7] We believe this is a pioneering application of the methods developed by Johansen (1988) to the problem of estimating the Armington elasticity.

B. Armington elasticity and Brazilian trade liberalization of 1990

The approach proposed by Armington (1969) to evaluate impacts of changes in trade policy on the volume of imports has been widely used, both in its original partial equilibrium formulation and in general equilibrium models.[8] The approach assumes that goods are differentiated by country of origin, and that domestic demand for each sector is supplied by a composite good which is a CES (Constant Elasticity of Substitution) aggregation of domestically produced and imported goods. This is represented in equation (1), for sector *i*.

$$(1) \qquad Q_i = \bar{Q}_i \left[\delta_i M_i^{-\rho_i} + (1-\delta_i) D_i^{-\rho_i} \right]^{-1/\rho_i}$$

where Q_i, M_i and D_i represent the quantity index of the aggregate good, the imported good, and the domestically produced good, respectively. The scale parameter is $\bar{Q}_i$, and ρ_i δ_i are the substitution and the distribution parameters, respectively. The first of these indicates the degree of substitutability between domestically produced and imported goods, and determines the shape of the indifference curve that represents the smooth transition between these two goods in the preferences of the representative consumer. Its role can be clearly seen by noting that, for ρ_i = -1, the composite good is a linear combination of M_i and D_i and the indifference curve is linear. The second parameter indicates the shares of imported and domestically produced goods in the composite good.

The solution to the problem of minimizing the cost of supplying total demand, given expenditure and the prices of the imported and domestically produced goods, gives the optimal mix of these two goods in the composite good Q_i, and is represented by equation (2). This shows that the proportions of the domestic and the imported good depend on the elasticity of substitution $\sigma_i = 1/(1+\rho_i)$ and on the ratio of their prices, represented by P_i^d and P_i^m.

$$(2) \qquad \frac{M_i}{D_i} = \left[\left(\frac{\delta_i}{1-\delta_i} \right) \left(\frac{P_i^d}{P_i^m} \right) \right]^{\sigma_i}$$

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8 A good example is Dervis, Melo and Robinson (1982) which uses the Armington specification in a computable general equilibrium (CGE) model, which has become a standard model for policy analysis. See also Melo and Robinson (1989) for a more detailed discussion of its use in CGEs. For examples of its use in a partial equilibrium framework, see the series of studies to measure the social cost of protection in several countries, started by Hufbauer and Elliot (1994) and sponsored by the Institute of International Economics, Washington, D.C.

The internal price of the imported good, which is the relevant value for the consumer's decision, is a function of P_i^e, its foreign-currency price on the foreign market, the exchange rate *(X)*, and the import tariff (τ_i) as shown in equation (3):

$$(3) \qquad P_i^m = P_i^e X\left(1+\tau_i\right)$$

It is useful to explore the behaviour of equation (2) for three extreme cases of the elasticity of substitution. If $\rho_i \rightarrow \infty$ and $\sigma_i \rightarrow 0$, there is no substitution between the two goods, and the ratio of their quantities does not depend on their relative prices. When $\rho_i \rightarrow -1$ and $\sigma_i \rightarrow \infty$, the two goods are perfect substitutes,[9] and small changes in the relative prices are sufficient to produce wide swings in the ratio M_i / D_i. Lastly, when $\rho_i = 0$ and $\sigma_i = 1$, the CES function in equation (1) reduces to the Cobb-Douglas function, and the ratio of expenditure on the imported and domestically produced goods is constant and equal to $\delta_i/(1-\delta_i)$.

Equation (2) also shows that an estimate of the elasticity of substitution, σ_i, can be obtained from the time series for the ratios M_i / D_i and P_i^d / P_i^m. Nonetheless, a relatively long time period is required for quantities to fully adjust to price changes. In the short run (a few months) the impact will probably be small, since several years are usually necessary for the quantity imported to fully reflect changes in the relative price of imports. The short-run and long-run elasticities are, therefore, different.

In this study we will estimate the long-run elasticities of substitution, using the argument employed by Gallaway, McDaniel and Rivera (2000) to justify adoption of that time horizon in their study. They point out that Armington elasticity estimates are most often used in comparative static analyses, in either partial- or general equilibrium models. In this type of analysis we compare the results of the controlled experiment to those obtained in the base year, assuming the economy has had long enough to adjust, so that the results of the experiment reflect the total effect of the policy experiment being evaluated.

To build a price series for imported goods that truly reflects the actual cost paid by importers to bring them into the country, a number of peculiarities of Brazilian tariff and import regulations in force between 1986 and 2002 need to be taken into account. The period can be divided into two distinct sub-periods that can be described in stylized fashion as follows.

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9 In this case we consider that the domestic price of good *i* is highly sensitive to the imported competitor, and that the ratio between them is approximately constant.

Before the unilateral trade liberalization that began in 1990, imports could be broadly classified in two categories:

(a) non-competing imports for which import tariffs were already low, precisely because there was no need to protect the domestic industry. Examples are metallurgy, coal, petroleum, some fertilizers, and capital goods with no close national substitute, etc.

(b) competing imports, for which import tariffs were extremely high. In this case, imports were only an economically viable alternative to domestic products if purchased by agents with access to reduced import tariffs or, in most cases, exemption. This special treatment was extended to agents eligible to claim special tax regimes, such as those applicable to State-owned enterprises, or those associated with special investment projects which, despite being private, were deemed to be of national interest. Examples are projects supported by the Amazonia Development Superintendency (SUDAM), the Northeast Development Superintendency (SUDENE), the Industrial Development Council (CDI), the Manaus Duty Free Zone, etc. This represents a situation of repressed demand, where potential importers were driven out of the import goods market by a combination of a prohibitive tariff and ineligibility for tariff-exempt status.[10] Consequently, there was a wedge between nominal tariffs, which were very high, and the tariffs that were actually paid, which were much lower because of these exemptions and reductions.

The dichotomy outlined above in the imported goods market was a consequence of the crisis in the Brazilian balance of payments which began in 1983 following a sudden stop in the flow of foreign savings to developing countries, and deepened in 1987 when the country declared a unilateral moratorium on its foreign-debt service. The situation only returned to normality in 2002, when the country signed a wide-ranging foreign-debt refinancing agreement with the international financial community.

After 1990, nominal tariffs were lowered, import restrictions were lifted, and most special tax regimes for imports were removed, except for those relating to the Manaus Free Trade Zone, the drawback regime, the special provisions for imports of computers and computer parts, and those relating to international agreements. As a consequence, the wedge between nominal and effectively paid tariffs, which had been very large, narrowed significantly.

10 According to Kume (1990), about 70% of imports, excluding oil, benefited from special tax regimes in this period. Competitive imports were also prohibitive because tariff redundancy was widespread.

The approach we propose captures these two situations, before and after liberalization, in a unified framework, by calculating the equivalent tariff faced by the average importer in each sector, taking into account the dichotomy between the two types of imports that existed before liberalization, as described above. Assuming that imports in each sector were determined by the condition represented in equation (2), but that the tariff faced by some of them was the nominal tariff, while for others it was zero because they were exempt, we can write equations (4) and (5), which determine the relative shares of imports and domestic production in each of these cases, respectively:

$$(4)\qquad \frac{\alpha_i M_i}{D_i} = \alpha_i \left[\left(\frac{\delta_i}{1-\delta_i}\right)\left(\frac{P_i^d}{P_i^e X}\right)\left(\frac{1}{1+\tau_i}\right)\right]^{\sigma_i}$$

$$(5)\qquad \frac{(1-\alpha_i)M_i}{D_i} = (1-\alpha_i)\left[\left(\frac{\delta_i}{1-\delta_i}\right)\left(\frac{P_i^d}{P_i^e X}\right)\right]^{\sigma_i}$$

where α_i denotes the proportion of imports that pay the full nominal tariff in sector *i*. Equation (6) sums these two types of imports, determined in equations (4) and (5) to calculate the ratio of imports to domestic production in sector *i*.

$$(6)\qquad \frac{M_i}{D_i} = \frac{\alpha_i M_i}{D_i} + \frac{(1-\alpha_i)M_i}{D_i} = \left[\left(\frac{\delta_i}{1-\delta_i}\right)\left(\frac{P_i^d}{P_i^e X}\right)\right]^{\sigma_i}\left[(1-\alpha_i)+\alpha_i\left(\frac{1}{1+\tau_i}\right)^{\sigma_i}\right]$$

For estimation purposes, a logarithmic transformation is applied to equation (6), to give equation (7).

$$(7)\qquad \log\left(\frac{M_i}{D_i}\right) = \sigma_i \log\left[\left(\frac{\delta_i}{1-\delta_i}\right)\left(\frac{P_i^d}{P_i^e X}\right)\right] + \log\left[(1-\alpha_i)+\alpha_i\left(\frac{1}{1+\tau_i}\right)^{\sigma_i}\right]$$

The fact that equation (7) is non-linear in the elasticity of substitution makes the estimation more difficult. To simplify, we take a Taylor series expansion of the second term in the right-hand side of (7) with respect to σ_i, in the neighbourhood of $\sigma_i = 1$, and obtain equation (8).

$$(8)\qquad \log\left[(1-\alpha_i)+\alpha_i\left(\frac{1}{1+\tau_i}\right)^{\sigma_i}\right] = \log\left[(1-\alpha_i)+\left(\frac{\alpha_i}{1+\tau_i}\right)\right] + \frac{\alpha_i}{1+\tau_i-\alpha_i\tau_i}(\sigma_i-1)\log\left(\frac{1}{1+\tau_i}\right) + \vartheta[\sigma_i-1]^2$$

Apart from the error term in the approximation, only the second term on the right-hand side of (8) depends on σ_i, and it does so linearly. That term is reproduced in equation (9), which also represents that fact explicitly.

$$(9) \qquad \sigma_i \left[\theta_i \log\left(\frac{1}{1+\tau_i} \right) \right], \quad \text{where} \quad \theta_i = \frac{\alpha_i}{1+\tau_i - \alpha_i \tau_i}$$

Using (9) as an approximation for the part of (7) that depends on σ_i, we obtain equation (10), where the constant term κ_i, defined in (11), consolidates terms in which the elasticity of substitution does not appear.

$$(10) \qquad \log\left(\frac{M_i}{D_i} \right) = \kappa_i + \sigma_i \log\left[\left(\frac{\delta_i}{1-\delta_i} \right) \left(\frac{P_i^d}{P_i^e X} \right) \right] + \sigma_i \theta_i \log\left(\frac{1}{1+\tau_i} \right)$$

$$(11) \qquad \kappa_i = \log\left[(1-\alpha_i) + \left(\frac{\alpha_i}{1+\tau_i} \right) \right] - \theta_i \log\left(\frac{1}{1+\tau_i} \right)$$

Lastly, equation (11) can be simplified to yield equation (12), which is similar to equation (2), but with the difference that the term capturing the effect of the tariff is raised to the power θ_i.

$$(12) \qquad \log\left(\frac{M_i}{D_i} \right) = \kappa_i + \sigma_i \log\left[\left(\frac{\delta_i}{1-\delta_i} \right) \left(\frac{P_i^d}{P_i^e X} \right) \left(\frac{1}{1+\tau_i} \right)^{\theta_i} \right]$$

The significance of the exponent θ_i in (12) can be understood by considering its definition in (9) and recalling that α_i is the proportion of imports that pay the full nominal tariff τ_i. It is then trivial to verify that equation (12) adequately reflects the two polar stylized types of imports that occurred before trade liberalization, because they correspond to the extreme values of α_i; and that it also correctly describes the intermediate situations existing afterwards.

We first discuss the situation before 1990. When the imported good was non-competing (case (a) described above), the tariff was low and applied to all import operations, i.e. $\alpha_i = 1$, so $\theta_i = 1$, and equation (12) reduces to (4). When the imported good is competing (case (b)), the pre-1990 tariff was very high, and the imports that actually entered the country did so under tariff exemption, which implies $\theta_i = 0$, and reduces equation (12) to (5).

Following trade liberalization, all goods are in an intermediate situation between the two polar cases that existed before 1990. They are also adequately represented by equation (12), whose behaviour can be extrapolated by examining a linear approximation for the expression that defines θ_i. This is given by equation (13), which is obtained by expanding equation (9) in a Taylor series with respect to α_i, $\tau_i \in [0,1]$ in the neighbourhood of the origin. This shows that for low values of the tariff $\tau_i \to 0$, when most imports are taxed, $\alpha_i \to 1$, and $\theta_i \to \alpha_i$. When only a small proportion of imports pay the nominal tariff $\alpha_i \to 0$, and $\theta_i \to (1 - \tau_i)$.

$$(13) \quad \theta_i = \alpha_i + \vartheta[\alpha_i]^2 + \tau_i\left(-\alpha_i + \vartheta[\alpha_i]^2\right) + \vartheta[\tau_i]^2$$

Estimation of equation (12) requires calculating θ_{it}, for which a measure of α_{it} is also needed.[11] This can be obtained by assuming that, in the context of the stylized situation described at the beginning of this section, the observed tariff is a weighted average of the nominal tariff applied to imports in case (b), and the zero tariff, which is relevant for imports in case (a), and where the weighting factor is the estimated value of α_{it}. This is represented in equation (14), where $\bar{\tau}_{it}$ stands for the *mean* observed tariff.

$$(14) \quad \bar{\tau}_{it} = \alpha_{it}\tau_{it}$$

From (14) we can then obtain an estimate for α_{it}, which is represented by $\bar{\alpha}_{it}$ in equation (15):

$$(15) \quad \tilde{\alpha}_{it} = \bar{\tau}_{it} / \tau_{it}$$

Lastly, we include a dummy variable in the estimated equations to represent the shift in the demand curve for imports, reflecting the possibility that became available to firms after the last quarter of 1990, of importing goods that were previously restricted provided the full import duty was paid.[12]

At this point the reader may be wondering why so much effort has made in these adjustments to allow the use of the data covering the period when foreign trade was restricted, and to take into account the impact of unilateral trade liberalization in Brazil, given that in our study this includes data for only five years (1986 to 1990). The main reason is not the additional degrees of freedom it gives the estimation, but rather to include a period in the empirical analysis where there were large

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11 In the notation for α_i we introduced the index t to emphasize that, besides being exogenous, it varies through time.

12 Import restrictions were officially lifted in March 1990 when the new Federal Government took office. Nonetheless, the non-tariff barriers that actually controlled most imports in practice were only eliminated in July of that year. Thus, only in the last quarter did economic agents effectively perceive and benefit from the new freedom to import.

changes in the relative internal price of imported goods. This is important because the aim of this study is precisely to estimate the curvature of the indifference curve for the CES function in equation (1), which can only be measured correctly if we have a database that includes a wide range of relative prices of domestically produced and imported goods.

At this point we speculate that the adjustments proposed here to make it possible to use a single unified framework for data relating to periods when trade restrictions were in place and for data relating to the liberalization and post liberalization periods, are also applicable to other countries besides Brazil. The reason is that the distortions we address, produced by very large nominal tariffs together with a large number of special import regimes that entail exception and special treatment for certain industries, firms or goods, are likely to arise in other countries that went through periods in which similar protectionist policies were used.

C. Empirical analysis

We used quarterly data for each sector of the Brazilian input-output matrix (IBGE – level 50), for the period 1986-2002. The database in electronic format is available to readers on demand, and its construction is described in Annex A.[13]

Estimating equation (12) for each sector requires examining the order of integration of the time series involved in it. A comparison of the stochastic characteristics of these series determined the model to be estimated to obtain the elasticity of substitution. This section describes the methodology used in these two steps of the estimation.

To implement the unit root test systematically with regard to inclusion of the constant and the time trend, we adopted the procedure proposed by Dolado, Jenkinson and Sosvilla-Rivero (1990). In cases where this indicated the existence of a unit root, we also applied the Perron (1989) test for a structural break in the fourth quarter of 1990, using the variant that specifies a break of the type represented by the *changing growth*

[13] The database in this study differs from that used in Tourinho, Kume and Pedroso (2002) because the foreign trade statistics for 1996 were revised in July 2002 (Funcex, 2002). The data for exports did not change significantly, but for several sectors expenditure on imports underwent major revision, mainly due to changes in the physical quantities imported. As the price and *quantum* indexes estimated by Funcex are chained in time, using 1996 as the base year, all the data were revised. Thus, the change in that year affected the level of the entire series, even though the rates of change were not altered, except for those calculated in relation to 1996.

model, according to the typology proposed there.[14] In all cases the level of significance adopted in the tests was 10%,[15] and the Akaike information criterion was used to determine the number of lags to be used. Annex B describes the methodology of the tests in greater detail.

To make the estimation methodology more explicit, we recall that equation (12) represents a long-term relation between p_{it} and q_{it}, defined in equations (16) to (19):

$$(16) \quad q_{it} = \log\left(\frac{M_{it}}{D_{it}}\right)$$

$$(17) \quad p_{it} = \log\left[\left(\frac{P_{it}^{d}}{P_{it}^{e} X}\right)\left(\frac{1}{1+\tau_{it}}\right)^{\theta_{it}}\right]$$

$$(18) \quad \theta_{it} = \tilde{\alpha}_{it} \big/ \left(1 + \tau_{it} - \tilde{\alpha}_{it}\tau_{it}\right)$$

$$(19) \quad \tilde{\alpha}_{it} = \overline{\tau}_{it} \big/ \tau_{it}$$

A stochastic version of equation (12) is represented in (20), where, to simplify the notation, we drop the product index *i*. This convention will be followed from now on and is justified by the fact that we apply the same methodology to all sectors. The elasticity is estimated for each of the products in isolation and individually, to be consistent with the Armington hypothesis that specifies zero cross-elasticities between all products.

$$(20) \quad q_t = \mu + \sigma p_t + \varepsilon_t \quad \text{where} \quad \mu = \kappa + \sigma \log\left(\frac{\delta}{1-\delta}\right)$$

$$\text{and} \quad \varepsilon_t \approx N\left(0, \upsilon^2\right)$$

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14 The Perron test for the *changing growth model* assumes, in the null hypothesis, the existence of a unit root and a change in the intercept of the stochastic process at the time of the structural break. The alternative hypothesis is that the process is stationary with a change in the slope of the deterministic time trend at the time of the break.

15 The significance level of the ADF test indicates the probability of incorrectly rejecting the existence of the unit root. We adopted the 10% level as a compromise solution, owing to the well-known low-power property of the ADF statistic, i.e. a bias towards non-rejection of the unit root when in fact is not present. A lower significance level would reduce the power of the test even further in this relatively small sample.

Each of these series can be integrated. When a unit root is not present, the series may or may not be stationary, but the procedure employed is the same, regardless, and is based on the assumption that the series is I(0). Table IX.1 presents the four possible combinations of the order of integration of the two series, along with the model employed in each case. As can be seen by comparing the lines in table IX.1, it is the order of integration of *q* that determines whether the estimation is done in terms of levels or first differences, because it is the dependent variable in equation (20).

Table IX.1
DECISION TABLE FOR THE TYPE OF MODEL USED IN ESTIMATION

Quantities *(q)*	Prices *(p)*	
	I (0)	I (1)
I (0)	A: levels	C: levels
I (1)	B: differences	D and E: cointegration

Source: Prepared by the authors.

All of the models we estimate also include as an exogenous variable the coefficient of variation of the ratio between the prices of domestic and imported varieties of the good.[16] This allows the uncertainty surrounding that relative price to affect the ratio between the amount imported and the amount produced domestically. The expected sign on its coefficient depends on the net effect of the speculative mechanism affecting imports, which may be positive or negative. For example, firms that depend heavily on imported inputs may react to greater uncertainty in their expected relative import costs by increasing their imports (positive effect) or else by substituting for them (negative effect). One cannot therefore anticipate the significance of this variable in equation (5), or the sign of its coefficient.

We also use control variables to take account of several important exogenous factors, as follows. The first is a dummy variable to capture the stepwise response of the quantity imported following the 1990 foreign trade liberalization. Its value is therefore $d_t = 1$ for $t \geq 1990{:}4$ and $d_t = 0$ for other periods. The second control variable is a time trend to capture other factors that may have provoked structural changes in the quantum of imports without affecting the relative price of imports. The third is a vector of seasonal dummies (z_t). The inclusion of a time trend and the dummy variable can be rationalized as an attempt to take account of variations in the quality of the goods and the composition of the sector price and quantity

16 The coefficient of variation is the ratio between the standard deviation of the variable and its mean. We chose this measure as a measure of variability because it preserves the non-dimensional nature of equation (2).

aggregates that could not be adequately considered when constructing the quantity index. Examples include imports of electro-electronic goods and personal computers, which grew strongly in the later years of the period, but for which there was also a significant quality change. Our formulation assumes that part of those changes occurred progressively throughout the period, while others happened suddenly in response to the change in the foreign-trade regime; and it allows the empirical equations to distribute these effects among the variables.

Lastly, the estimated equation is shown in (21), which includes all the effects discussed above.

$$(21) \quad q_t = \mu + \sigma p_t + \lambda d_t + \xi \cdot \mathbf{z_t} + \gamma t + \varepsilon_t$$

In the estimation of all models mentioned in table IX.1 we start with the most general specification, assuming the maximum number of lags for the price variable; and we progressively eliminate the non-significant variables to arrive at the final equation. In the next section we discuss the estimation of each of the models mentioned in table IX.1.

1. Model A

The simplest case is when both series are stationary, and we can obtain the long-term elasticity in equation (21) from a regression on the level variables. The equation is initially estimated by ordinary least squares; but, when the Durbin-Watson statistic indicates the existence of first-order serial correlation among the residuals, it is re-estimated using the maximum likelihood method, assuming a first-order autoregressive structure for the errors. This provides estimates of the coefficients and confidence intervals for the parameters of equation (6), and for the parameter of the autoregressive term (ρ), which allows us to calculate the long-term Armington elasticity σ/(1-ρ).

In cases where this procedure suggests the possible existence of a unit root on the residuals, i.e. the confidence interval of ρ includes 1, the equation is re-estimated in first difference terms, in the form of equation (7), which also includes lagged values of the price variable among the explanatory variables. The number of lags included in the equation is the same as used in the procedure to determine the order of integration of the price series, and may be zero.[17]

$$(22) \quad \Delta q_t = \mu + \sigma \Delta p_t + \sum_{l=0}^{\tau} \nu_l \, p_{t-l} + \lambda d_t + \xi \cdot \mathbf{z_t} + \gamma t + \varepsilon_t$$

[17] Appendix B shows how we used a sequence of chained tests to endogenously obtain the number of lags used in the ADF test.

2. Models B and C

Cases where the order of integration of the two series is not the same are hard to rationalize from an economic point of view. Moreover, these unbalanced equations are quite troublesome to estimate. This difficulty has been noted by other authors, who nonetheless recognize the need to overcome the problem in the best possible way.[18] Below we indicate how we treat the two unbalanced cases of table IX.1.

When q is I(1) and p is I(0), the equation is estimated in terms of first differences, as in (22). This avoids the possibility of spurious correlation, because differentiation produces stationary series.[19] When q is I(0) and p is I(1), we estimate the equation in terms of levels, including as many lags as those used in the tests of order of integration, plus one, as indicated in equation (23).

$$(23) \quad q_t = \mu + \sigma\, p_t + \sum_{l=0}^{\tau+1} \nu_l\, p_{t-l} + \lambda\, d_t + \boldsymbol{\xi} \cdot \mathbf{z_t} + \gamma\, t + \varepsilon_t$$

The asymmetric treatment of these two cases is justified by the need to deal with the "integratedness" of the dependent variable, when it has that property. On the other hand, the differencing operation we perform when it is integrated does not impair the estimation when the price variable is already stationary; and it preserves the possibility of interpreting the coefficient on the price variable as the elasticity of substitution. Any signs of serial correlation among the residuals when estimating equation (23) are dealt with by using the same procedure as in Model A.[20]

When the procedure described above is unable to produce an elasticity that is significantly different from zero, we try to estimate it by using the co-integration model described in the next section. We call this case Model E. This procedure is adopted even though the series have not been classified as integrated; but this can be justified in two ways. The first is that there is a margin of error in the tests of order of integration described at the start of this section and in appendix B, which may have led to rejection of the unit root for one of the series, when it is in fact present. The second argument has already been put forward above: there is no entirely satisfactory procedure available to deal with the case of unbalanced equations; and each of the procedures entails a compromise.

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18 See, for example, Maddala and Kim (1998, p. 252): "Should one estimate unbalanced equations? Of course not, if it can be avoided. But if it has to be done, one has to be careful in their interpretation and use appropriate critical values".

19 This procedure is also adopted in Gallaway, McDaniel and Rivera (2003).

20 This procedure is the estimation via quasi-first differences, as described in the previous section (equation (22)).

In our opinion these arguments justify the attempt to estimate the equation by cointegration methods when the other methods fail to find an elasticity that is significantly different from zero.

3. Models D and E

When prices and quantities are integrated, the cointegration relation provides an estimate of the long-term Armington elasticity, for which we use the general formulation contained in Johansen (1988). We write equation (20) in vector notation as equation (24).

$$\text{(24)} \quad \boldsymbol{\beta}'\mathbf{x}_t - \boldsymbol{\mu} = \boldsymbol{\varepsilon}_t$$

where $x_t' = (p_t, q_t)$, $\beta' = (1, -\sigma)$. This VAR model, can be put in a restricted form as a vector error correction (VEC) model which can be written as equation (25) when there is no time trend and the variables are lagged by just one period,

$$\text{(25)} \quad \Delta\mathbf{x}_t = \boldsymbol{\alpha}\cdot\left(\boldsymbol{\beta}'\mathbf{x}_{t-1} - \mu\right) + \boldsymbol{\varepsilon}_t$$

where β is the co-integration vector and α is a vector that contains the weights applied to components of the cointegration term, and is used to adjust the value of x; in other words, it is the vector or coefficients of the error correction term. The vector of residuals ε must be i.i.d. with mean zero and variance matrix Ω.

The VEC of equation (25) can be generalized and written as equation (26) by including *k* lags of the first difference of the vector of variables, and including the exogenous variables used to obtain equation (21): a *dummy* variable that captures the shift in the intercept caused by the trade liberalization, a time trend, and seasonal *dummy* variables:

$$\text{(26)} \quad \Delta\mathbf{x}_t - \boldsymbol{\gamma} = \boldsymbol{\alpha}\left[\boldsymbol{\beta}'\mathbf{x}_{t-1} - \left(\boldsymbol{\beta}'\boldsymbol{\gamma}\cdot(t-1) + \lambda d_t + \mu\right)\right] + \sum_{l=1}^{k-1}\boldsymbol{\Gamma}_l\left(\Delta\mathbf{x}_{t-l} - \boldsymbol{\gamma}\right) + \boldsymbol{\varepsilon}_t$$

In our case, the matrices Γ_τ are 2x2 and contain the weights of the autoregressive components of the process. We chose the number of lags to be included in the equation, represented by *l*, so as to maximize the likelihood statistic for the system of equations.[21] In equation (26), γ is a

[21] The number of lags was reduced progressively starting from a maximum of eight quarters, until the remaining terms were significant. To choose the maximum number of lags we assumed that the effects of a given shock would mostly have been absorbed by the system within two years.

2x1 vector containing the time-trend parameters for the growth of the variables. Thus, $\beta'\gamma \cdot (t-1)$ is a scalar term that shows how the time trend of prices and quantities affects the cointegration relation.

Since the cointegration relation was normalized with respect to the quantities (the first dimension of vector x), one can interpret the term in parentheses in equation (26) as the long-term effect that would occur if the distribution parameter of the CES function in equation (1) had a time trend and were independent of the foreign trade regime. That dependence is captured in our formulation through d_t, the dummy variable that captures the effect of liberalization; and it is represented by its effect on a generalized distribution parameter of the CES formulation, δ_t which is defined (implicitly) by equation (27).

$$(27) \quad \log\left[\delta_t / \left(1-\delta_t\right)\right] = \boldsymbol{\beta}'\boldsymbol{\gamma} \cdot (t-1) + \lambda d_t + \mu$$

To summarize, equation (26) takes account of the major shifts that may have occurred in the demand function for imports, based on the hypothesis that the elasticity of substitution σ was constant throughout the period.

D. Results

We applied the procedure described above to identify the order of integration of the series and to choose the most suitable model to estimate data from the period 1985-2002, for the 28 sectors of the Brazilian input-output matrix where imports were positive in 2002, except for agriculture (including livestock) and the service sectors.

Table IX.2 sets out the types of series considered in our classification, in terms of their stochastic properties and the code adopted for each one. We also show, for each variable in the model, the frequency with which each type of series was encountered. Only 16 quantum series and 11 price series do not have a unit root; but, of these, only six quantum series and five price series are stationary. For 10 quantum series and 17 price series, we are able to find evidence of a unit root.[22] For two quantum series we cannot rule out the existence of a unit root. Lastly, there is evidence of a structural break in the fourth quarter of 1990 for 10 quantum series and six price series.

[22] It is possible to put forward theoretical arguments against the possibility that a price series is integrated. However, we admit that if they behave like integrated series in our sample, it is preferable to treat them as such in the estimation.

Table IX.2
TYPOLOGY OF QUANTUM AND PRICE SERIES

Code	Type	Number of series	
		Quantum	Price
1	Stationary around a non-zero average	2	3
2	Stationary around a zero average	–	1
3	Stationary around a linear trend	4	1
4	Has a unit root with zero time trend	10	17
5	Has a unit root with non-zero time trend	–	–
6	The existence of a unit root cannot be rejected	2	–
7	Does not have a unit root	10	6
–	Evidence of the existence of a structural break in 1990:4	10	6

Source: Prepared by the authors.

The classification of model types presented in table IX.1 shows that most cases refer to situations where the order of integration of the price and quantity series coincide: this happens in 11 instances of Model A (estimation in terms of levels), and 11 of Model D (cointegration). There are six cases of unbalanced equations, but we estimated four of them using cointegration (Model E) since Models B and C did not produce estimates that were significantly different from zero in these cases. There are also two cases where the estimation through models B and C was satisfactory — one of each type.

The results, in table IX.3, show that the estimate of the Armington elasticity has the correct sign and is significantly different from zero for 20 sectors at the 5% significance level. For two sectors it is significant only at the 10% level, and for two others it is significant only at 20%. For one sector the estimated value is significant but its sign is incorrect (negative); and for the three remaining sectors the estimated elasticity is not significantly different from zero.

The coefficient of variation of prices proved significant in just two sectors, and in both cases it is positive, indicating an increase in the share of imports relative to domestic production, in response to greater uncertainty in the relative price of imports. The lack of significance of this variable in most sectors was somewhat surprising, because we had expected it to be significant in several sectors.

The dummy variable that captures the occurrence of a structural break in 1990:4 was significant at the 5% level in 11 sectors, and at 10% in one other, thus confirming the importance of the point discussed in section 2 regarding the nature of the impact of the liberalization that began in 1990. Its coefficient is positive for eight sectors, where the proportion of imports increased, and it is negative in the other four.

When interpreting the coefficient on this variable it is important to remember that part of the impact of liberalization appears in the equation as a tariff reduction and is therefore already taken into account in the estimated value of the Armington elasticity. The dummy variable captures the rest of the impact of liberalization, which can be attributed to other factors such as the existence of repressed demand for imports, which was revealed when non-tariff barriers were removed on that occasion.[23]

The coefficient of the time trend variable is significant in 20 sectors, and is positive in all cases except one. This is consistent with the interpretation that during the period 1986-2002 there was an increase in the relative demand for imports that is not explained by the other three factors; and it is possibly related to the modernization and internationalization of the basket goods produced and consumed by domestic industry.

Table IX.4 summarizes the values we obtained for the Armington elasticity for the different the sectors, classified as: very high, high, average, low, zero or negative. The range of values in each category is not uniform because the purpose of the classification is to provide an indication of the curvature of the indifference curve between imports and domestically produced goods, for each sector. This curvature does not vary linearly with the elasticity, however, as can be seen in figure IX.1.

Figure IX.1 is also useful for illustrating the difficulty of estimating the Armington elasticity when its value is small and the indifference function has a high degree of curvature. This is the case in the two sectors where the elasticity is classified as low, and in the three cases where it is zero or negative (table IX.4). Considering that, at each point on the plane in figure IX.1, the elasticity of substitution is the ratio between the slope of the indifference curve passing through that point and the slope of the line segment that connects it to the origin of the coordinate

23 The presence of repressed demand is a possible explanation for the sign of the dummy variable in imports of automobiles, trucks and buses, tractors and machinery, other vehicles and autoparts, textiles and clothing. The negative value of that variable can be rationalized for cases where imports did not grow by as much as expected, given the tariff reduction. This may have occurred because of pricing policies implemented by domestic producers to control import penetration in the following sectors: the rubber industry, electronic equipment, miscellaneous chemical products and mineral extraction.

Table IX.3
ARMINGTON ELASTICITIES FOR BRAZIL, 1986-2002

Sector	Classification[a]		Structural break[b]		Model[c]	Share (%)[d]	Estimated coefficients (t-statistic in brackets)			
	Quantum	Price	Quantum	Price			Substitution	Variation	Break	Trend
Other food products and beverages	4	4	-	-	D	1.9	3.59	5.13	-	-
							(6.12)	(1.58)	-	-
Textile industry	4	4	-	-	D	2.0	3.36	-	0.65	0.03
							(9.66)	-	(3.52)	(6.53)
Miscellaneous industry	4	4	-	-	D	4.3	2.42	-	-	0.02
							(6.08)	-	-	2.82
Clothing articles and accessories	4	4	-	-	D	0.4	2.23	-	0.63	0.05
							(8.07)	-	(1.90)	(6.07)
Rubber industry	4	4	-	-	D	1.3	2.16	-	-0.54	0.01
							(7.61)	-	(-3.11)	(2.21)
Meat preparation and animal slaughtering	7	7	-	-	A	0.3	2.03	-	-	-
							(3.30)	-	-	-
Wood products and furniture	4	4	-	-	D	0.4	1.86	-	-	0.03
							(6.04)	-	-	(5.26)
Tractors and machinery	4	4	-	-	D	12.3	1.78	-	0.89	-
							(8.29)	-	(7.84)	-
Plastics	4	4	-	-	D	0.6	1.75	-	-	0.03
							(9.54)	-	-	(5.23)
Other metallurgical products	4	4	-	-	D	1.9	1.50	-	-	0.03
							(6.84)	-	-	(5.99)
Milk and derivatives	7	1	-	-	A	0.7	1.47	-	-	-
							(2.53)	-	-	-
Automobiles, trucks and buses	3	3	S	S	A	4.3	1.43	-	1.60	0.06
							(2.80)	-	(2.88)	(3.61)
Processing of plant products and tobacco	7	7	-	-	A	2.5	1.18	-	-	-
							(3.80)	-	-	-
Petroleum refining and petrochemicals	6	4	S	-	D	10.8	1.18	-	0.53	0.01
							(4.22)	-	(3.49)	(3.81)
Paper, pulp and print	4	2	-	-	B	2.0	1.01	0.71	-	-
							(6.11)	(1.77)	-	-
Metallurgy of non-ferrous materials	7	1	-	-	A	2.1	0.98	-	-	0.01
							(5.04)	-	-	(8.09)

(Continues)

Table IX.3 (concluded)

Sector	Classification[a]		Structural break[b]		Model[c]	Share (%)[d]	Estimated coefficients (t-statistic in brackets)			
	Quantum	Price	Quantum	Price			Substitution	Variation	Break	Trend
Non-metallic minerals	7	2	S	S	A	0.9	0.75	-	0.55	0.02
							(4.00)	-	(3.43)	(5.17)
Vegetable oils and edible fats	3	7	S	S	A	0.5	0.61	-	-	-
							(1.77)	-	-	-
Steel	7	4	-	-	E	1.3	0.57	-	-	0.02
							(2.16)	-	-	(4.24)
Other vehicles, parts and accessories	6	4	S	-	D	8.9	0.41	-	0.32	0.02
							(2.27)	-	(2.26)	(7.27)
Pharmaceuticals and perfumery	7	4	-	-	C	3.8	0.40	-	-	0.03
							(3.62)	-	-	(11.34)
Electrical materials	7	4	-	-	E	5.2	0.36	-	-	0.02
							(1.94)	-	-	(10.03)
Petroleum, natural gas, coal and other fuels	1	7	S	S	A	1.0	0.27	-	0.45	-0.03
							(1.34)	-	(2.45)	(-5.95)
Electronic equipment	3	4	S	-	E	12.0	0.16	-	-0.46	0.01
							(1.37)	-	(-4.67)	(5.72)
Non-petrochemical chemicals	3	2	S	S	A	5.3	0.36	-	-	0.03
							(1.23)	-	-	(5.28)
Miscellaneous chemicals	7	7	S	S	A	4.6	0.14	-	-0.24	0.02
							(0.58)	-	(-2.28)	(7.82)
Footwear, leather articles and fur	7	4	S	-	E	0.5	-0.18	-	-	0.02
							(-0.54)	-	-	(5.21)
Mineral extraction	1	7	-	-	A	1.8	-1.34	-	-0.70	-
							(-3.24)	-	(-3.02)	-
Weighted average share of imports							**0.93**	**0.11**	**0.20**	**0.02**

Source: Prepared by the authors.

Notes: For 60 degrees of freedom and bilateral confidence intervals, $t_{20\%} = 1.296$, $t_{10\%} = 1.671$ and $t_{5\%} = 2.000$.

[a] Code for series classification:
1 – Stationary around a non-zero average
2 – Stationary around a zero average
3 – Stationary around a linear trend
4 – Has a unit root with zero time trend
5 – Has a unit root with non-zero time trend
6 – The existence of a unit root cannot be rejected
7 – Does not have a unit root

[b] Code for structural break: S when there is structural break in 1990:4

[c] Code for estimated model (see text for details):
A – Regression on levels
B – Regression on first differences
C – Regression on levels, with lagged dependent variable
D – Cointegration model between integrated series
E – Estimation by cointegration owing to zero elasticity in models B and C

[d] Share of the sector in Brazil's total imports, average 1997-2002

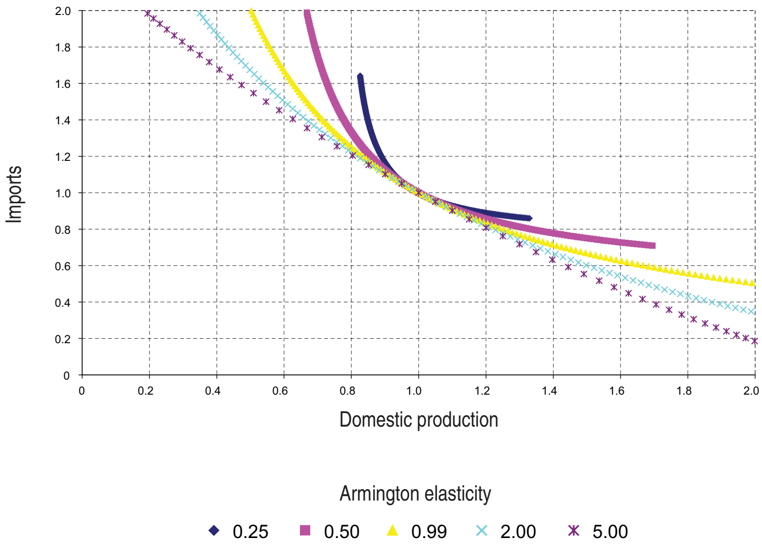
Table IX.4
RANGES OF THE ESTIMATED ELASTICITIES AND INTERNATIONAL COMPARISON

Sector	This paper			USITC and GTAP		
	Category	Definition	Elasticity	Average	USITC	GTAP
Other food products and beverages	Very High	$\sigma \geq 3$	3.59	3.2	4.2	2.2
Textile industry			3.36	2.3	2.3	2.2
Miscellaneous industry			2.42	2.3	1.7	2.8
Clothing articles and accessories			2.23	3.2	2.0	4.4
Rubber industry			2.16	2.0	2.0	1.9
Meat preparation and animal slaughtering	High	$1,5 \leq \sigma < 3$	2.03	2.5	2.7	2.2
Wood products and furniture			1.86	2.8	2.8	2.8
Tractors and machinery			1.78	2.5	2.2	2.8
Plastics			1.75	2.0	2.0	1.9
Other metallurgical products			1.50	3.5	4.1	2.8
Milk and milk derivatives			1.47	3.6	5.0	2.2
Automobiles, trucks and buses			1.43	4.0	2.7	5.2
Processing of plant products and tobacco			1.18	3.3	3.5	3.1
Petroleum refining and petrochemical industry			1.18	2.2	2.5	1.9
Paper, pulp and print	Average	$0,5 \leq \sigma < 1,5$	1.01	2.9	3.9	1.8
Metallurgy of non-ferrous materials			0.98	3.6	4.4	2.8
Non-metallic minerals			0.75	2.7	2.5	2.8
Vegetable oils and edible fats			0.61	3.6	5.0	2.2
Steel			0.57	3.5	4.1	2.8
Others vehicles, parts and accessories			0.41	4.0	2.7	5.2
Pharmaceuticals and perfumery			0.40	2.0	2.0	1.9
Electrical materials	Low	$0 < \sigma < 0,5$	0.36	2.5	2.2	2.8
Petroleum, natural gas, coal and other fuels			0.27	2.8	2.8	2.8
Electronic equipment			0.16	2.7	2.6	2.8
Non-petrochemical chemicals			0.00	2.0	2.0	1.9
Miscellaneous chemicals	Null	$\sigma \approx 0$	0.00	2.0	2.0	1.9
Footwear, leather articles and fur			0.00	3.1	1.7	4.4
Mineral extraction	Wrong sign	$\sigma < 0$	-1.34	2.4	2.0	2.8
Arithmetic mean[a]			1.24	2.8	2.8	2.8
Minimum[a]			0.00	2.0	1.7	1.8
Maximum			3.59	4.0	5.0	5.2
Range			3.59	2.0	3.3	3.4
Standard deviation[a]			0.98	0.6	1.0	0.9

Source: Donnelly et al (2004) and authors' calculations.

[a] Excludes sector where the elasticity has the incorrect sign.

Figure IX.1
INDIFFERENCE CURVES BETWEEN IMPORTS AND DOMESTIC PRODUCTION



Source: Prepared by the authors.

system, we note that when the curvature is high, only a small segment of the indifference curve is spanned when the relative price changes, even when that change is large. This make the estimation more difficult because large variations in p elicit only small changes in q; and in this situation the error term ε in equation (21) is larger. One might therefore expect a large standard error in the estimate of σ in those cases.

Table IX.4 also shows that there are two sectors with very high elasticity ($\sigma \cong 3$), eight with high elasticity ($\sigma \cong 2$), nine average ($\sigma \cong 1$), five low ($\sigma < 0.5$), three zero and one with negative elasticity. The arithmetic mean of the estimated elasticities is 1.24 and their frequency distribution is roughly symmetric.

We now compare the elasticities we estimated for Brazil with those produced by other studies. As noted above, there are no previous estimates for our country; and the difficulty in contrasting them with elasticities for other countries is compounded by differences in the sector classification and the small number of studies available in the international literature. Nonetheless, this obstacle can be partially avoided using the study by Donnelly et al (2004), which presents the Armington elasticities adopted in the applied general equilibrium models of the

United States International Trade Commission (USITC) and the Global Trade Analysis Project (GTAP),[24] using a sector classification similar to the one we adopted.[25]

Although table IX.4 compares these three measures of sector elasticities, the comparison may be of questionable relevance in itself, since there is no *a priori* reason for them to be equal, or even similar, because they reflect country-specific characteristics in the respective consumption and production structures. We proceed with the comparison nonetheless.

Initially we note that the arithmetic mean of the estimated sector elasticities (1.24) is only 44% of the 2.8 average value of the USITC and GTAP sector elasticities. This suggests that substitution between imports and domestic production is more difficult in Brazil than in the United States, or in other hypothetical "conventional" countries, because, to produce the same relative change in import share the change in relative price needed in Brazil is twice as large, owing to the lower elasticity. This is consistent with the perception that, even after liberalization, Brazil still is relatively closed to international trade. Another possible interpretation of this very large difference in elasticities is that those calibrated by USITC and GTAP are in fact too high for a country with our characteristics, and thus do not represent the behaviour of the import share our case.

On the other hand, this difference in elasticity values does not represent a very significant difference in the curvature of the indifference curve, as can be inferred from figure IX.1 by noting that the curves for $\sigma = 1$ and $\sigma = 3$, are very close to each other.[26]

As can be seen by comparing the ranges and standard deviations of the sector-level Armington elasticities in our study and those of USITC and GTAP, shown in the bottom lines of table IX.4, the variability of our estimates is similar to those reported by Donnelly et al (2004).

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[24] The United States International Trade Commission (USITC) developed a CGE model for the United States to evaluate the impacts of changes in trade policy in that economy. Its elasticities refer to the United States, and they were obtained from the literature and later calibrated by sector experts. The Global Trade Analysis Project (GTAP) developed a global-scope multi-regional model which is used to evaluate the global impacts of trade agreements. The generic elasticities obtained from it are used by the model when specific values are not available for a given country. They are derived from the SALTER project of the Australian Industry Commission (Huff, 1997) and other data contained in the literature.

[25] In sectors for which we were able to find a direct correspondence with those in our study, we transcribed the values directly. For the other sectors we repeated the values of that study for the broader sector classification.

[26] Although the curve for elasticity equal to 3 is not shown in the figure, to preserve clarity, its position can be easily inferred from the curves for elasticities equal to 2 and 5, as between these two but closer to the former.

Nonetheless, the fact that the minimum value of the sector elasticities is so high (1.7 and 1.8 for USITC and GTAP, respectively) clearly stands out in comparison to the zero value we found for Brazil. There is also a significant numerical difference in the maximum value of the sector elasticities (3.6 versus 5). Nonetheless, that does not imply a large difference in the curvature of the indifference curve, which is practically zero in both cases, since substitution is near perfect in those sectors, both in Brazil and in the countries represented in those two CGE models.

Lastly, it is important to note that the differences between our estimates and those reported in Donelly et al (2004) may be due not only to the characteristics of the countries, but also to differences in the methodologies used to obtain them. The calibration of elasticities, based on expert opinions expressed in relation to sector studies using the Delphi methodology, may have led the institutions in question to eliminate outlying or uncommon values, as a result of error-risk aversion on the part of the specialists consulted.

We can conclude this international comparison by making a more general assessment of the comparison of sector elasticities, noting that the discrepancies are substantial and more frequent than the coincidences. This eloquent evidence advises against a procedure that is frequently encountered in the literature, whereby the impacts of trade and exchange-rate policies are analysed using elasticities calibrated on the basis of values adopted for other countries, on the assumption that differences are insignificant. This procedure is not valid for Brazil, at least for the period we have analysed, and it may easily lead to false conclusions.

E. Conclusions

This paper has estimated a new set of Armington substitution elasticities for the 28 industrial sectors of the Brazilian input-output matrix, for the period 1986-2002. We develop an estimation methodology that measures the effects on observed data of the trade restrictions that existed before 1990, and the impact of the trade liberalization that began in that year. The methodology also carefully examines and takes into account the stochastic and dynamic properties of the variables involved, and chooses the estimation method so as to be consistent with those properties. We speculate that this methodology could be applicable to other countries that have undergone trade liberalization; and we argue that this is relevant and needs to be included in the estimation data for the trade liberalization period, since it contains information that is very important for estimating the elasticity of substitution and the curvature of the indifference curve between imports and domestic production.

The Armington elasticities we estimate have the correct sign; and they are significant at the 5% level for 20 sectors, at 10% for two sectors, and at 20% for two others. In one sector the estimated value is significant but has the incorrect sign (negative). Although the estimated elasticity is not significantly different from zero in three sectors, these represent only 12% of the average total value of imports in the period 1997-2002. The point estimate of the elasticity of substitution ranges from 0.16 to 3.6 in the sectors where it is positive and statistically non-zero; and its weighted average value is 0.93, the weights being the value of sector imports.

A classification was used to group sectors according to the range of variation of the substitution elasticity. This shows that the Armington elasticity is very high in two sectors, high in eight, average in nine, low in five, zero in three, and negative in one. The arithmetic mean of their value is 1.24 and the frequency distribution of the elasticities is roughly symmetric.

The international comparison shows that the average of the sector elasticities we obtain is only 44% of those used in the USITC and GTAP general equilibrium models; but this numerical difference in elasticity values only has a small effect on the curvature of the indifference curve between imports and domestic production. Nonetheless, there are large sector differences in elasticity values when the estimates for different sectors are compared individually; and there are also differences in the minimum and maximum elasticities across sectors.

Lastly, we believe that using these elasticities will enable researchers to more precisely evaluate the economic impacts of a change in trade policy, in both partial and general equilibrium models.

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Annex A

Source and treatment of the data

We used quarterly data for each sector of the Brazilian input-output matrix (IBGE – level 50), for the period 1986-2002.

The price and quantum indices (P_i^m) and (M_i), respectively, are those produced by Fundação Centro do Comércio Exterior (Funcex), using the methodology described in Markwald et al (1998); they are available in the electronic database system IPEADATA (www.ipea.gov.br). The exchange rate (e) is the monthly average of the official dollar selling price. We approximated the domestic price index (PD_i) by the corresponding wholesale price index, *Índice de Preço no Atacado (Oferta Global)* calculated by the Getulio Vargas Foundation (IPA-OG-FGV), having reconciled its sectors with those of the input-output matrix described in table A1. We calculate the average price index in cases where an activity of the input-output matrix corresponds to more than one IPA sector, using a weighted average when the necessary data was available, otherwise the simple average.

The coefficient of variation of the relative price (P_i^d / P_i^m) measures the effect of uncertainty and was calculated as the ratio between the standard deviation and the average of this price ratio over a six-month "window" centred on the median month of the period in question.

The domestic sales *quantum* index (D_i) was estimated by deflating the value of domestic sales for each sector (VDT_i) by the corresponding domestic price index (P_i^d).

The sector VDT_i was calculated by deducting the value of exports from the corresponding sector-production value (VP_i), which was inferred from its value in the most recent input-output matrix, together with the variation in the production and price indices between the year to which it refers and the date for which the calculation is being made. As data availability prior to 1990 is limited, the procedure was slightly adapted in the earlier period, as follows.

For 1986-1990 the value of total domestic sales (VDT_i) was estimated by equation (A1), which shows that the value for each month was calculated by applying the observed monthly variations in the quantum and price indices to the average value of total domestic sales in 1985, and then deducting the value of exports for the respective month. This uses the domestic production *quantum* index calculated by IBGE (www.ipea.gov.br) for each sector of the matrix, adjusted to the aggregation used here as described in table A2.

Table A1
RECONCILIATION BETWEEN THE SECTORS OF THE INPUT-OUTPUT MATRIX AND THE IPA INDUSTRY CLASSIFICATION

Sector of input-output matrix (level 50)	Sector of the IPA-OG-FGV (column)
Mineral extraction	Mineral extraction (28)
Petroleum, natural gas, coal and other fuels	Fuels and lubricants (54)
Non-metallic minerals	Limestones and silicates (30)
Steel	Iron, steel and derivatives (32)
Metallurgy of non-ferrous materials	Non-ferrous metals (33)
Other metallurgical products	Total metallurgical (31)
Tractors and machinery	Machinery and industrial equipment (36)
Electric material	Total electric material (38)
Electronic equipment	Electric material and others (41)
Automobiles, trucks and buses	Motor vehicles (43)
Other vehicles, parts and accessories	Motor vehicles (43)
Wood products and furniture	Wood (45), total furniture (46)
Paper, pulp and print	Paper, paperboard (50)
Rubber industry	Rubber (51)
Non-petrochemical chemicals	Chemicals and others (58)
Petroleum refining and petrochemical industry	Total chemicals (53)
Miscellaneous chemicals	Total chemicals (53)
Pharmaceuticals and perfumery	Pharmaceutical products (81), perfumery, soaps and candles (82)
Plastics	Plastics (56), plastic products (83)
Textile industry	Natural fabrics and yarns (60), man-made fabrics and yarns (61), knitted or crocheted fabrics (62)
Clothing articles and accessories	Clothing (63)
Footwear, leather articles and fur	Footwear (64)
Processing of plant products and tobacco	Plant products (71)
Meat preparation and animal slaughtering	Meat and fish (78)
Milk and milk derivatives	Milk and milk derivatives (79)
Vegetable oils and edible fats	Vegetable oils and fats (74)
Other food products and beverages	Salt, animal feed and others (80), beverages (66)
Other miscellaneous industries	Total manufacturing industry (29)
Meat preparation and animal slaughtering	Meat and fish (78)

Source: IBGE and FGV. Prepared by the authors.

$$\text{(A1)} \quad VDT_{it} = \left(\frac{VP_{i85}}{12}\right) \cdot \left(\frac{q_{it}}{q_{i85}}\right) \cdot \left(\frac{P_{it}}{P_{i85}}\right) - VE_{it}$$

where:

VDT_{it} = value in R$ of total sector *i* domestic sales in month *t*;

VP_{it85} = value in R$ (base price) of sector *i* production in 1985;

q_{it} = index of sector *i* physical production in month *t*;

q_{it85} = index of sector *i* physical production, monthly average in 1985;

P_{it} = index of sector *i* domestic price in month *t*;

P_{i85} = index of sector *i* domestic price, monthly average in 1985; and

VE_{it} = value in R$ of sector *i* exports in month *t*.

After 1991, the procedure described above for the reference year 1985 was repeated, but using previous year's average values as the base, because the value of domestic production each year is available in the input-output matrix for 1991-1996, and in the National Accounts for 1997-2002.

We used a two-step procedure to calculate the nominal tariff in each sector τ_{it}. First, we distributed the products and respective tariffs obtained from the foreign trade classification table —the Brazilian Merchandise Nomenclature: Harmonized System *(NBM-SH)* and the Common Mercosur Nomenclature *(NCM-SH)*— for each sector (level 80) of the input-output matrix. Next, we calculated the average nominal tariff for each activity in the input-output matrix (level 50), weighted by the value of production of each sector (level 80) belonging to each activity (level 50).

The effectively paid tariff series ($\bar{\tau}_{it}$) was calculated as the ratio between tariff revenue and the total value of imports for each category of use, using data obtained from the Brazilian Internal Revenue Service (SRF/MF). This was adjusted to be consistent with the sector classification of the input-output matrix.

Table A2
RECONCILIATION BETWEEN THE SECTORS OF THE INPUT-OUTPUT MATRIX AND THE INDUSTRY CLASSIFICATION

Classification	Sector: IBGE (g) or matrix (m): 1986-1990	Sector: IBGE (g) or matrix (m): 1991-1999
Mineral extraction	Non-metallic minerals extraction (m)	Mineral extraction (g)
Petroleum, nat'l gas, coal & other fuels	Petroleum and natural gas extraction (m)	Petroleum and natural gas extraction (m)
Non-metallic metals	Non-metallic metals products (g)	Petroleum and natural gas extraction (g)
Steel	Flat-rolled steel products (m)	Steel (m)
Metallurgy of non-ferrous materials	Basic metallurgy (g)	Metallurgy of non-ferrous materials (m)
Other metallurgical products	Other metallurgy (g)	Other metallurgical products (m)
Tractors and machinery	Mechanical products (g)	Mechanical products (g)
Electrical materials	Electrical and communication materials (g)	Electrical machinery and equipment, including household appliances (m)
Electronic equipment	Electrical and communication materials (g)	Material for electronic & comm. equipment (m), and TVs, radios & sound equipment (m)
Automobiles, trucks and buses	Automobiles (m)	Automobiles, trucks and buses (m)
Other vehicles, parts & accessories	Autoparts and accessories (m)	Autoparts and accessories (m)
Wood products and furniture	Total manufacturing industry (g)	Wood (m) and furniture industry (m)
Paper and pulp	Paper and paperboard (g)	Paper and paperboard (g)
Rubber industry	Rubber (g)	Rubber industry (g)
Non-petrochemical chemicals	Total chemicals (g)	Chemical, non-petrochemical or carbon-based (m) and alcohol distilling (m)
Petroleum refining & petrochem. ind.	Petrochemicals, coal refining (g)	Petroleum refining (m), basic & interm. petrochem. (m), resin, fibres & elastomers (m)
Miscellaneous chemicals	Other chemicals (g)	Fertilizers (m) and miscellaneous chemical products (m)
Pharmaceuticals and perfumery	Pharm. & vet. prod's (g) & perf., soaps & candles (g)	Pharmaceutical industry (m) and perfumery, soaps and candles (m)
Plastics	Plastics (g)	Plastics (g)
Textile industry	Textiles (g)	Textiles (g)
Clothing articles and accessories	Clothing, footwear and fabrics (g)	Clothing articles and accessories (m)
Footwear, leather articles and fur	Footwear (m)	Footwear (m)
Processing of plant prod's & tobacco	Food products (g)	Processing of rice (m), wheat milling (m) & processing of other plant prod's for food (m)
Meat prep. & animal slaughtering	Meat preparation & animal slaughtering (m)	Meat preparation and animal slaughtering (excl. poultry) (m) & meat prep. poultry (m)
Milk and milk derivatives	Milk derivatives (m)	Preparation of milk and milk derivatives (m)
Vegetable oils and edible fats	Oil refining and edible fats (m)	Natural vegetable oils (m) and refining and edible fats (m)
Other food products and beverages	Food products (g) and beverages (g)	Other food industries (m) and beverage industry (m)
Miscellaneous industries	Total manufacturing industry (g)	Total manufacturing industry (g)

Source: IBGE, prepared by the authors.

Annex B

Determination of the order of integration of the price and quantum series

We used the methodology proposed by Enders (1995) to determine the order of integration of the price and quantity series involved in the estimation equation, complemented by the Perron (1989) test to deal with the possibility of structural breaks in the series.

We initially estimate equation (B1), which contains a trend, a constant term and autoregressive components; and we test for the existence of a unit root ($\gamma = 0$), using the augmented Dickey-Fuller (ADF) statistic.[27] If that hypothesis is rejected, we conclude there is no unit root and terminate the search.

$$\text{(B1)} \quad \Delta x_t = a_0 + \gamma x_{t-1} + a_2 t + \sum_{i=1}^{p} \beta_i \Delta x_{t-i} + \varepsilon_t$$

As this is a low-power test, if the unit root cannot be rejected we must also test the joint hypothesis of its existence and the absence of a trend ($a_2 = \gamma = 0$), using the Dickey-Fuller ϕ_3 statistic (1981). If this joint hypothesis is rejected, we test again for $\gamma = 0$, using a normal distribution, and the procedure is then ended. If this joint hypothesis cannot be rejected, we assume that we can cast the data-generating process in the form of equation (B2), and we again test for a unit root with the ADF statistic.

$$\text{(B2)} \quad \Delta x_t = a_0 + \gamma x_{t-1} + \sum_{i=1}^{p} \beta_i \Delta x_{t-i} + \varepsilon_t$$

If the null hypothesis of a unit root is rejected in this specification, we terminate the procedure. If it cannot be rejected, we test for the joint null hypothesis $a_0 = \gamma = 0$ using the Dickey-Fuller ϕ_2 statistic (1981). If joint hypothesis is rejected, we test again for $\gamma = 0$, using the normal distribution, and the procedure is completed. If the hypothesis $c_1 = \gamma = 0$ is not rejected, we test for the existence of the unit root in the specification of equation (15), again using the ADF statistic. If $\gamma = 0$ is accepted (rejected), we conclude that the series contains (does not contain) a unit root.

$$\text{(B3)} \quad \Delta x_t = \gamma x_{t-1} + \sum_{i=1}^{p} \beta_i \Delta x_{t-i} + \varepsilon_t$$

[27] The critical values for the ADF statistics were taken from Hamilton (1994) for a 10% significance level.

In equations (B1), (B2) and (B3), the number of lags (p) was chosen according to the *general-to-simple* criterion, starting with a maximum of five. The fifth lag is retained if it is significant at the 5% level. Otherwise, we re-estimate the equation with four lags, and again assess the level of significance of the last lag. The procedure continues until the coefficient of the last autoregressive component is significant at the 5% level.

It should be noted that the results of the tests described above may not be conclusive if there is a structural break in the series; in that case the ADF statistic has a bias towards non-rejection of the unit root. To account for this and take into consideration the likelihood of a structural break in the fourth quarter of 1990, we apply the Perron (1989) test to series displaying a unit root. Using the taxonomy proposed by that author, we assume the break is of the type represented by the *changing growth model*. Equation (B4) describes this model, and accommodates both the null and the alternative hypothesis of the test. In the null hypothesis, a unit root is assumed with a change in the intercept of the process at the time of the structural break. The alternative hypothesis assumes that the process is stationary with a change in the slope of the deterministic trend line at the time of the break.

$$\text{(B4)} \quad x_t = \mu + \theta U_t + \beta t + \gamma T_t + \alpha x_{t\text{-}1} + \sum_{i=1}^{p} \beta_i \Delta x_{t\text{-}i} + \varepsilon_t$$

where:

T_B = date of the structural break;

$DU_t = 1$, if $t > T_B$ and $DU_t = 0$; and

$DT_t^* = t - T_B$, if $t > T_B$ and $DT_t^* = 0$, otherwise.

The null hypothesis imposes the following restrictions on the parameters of equation (B4):

$\alpha = 1$, $\gamma = 0$, $\theta \neq C$

The alternative hypothesis imposes the following restrictions on the parameters of equation (B4):

$\alpha < 1$, $\gamma \neq 0$, $\theta = C$

We assumed that the structural break occurred in the fourth quarter of 1990, and the critical values used were those of Perron (1989), with a 10% significance level. We applied the test sequentially, adding autoregressive components until the hypothesis of residual autocorrelation was rejected in the Ljung-Box test, at a 5% significance level.