

A Comprehensive Empirical Analysis of Trade Policy in a Small Country with Monopolistic Competition*

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Abstract

A comparative advantages model with monopolistic competition is proposed to perform a comprehensive analysis of the effects of protectionism on competition, economies of scale and resource allocation, plus a non-cost competition term. Evidence is based on Brazil's period of import substitution industrialization. The foreign economy is a set of (integrated) developed countries, which both improves the accuracy of the comparative advantage calculations and raises access to data for the computations of fixed costs. Only the period under protection is considered, so that the comparative statics, aimed to corroborate the above policy effects, are drawn on within-period counterfactuals.

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1 Introduction

We attempt a comprehensive industry-level analysis of trade-policy effects under imperfect competition, building on the efficiency basis of comparative advantages (Deardorff, 1980). Evidence is based on Brazil's import-substitution industrialization (ISI) from the late 1960s to the late 1980s, which offers a rich experience for examining this policy issue.

To assure a robust index of Brazil's revealed comparative advantages (RCA), we define the foreign economy as a set of integrated developed countries, which further avoid biases from characteristics (e.g., trade costs) we are not controlling. Although this three-country (adding the rest of the world) framework assumes away the geographic basis of trade (Eaton and Kortum, 2002), it makes ample room to define and compute comparative advantages.

In theory, imperfect competition and trade-policy barriers can weaken and even invert comparative-advantage linkages, as given by the negative N-industry correlation between net exports (or RCA) and countries' characteristics (Deardorff, 1979, 1980). However, to avoid misleading statistical inferences, we both use that integrated foreign economy and expand the period to the entire ISI era. This long time span also enables us to refrain from disputable statistics experiments with the ensuing trade-openness period, which was largely affected by simultaneous non-trade related reforms in Brazil, making the isolation of trade policy difficult.¹ Our comparative-static exercises are, instead, based on within-sample counterfactuals that also allow simultaneously addressing some related development issues over that ISI period.

At the same time, using the pre-1980 years greatly reduced the availability of internationally comparable data. The number of total employees, for instance, is the only direct marginal-cost information available for the seven included countries. This prompted us to develop a latent opportunity-cost variable, based on the impact of unobserved differences in factor proportions. Additionally, given that Brazil's wide-ranging protectionism cannot be adequately characterized by either nominal tariffs or the effective rate of protection (ERP), we attempt a counterfactual based on a more accurate (than

¹A standard problem in developing countries' reform, as stated by Trefler (2001) and witnessed in Tyler and Gurgel (2008).

RCA) trade-performance measure: the revealed comparative efficiency in manufacturing (RCEM), which provides more conclusive evidence about the trade-policy effect on allocative-efficiency.

Underlying the RCA index is the possibility of exporting from sectors without comparative advantages, which we draw from monopolistic competition (i.e., intra-industry trade) that eventually introduces a non-price competition argument through product-differentiation. The latter relationship is empirically captured by an industry-level variable proxying changes in world preferences.

Moreover, the spatial monopolistic competition (Lancaster, 1984; Schmitt, 1990) also enables, from its non-constant markup pricing, both the scale (productive efficiency) and the pro-competitive effects, which are, respectively, the firm-size and the industry-size effects from protection (Feenstra, 1995). In the Chamberlinian monopolistic competition (Melitz, 2003), the productive effect stems instead from selection of heterogeneous firms in productivity, as empirically supported by previous studies (Tybout et al, 1991; Tybout, 1993; Head and Ries, 1999), rather than from elimination of homogeneous plant.

Unlike these quoted partial-equilibrium empirical analyzes around a production function, ours is around a transformation function – from its dual (i.e., costs) – that focuses in inter-industry trade as well, similarly to the Chamberlinian analysis by Bernard et al (2007). The consequent introduction of comparative-cost advantages, as control for assessing non-observed fixed cost, outweighs the inferior technological accuracy of an industry-level analysis.

We consider both corporate and plant fixed costs, where the former is proxied by a variable expressive of each industry's economic activity, whereas the latter, aimed to express the activity level (i.e., size) of a firm, is proxied by a composite variable of the operative workers input and the average firm size. The integrated foreign economy is crucial for assessing these fixed costs, since relevant data is not available for all developed countries.

A derived variable of market structure proxies the pro-competitive effect, the effective rate of protection, an association that is reinforced by further statistics experiments.

Three policy effects are then examined: on allocative efficiency (opportunity cost), on productive efficiency (average cost), and the pro-competitive effect (pricing).² The non-price competition term

²From Melitz and Ottaviano (2008), which rests on a Chamberlinian quasi-linear utility, we draw another comprehensive

has no definite relationship with policy, nor with the varieties effect to consumer welfare.(Arkolakis et al, 2008). It is worth stressing an empirical complementarity in identification: marginal costs help to single out the pricing and fixed cost terms, while the latter reduce the risk of spurious comparative-advantages linkages from trade policy. As considered below, this is a convention in the the empirical IO literature, for identifying fixed costs and imperfect competition.

The statistics results of both the exploratory and regression analyzes corroborate all those positive basis of economic inefficiency, showing that Brazil's ISI regime not only caused extreme allocative inefficiency, as already shown in Tyler (1985), but also other inefficiencies related to higher average costs and market power.

The paper is structured as follows. The models are worked out in Section 2, followed by a description of the empirical variables in Section 3. In Section 4, an exploratory statistical analysis briefly describes Brazil's experience, while Section 5 presents the basic regression results, and Section 6 takes up further statistical experiments. Conclusions follow.

2 Theory and Empirical Specification

We start with the closed-economy, focusing on both market conduct and the temporal change in industry sizes, and eventually shift to the international-economy model, focusing on relative export sizes, as initially given by both comparative costs and distorted prices, and subsequently by inefficient firm entry.

2.1 Industry Size in Autarky

Consider an economy having a competitive sector, y , produced with unskilled labor, and N manufacturing industries X_i , each producing horizontally differentiated varieties with unskilled and skilled labor under internal increasing returns to scale. Consumers are heterogeneous in their preferences to

analysis of international upon economic performance.

varieties, as described by this upper-level utility function:

$$U(y, x_i, d_i) = y^{1-\xi} \prod_{i=1}^N x_i^{\xi_i}, \quad \xi = \sum^N \xi_i < 1, \quad (1)$$

where $x'_i = x_{i\omega}/h(d_{i\omega})$ is the quality-adjusted quantity a specific consumer attains from the most preferred variety, $x_{i\omega}$, after applying the compensating function, $h(d_{i\omega}) \geq 1$, in the distance $d_{i\omega}$ from her *ideal variety* (see Lancaster, 1979, 1984). The quality space is in a unit circle, where the n_i firms (varieties) are symmetrically spaced, so that $d_i = 1/n_i$, and the heterogeneous consumers are uniformly distributed. Aggregating those for whom $p_{i\omega}h(d_{i\omega})$ is minimized, one reaches the firm's clientele, whose price-elasticity $\sigma_i(n_i)$ is fully determined by the number of firms³.

Given (1), the aggregate demand for each manufactured product i , after accounting for the symmetric (in price and size of firms) zero-profit equilibrium, will be

$$X_i = S_i \left(\frac{1}{\theta_i c_i} \right), \quad i = 1, \dots, N, \quad (2)$$

where $S_i = \xi_i Y$ is the size of industry i , irrespective of prices, and Y is the available income of the economy. The denominator within brackets is the optimum price (relatively to the numeraire y): marginal cost, c_i , times the markup $\theta_i = [1 - 1/\sigma_i(n_i)]^{-1}$.

Normalizing (2) by $Y = X$, yields:

$$x_i = \xi_i [\theta_i(\mathbf{w}\mathbf{a}_i(\mathbf{w}))]^{-1}, \quad (3)$$

where $x_i = X_i/X$ is the relative sales of manufacturing industry i and w and a_i are, respectively, the factor-price and factor-input vectors in marginal cost.

To cast (3) in a time dimension, we add subscript t to each variable and substitute ξ_i by $\eta_{it} = d \log \xi_{it} / d \log Y_t$, yielding:

$$x_{it} = \eta_{it} [\theta_{it}(\mathbf{w}_t \mathbf{a}_{it}(\mathbf{w}_t))]^{-1}. \quad (4)$$

As indicated, the η_{it} terms come from temporal changes in consumer preferences for products.

³This result was noted by Helpman and Krugman (1985, 6.3), and demonstrated by Cinquetti and Balistreri (2010) in a fully specified model, using $h(d)$ as in Lancaster (1984). Other developments of this approach to trade theory are Schmitt (1990) and Vogel (2008), with heterogeneous firms and non-local competition.

2.2 Export and Opportunity Cost under Free Trade

Suppose we partition the economy into several non-symmetric countries that are integrated only by free international trade in goods, without geographic barriers. Then the share of each country k in the world market of i , x_i^k , can be drawn from (4), adjusting its arguments to country characteristics.

If total employees and their wages are the sole evidence, in our database, of marginal costs, $\mathbf{w}_t^k \mathbf{a}_{it}^k(\mathbf{w}_t^k)$, rather than a full vector of input uses and prices, how can we further characterize opportunity costs in each country? The Rybczynski theorem provides the clue. That is, controlling for fixed-input requirement in a sector i , \bar{v}_i , the relationship between a positive factor-endowment change in v_l and relative output change in a small country is:

$$\hat{x}_i = FT\hat{v}_l > 0, \quad \hat{x}_j = FT\hat{v}_l < 0, \quad (5)$$

where x_i and x_j are sectors intensive and not intensive in v_l , respectively, whereas FT is the transformation function. Similar relationships hold for the foreign economy. Therefore, if v/v^* is the ratio of factor endowment of home and foreign countries, its relationship with the vector of relative exports, x^T/x^{T*} , can be indirectly conveyed by the following correlation with the vector of relative output, x/x^* :

$$\text{corr}[(x_i^T/x_i^{T*}), x_i(v)/x_i^*(v^*)] = \gamma, \quad \gamma > 0. \quad (6)$$

Hence, γ indirectly conveys the efficiency relationship between comparative exports and factor proportions (Deardorff, 1980), with the relative size x_i/x_i^* expressing the latent opportunity (or marginal) cost. In general, if a country k , has an abundance of factors in which good i is relatively intensive, it will tend to have a higher share in the world output (and exports) of that good.(Harrigan, 1997).

2.3 Fixed Costs, Protection and Comparative Exports

We assume a separable unit cost function in the marginal and the fixed costs. The latter encompasses both plant and corporate fixed cost, $G_i(q_i)$ and $F_i(q_i)$, which are associated with fixed input

coefficient of unskilled and skilled labor respectively(see Markusen and Venables, 2000):

$$c_i(w^k, q_i) = a_i^k(w^k)w^k + (G_i/q_i)w^k + (F_i/q_i)w_s^k. \quad (7)$$

Unlike the fixed costs, the input coefficient of marginal cost, $a_i^k(w^k)$, is irrespective to firm's output, q_i . Yet, the unit values of the former vary with the unskilled and skilled-labor prices, w^k and w_s^k respectively. In a developing country, the corporate fixed cost is mostly related to technology transfer, rather than technology generation, so we also test home's corporate fixed cost with an expected value $F_i^h < F_i$.

We now introduce domestic trade-policy barriers (and incentives), T_i^k , and assume it to impact on the product markets alone. This makes the markup in good i , which is a function of trade policy, $\theta_i^k(T_i^k)$, higher than the measure of economies of scale, ψ_i , in proportion to the policy revenues (or positive profits). That is

$$\theta_i(T_i) = \left(1 - \frac{1}{\sigma_i(n_i^T)}\right)^{-1} > \frac{c_i(w, q_i)}{c_{iq}(q_i, w)} = \psi \equiv \theta_i, \quad (8)$$

where $c_{iq}(q_i, w)$ stands for marginal cost and n_i^T for the number of firms with protected trade. Higher markup causes a rationalization on industry output, as can be drawn from (4), inversely expressing the pro-competitive effect from international trade (Markusen, 1981). The alternative cost impact from T_i^k is discussed next.

We may now substitute (6)-(8) into (4), with the latter defined in terms of each country's relative supply (exports) to the rest of the world (ROW), x_{it}^T and x_{it}^{T*} , which introduces ROW as a third country. In this international context, S_{it} (and S_{it}^*) must be replaced by its corresponding industry size in the international economy: $S_{it} = \delta_i(n_{it})\xi_t^{\eta_{it}}Y_t^w$, where $\delta_i(n_{it})$ stands for the home economy's share in the world sales of i (there is a similar share for foreign) and Y^w for the world income. Lastly, the transformed equation (4) is rewritten as comparative exports and then linearized into the following stochastic form:

$$x_{it}^T/x_{it}^{T*} = \alpha_i + (\delta - \delta^*)\eta_{it} - \beta_2 \left(\frac{w_t a_{it}}{w_t^* a_{it}^*}\right) + \beta_3 \left(\frac{Y_{it}}{Y_{it}^*}\right) - \beta_4 \tilde{G}_{it} - \beta_5 \tilde{F}_{it} - \beta_6 T_{it} + \mu_{it}, \quad (9)$$

where α_i stands for unmeasured industry-specific characteristics, μ_{it} for the random error, and the

subscript of $(\delta - \delta^*)$ were dropped since these coefficients are constrained to be the same across industries.

We thus have a comparative cost and pricing model, where $\beta_1 = (\delta - \delta^*)$, expressing the relative increase in home's exports as world preferences change over time, stands for a non-price competition term. Besides implying that the size of each regional industry is proportional to the number of competitive varieties, this relationship requires that η_{it} , the rank of industries growing most rapidly in international markets, is correlated to the intensity of non-price competition.⁴ A possible extension is that η_{it} is proportional to activities intensive in skilled-labor, so that $\beta_1 > 0$ could be related to a stride in this factor's endowment by the developing country (Currie et al., 1999).

The coefficient β_2 captures the marginal comparative cost advantages and should be negative, while β_3 captures the latent opportunity cost attempted in (6) and should be positive. Policy distortions weaken these comparative-advantage linkages, however the analytical method of identifying this *allocative inefficiency* from β_2 and β_3 is left to Section 6. Next, the fixed-cost $G_{it}(= G_{it}w^k/y_{it})$ and $F_{it}(= F_{it}w_s^k/y_{it})$ are affected by both factor proportion (or prices) and a certain home-market effect, though their empirical form, as explained bellow, measures only cross-industry differences. To some degree, the association of β_6 with pricing distortions is partially granted by the fact that both β_2 and β_3 capture the parallel effect of T on cost distortions. However, a negative impact from trade policy T will have to be further decomposed into cost and the pricing (market-power) distortions, which is done in Section 6.

⁴Suppose firms are identical in size, q , so $X_i^k = n_i^k q_i$, where k refers to either home or foreign. we can then rearrange $X_i^k = \delta_i^k \xi_i Y^w / \theta_i^k \mathbf{w}_i^k = \delta_i^k \xi_i Y^w / p_i^k$ to

$$\delta_i^k = p_i^k n_i^k q_i (\xi_i Y^w)^{-1},$$

making clear the association between δ_i^k and n_i^k , once p_i^k has already been accounted for. Consider now the international form of the normalized temporal equation (4):

$$x_{it}^k = \delta_i^k \eta_{it} (p_i^k)^{-1}.$$

If the inter-period changes in varieties are internationally uneven (e.g., Grossman and Helpman, 1991, ch. 9), then $\dot{n}_i \geq \dot{n}_i^* \Rightarrow \delta_i^k(\eta_{it})$, and so $\delta - \delta^* \leq 0$ reflects countries' relative positions in this non-price competition.

2.4 Protection and Productive Efficiency

Protection can also affect the right-hand side of (8) by inducing inefficient entry of firms, and so increasing average cost (Horstmann and Markusen, 1986), $c_i(w, y)$. Despite some analytical limits of this homogenous spatial monopolistic competition (see Vogel, 2008), its symmetric distribution of firms across regional markets (see Schmitt, 1990; Lancaster, 1984), makes the scale effect less dependent of the hypothesis of non-segmented market (Markusen and Venables, 1988).

Similar to most developing countries in the analyzed period (see Santos-Paulino, 2002), Brazil's protectionism encompassed a myriad of trade policy (tariff barriers and several non-tariff barriers) and industrial-policy instruments. Assuming that T is proxied effective rate of protection (or nominal tariffs), it can thus capture just part of the effects from those policy, so this additional scale effect should be referred to unobservable (in our empirical model), \bar{T} .

Plant fixed cost, G_{it} , is the natural index of economies of scale, since its technology coefficient does not change internationally, unlike F_{it} . Using the separability in factor content in (7), the whole imperfectly trade-policy effect on each G_i and θ_i can be thus decomposed:

$$\Theta_i(T, \bar{T}) = G_i/x_i[n_i(\bar{T}_i)] + \sigma[N_i(T_i)], \quad (10)$$

where $N_i = n_i + n_i^*$ stands for the numbers of firms (varieties) in i . Under the hypothesis of free entry, the average costs effect responds to unobserved instruments \bar{T}_i , by means of adjustment in the number of local firms, n_i , whereas market power responds to T_i , by means of adjustment in the number of varieties in the market, N_i .

Replacing the internationally equal G_i by the local G_i^n , and so transforming model (9) into:

$$x_{it}^T/x_{it}^{T*} = \alpha'_i + (\delta - \delta^*)'\eta_{it} - \beta'_2 \left(\frac{w_t a_{it}}{w_t^* a_{it}^*} \right) + \beta'_3 \left(\frac{Y_{it}}{Y_{it}^*} \right) - \beta'_4 G_{it}^n - \beta'_5 F_{it} - \beta'_6 T_{it} + \varepsilon_{it}. \quad (11)$$

we test the impact the mentioned economies of scale effect. More to the point, in this counterfactual to (9), the inefficient entry, causing $G_{it}^n > G_{it}$, is identified by the weakening linkage of plant fixed cost with RCA . That is, using the true, local-adjusted fixed cost, causes $\beta'_4 > \beta_4$.⁵ An exploratory

⁵The non-homothetic cost function (7) does not compromise the entailed framework, since the comparative cost relationship does not rely on autarky prices.

analysis is previously made to identify if firm size is, indeed, beneath the $G_{it}^n > G_{it}$.

3 Variables and Data

To discuss the empirical specification, let us transform (9) and (11) to a nominal form:

$$\begin{aligned} RCA_{it} = & \alpha_i + \beta_1 WYEL_{it} - \beta_2 CPCOST_{it} + \beta_3 SIZE_{it} - \beta_4 PLANT_{it} - \beta_5 CORPO_{it} \\ & - \beta_6 FPROT_{it} + \epsilon_{it}, \quad i = 1, \dots, 20 \quad \text{and} \quad t = 1, \dots, 4 \end{aligned} \quad (12)$$

where RCA_{it} (revealed comparative advantages) = x_{it}/x_{it}^* , $WYEL_{it} = \eta_{it}$, $CPCOST_{it} = (w_t a_{it})/(w_t^* a_{it}^*)$, $SIZE_{it} = y_{it}/y_{it}^*$, $PLANT_{it} = G_{it}$, $CORPO_{it} = F_{it}$ and $FPROT_{it} = T_{it}$. Alternative to $CPCOST$ is $CPROD = (a/a^*)$ and to $PLANT$ is $PLANTBR = G^n$. The three-digit twenty manufacturing industries – with some adjustments to available data – are described below, while the four years are 1967, 1973, 1980, and 1987-88 (average, due to the extreme disturbances of these two years), with slight deviations for some variables. Pre-1980 years made the dearth of international compatible data more stringent, the sources of which are described in the Data Appendix.

The then six largest industrialized economies (USA, Japan, Germany, UK, France, and Italy) make up the foreign economy, which assures less-biased RCA_{it} and comparative costs than one built on a single developed country, given their large differences in size and factor endowments⁶. Hence, in the RCA, $x_{it}^* = \sum_{it} (X_{it}^j / X_t^j)$, X_{it}^j stands for the j th foreign country's exports of i and X_t^j for its total manufactured exports at t .

Variation in i 's world-market demand, $WYEL_{it}$, is given by:

$$\eta_{it} = \frac{X_{it}^w / X_{i,t-1}^w}{Y_t^w / Y_{t-1}^w},$$

where X_i^w is the world's exports of i , $Y_t^w = \sum_i X_t^w$ the world total exports of all products (i.e., not only manufactured), the $t - 1$ obliges us to take 1963 data. Y_t^w can be thought as proxying the world output (income) of tradable-goods sectors.

⁶This multi-country aggregation is an alternative to the multiple regressions made by Golub and Hsieh (2000).

Marginal and fixed costs, which are not directly observable, as in most inter-industry studies (Bersnahan, 1989), are taken as distinct components of total labor input. Accordingly, marginal comparative labor costs is given by:

$$CPCOST_{it} = \frac{a_{it}w_t}{a_{it}^*w_{it}^*} = \frac{(l_{it}/x_{it}) \cdot w_t}{\left(\sum_j l_{it}^j / \sum_j x_{it}^j\right) \cdot w_{it}^*},$$

where l/x stands for “total employees/value added”, and w and w^* are the manufacturing wages in constant US dollars of Brazil and foreign. A pure productive measure, $CPROD_{it}$, is obtained by dropping w_t/w_{it}^* . The social opportunity cost, $SIZE_{it} = x_{it}/x_{it}^*$, based on (6), makes room for some cross-time scale (or home-market) effects.

Similarly to Brainard (1997), corporate fixed cost, $CORPO_{it}$ ($CORPBR_{it}$), is proxied by the ratio “office labor/total employees”, but instead of proxying plant fixed cost with operative labor input alone, we expand it by an additional measure of economies of scale: average firm size. Thus, plant fixed cost is defined as:

$$PLANT_{it} = \frac{\left(\sum_j l_{it}^{u*} / x_{it}^*\right) \cdot N_{it}^*}{\bar{G}_t^*}$$

where l_{it}^{u*} and x_{it}^* stand, respectively, for operative workers employment and output in the US industries, N_{it}^* for the number of firms therein. Hence, the unitary plant fixed cost, $PLANT_{it}$, is given by the operative labor input of the average firm (in size) in each industry. The normalization by \bar{G}_t^* , the yearly average of the numerator, removes a likely general higher size of foreign firms, so that this regional difference reflects only relative cross-industry difference. By transforming them into a stationary panel data, it also removes a temporal home-market (size) effect, or the difference of $PLANT_{it}$ from $PLANTBR_{it}$. This plant fixed cost in Brazil is similarly calculated to $PLANT_{it}$.

Variable $CORPO_{it}$ aims to assure the separability of fixed costs related to economic activity, thus us disregarding the output level and considering only the intensity of skilled (office) labor (see Antweiler and Trefler, 2002). Without ignoring the advantage of firm-level data, this corporate fixed cost, which is absent in the empirical studies on trade policy and its effects on either plant elimination or selection (Head and Ries, 1999; Tybout et al, 1991; Feenstra, 2003), helps to isolate

plant fixed costs; to reinforce the attempted meaning underlying the comparative impact of $PLANT_{it}$ and $PLANTBR_{it}$. Moreover, as conventional in the empirical IO literature Berry and Reiss (2007); Bersnahan (1989), an inter-industry study about these fixed costs becomes more robust if the model has controls for marginal costs – absent in the mentioned empirical studies – and pricing (or market conduct) as well.

A derived market-structure variable (Schmalensee, 1989) proxies market power stemming from import tariffs (and export subsidies), $FPROT_{it}$: the effective rate of protection in Brazil. The reason is straightforward: the literature makes a direct association between protection and market power. Nonetheless, given the likely impact on higher opportunity costs as well, further regression experiments are performed, in Section 6, to prove that $FPROT_{it}$ stands mostly for prices rather costs distortions. Nominal tariffs in Brazil, $TNOM_{it}$, are also tested for robustness. Disregarding the foreign economy, whose correspondent panel data were not available, amounts to assuming it as if under free trade as compared to Brazil – quite reasonable for that period

Policy endogeneity is dismissed in (11) on the ground that protection in Brazil was unrelated to sectors' comparative advantages (Gonzaga et al., 2006).

4 Trade Policy in Brazil

A brief overview of Brazil's policy experience is useful and enables us to better grasp of some of our variables. We begin it with a graphical analysis of a centered RCA, as in Benedictis (2005) :

$$b_{it} = \frac{RCA_{it} - 1}{RCA_{it} + 1},$$

with $-1 \leq b \leq 1$, where positive (negative) values, from $RCA > 1$ (< 1), indicate comparative advantages (disadvantages). These b_{it} are further classified into the four technology groups (Lall, 2000): **RB (resources-based)**: food products, beverages, paper & paperboard, rubber, non-metallic minerals, wood & cork; **LT (low technology)**: furniture, leather & furs, clothing & shoes, metals and textiles; **MT (medium technology)**: transport equipment, plastics, printing & publishing, mechani-

cal equipment, chemicals and tobacco⁷; **HT (high technology)**: other chemicals, electrical material and other sectors.

The b_{it} are plotted in the below diagrams, each having the original and final periods on the horizontal and the vertical axes, respectively, so that points below the diagonal indicate industries whose final RCAs were smaller than the original ones. In 1967, Brazil had comparative advantages in only two manufacturing industries,

The bit are plotted in the below diagrams, each having the original and final periods on the horizontal and the vertical axes, respectively, so that points below the diagonal indicate industries whose final RCAs were smaller than the original ones. In 1967, Brazil had comparative advantages in only two manufacturing industries, although the concentration of points above the diagonal, in both figures, shows a steady upward movement. However, all seven sectors having comparative advantages by 1987-88 belonged to either the *RB* or the *LT* groups.

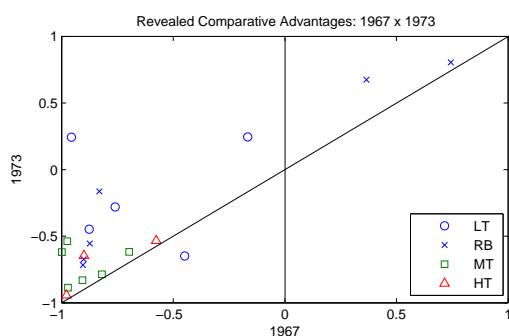


Figure 1

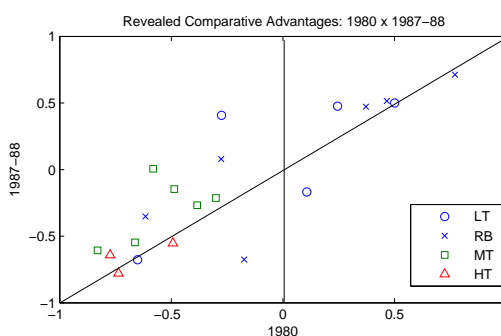


Figure 2

However, the Brazilian HT and MT industries, whose trade performance improved the least, were among those having the highest output growth, which indirectly suggest the allocative inefficiency of this inward-growth experience.

The whole picture becomes clearer once we consider the evolution of factor endowments. Table 1 below shows that Brazil's proportion of skilled to unskilled labor, relatively to the developed countries, did not change from 1967 to 1980, having even decreased relative to arable land. This slow pace of human capital formation is a key difference between this industrialization strategy and that of

⁷This manufacturing sector is highly concentrated around few firms due to a high expenditure on advertising.

the Asian NICs – see UN, *Human Development Report 1999* and World Bank, *World Development Indicators 1998*.

Table 1: Factor Endowments: Brazil/Developed Countries

Countries	Skilled/Unskilled Labor*		Skilled Labor/Land**	
	1967	1980	1967	1980
Brazil	0.07	0.12	57.2	59.8
Developed Countries	0.15	0.26	96.8	297

Source: for 1967: Bowen et al (1987); for 1980: The World Institute Resource (1998) and World Bank, *World Development Report 1982*.

*Skilled Labor: 1967, percentage of clerical and management in the economically active population; 1980, complete secondary education as % of relevant age group.

**Land: Arable in hectares.

Trade protection is other factor to consider, starting with the average (and standard deviation) of the effective rate of protection: 79.7 (45.2) in 1967; 34.1 (32.4) in 1973; 36.0 (53.4) in 1980; and 41.4 (51.6) in 1987-88. It is worth noticing that Brazil’s GDP and relative (to the world) total export grew the most in the only period of steady and general fall in the FPROT, 1967-73. Protectionism resumed strongly afterwards. Note that the slight increase in the average ERP by 1980 was accompanied by a sharp increase in the standard deviation, from 32.4 to 53.4. This period featured an erratic sectoral trade policy, which included negative protection in some industries (see also Tyler, 1985; Savasini, 1983) and typified the uncontrolled consequences of expanding trade barriers without policy coordination. In particular, huge export subsidies, in many cases aimed at compensating the anti-export bias of the import-substitution policy, were the main trait of this new policy (Bruton, 1989; Moreira, 1995).

5 Estimation Results

The main goal of the ensuing regression analysis is estimating the qualitative effects of trade policy, including the counterfactuals seeking to identify those effects via the shifts in the parameters. Accordingly, we work with centered variables: $z_i - \bar{z}_i$, where the “within” mean is $\bar{z}_i = \sum_t z_{it} / \sum_i$ (of industries i) which further avoids the scale nature of some variables, and apply a WGLS-White estimator to models (9) and (11), as justified in the Statistical Appendix.

As shown in Table 2, all variables are statistically significant in most of model specifications, in spite of some regressions having low \bar{R}^2 , which can be attributed to both the small sample and high number of regressors. Average values of the fixed effects, for models (i)-(iv), without the dummies, and (v)-(ix), respectively, clearly characterize industries' components of the RCA_{it} , as indicated by the slight change in their ordering with the dummies.

The negative WYEL confirms that Brazil did not thrive in the world's most expansive markets. Given the empirical form of WYEL and all cost and pricing controls, this result suggests that Brazil did not respond well to global non-price competition. Assuming that these demand-expansive sectors were high-tech intensive, the failure in this favored target of the import substitution (ISI) policy can be explained by Brazil's sluggish human-capital formation (Bruton, 1989, see). The assumption that these were high-tech sectors is not necessarily accurate, though.

The coefficient of CPCOST is generally lower than that of CPROD and insignificant in some cases. This likely reflects the sudden impact of sizeable devaluation in 1987-88 on manufacturing wages in Brazil, relative to the composite foreign economy. CPROD does not contain wages and is simply a comparison of relative (inverse) labor productivity. However, in models (i) to (iv), the positive partial correlation between comparative cost and relative exports stands for an extreme "allocative inefficiency", confirming previous studies of Brazil's ISI (Tyler, 1985; Savasini, 1983; Bruton, 1989). However, a closer look at resource-based sectors, such as Food and Wood, having high CPCOST, suggests that this variable may misrepresent comparative cost advantages because of the unobserved cost advantages from abundance of non-tradable natural resources.

With a cost-dummy variable for both Food and Wood, in models (v)-(ix), CPCOST and CPROD then become negative. The world is not Ricardian: the comparative advantage linkages (RCA inversely correlated to comparative costs) only show up when other productive factors are taken into account. However, as suggested by the model in column (ix), we cannot rule out that there were significant resource misallocations, which have not been sufficiently controlled by either FPROT or TNOM, to the point of inverting the weak linkages in comparative advantage in the sense of Deardorff

Table 2: Estimates of the Comparative Advantages Model

Independent Variables	Dependent Variable: RCA								
	(i)	(ii)	(iii)	(iv)	(v)	(vi)	(vii)	(viii)	(ix)
YEL	-0.266	-0.208	-0.041	-0.072	-0.221	-0.051	-0.068	-0.055	-0.348
	0.024	0.013	0.017	0.013	0.021	0.018	0.014	0.024	0.048
CPCOST	0.289		-0.042		-0.169	-0.196		-0.109	0.162*
	0.065		0.105		0.080	0.099		0.096	0.091
CPROD		0.207		0.253			0.236		
		0.024		0.027			0.030		
DCPCOST					0.773	0.897		0.647	0.413
					0.211	0.184		0.280	0.247
DCPROD							0.126		
							0.038		
SIZE	2.257	2.245	2.963	3.231	1.632	2.847	3.175	2.657	1.714
	0.375	0.359	0.386	0.323	0.406	0.390	0.311	0.447	0.452
PLANT	-1.283	-1.295	-1.051	-1.115	-1.745	-1.246	-1.328	-1.048	-1.740
	0.186	0.151	0.324	0.216	0.220	0.280	0.228	0.378	0.237
CORPO	3.405	1.975			2.476				4.087
	0.268	0.164			0.215				0.389
CORPBR			0.276	0.101		0.244	0.122	0.286	
			0.042	0.029		0.038	0.029	0.056	
FPROT	-0.183	-0.195	-0.231	-0.205	-0.328	-0.277	-0.238		
	0.020	0.011	0.027	0.018	0.023	0.028	0.022		
TNOM								0.037	-0.249
								0.042	0.064
N. Observations	77	77	77	77	77	77	77	77	77
Adjusted R2	0.739	0.845	0.641	0.687	0.806	0.693	0.722	0.594	0.720
F statistics	48.04	87.74	32.13	38.30	57.12	32.99	37.27	22.85	36.85
Ordered Fixed Effects									
FOOD			7.275				FOOD	0.415	
WOOD			5.21				PAPER	6.252	
CLOTHSHOES			4.169				EQTRANS	5.586	
METAL			2.034				MECH	3.055	
MECH			1.692				PRNTNG	2.469	
LEATFUR			1.358				PLAST	2.364	
PRNTNG			1.286				CHEM	1.765	
PAPER			0.886				FURN	1.261	
DIVERSES			0.488				BEVER	1.236	
FURN			0.457				DIVERSES	1.225	
TEXT			0.439				METAL	1.04	
RUBB			0.302				TEXT	0.812	
ELETR			0.173				LEATFUR	0.755	
NONM			0.087				OTHCHM	0.707	
EQTRANS			-0.097				TOBAC	0.313	
PLAST			-0.427				NONM	-0.093	
TOBAC			-0.908				CLOTHSHOES	-0.593	
OTHCHM			-1.103				ELETR	-0.674	
CHEM			-1.169				RUBB	-0.787	
BEVER			-1.297				WOOD	-0.788	

Heteroskedasticity corrected models (cross-section weights and White covariance matrix). Number in brackets are standard errors.

Statistical significance: bold letter stands for statistical significance of 5% or higher, whereas * stands for 10% significance.

†: Three industry observations for WYEL were not available for 1967, reducing the total to 77.

Description of Variables in Section 3. Data Source: Appendix I.

(1979).⁸ We take up this point at the next section.

Since the fixed-cost variable PLANT stands for an input/output rate, whereas CORPO (and CORPBR) refers to a factor intensity, their coefficients show that Brazil's RCA partially rested on both plant-level economies of scale and skilled-labor intensive sectors. The positive coefficient on CORPBR is unexpected since it measures a Brazilian relative cost. It may reflect Brazil's regional comparative advantages in skilled-labor intensive goods, as happened with Japan's early manufacturing exports (Heller, 1976). This simultaneously means that this variable contains (or is correlated with) unobserved fixed trade costs, which are smaller for exports to neighboring countries.

Another possibility is that PLANT, which incorporates only unskilled labor, under-controlled the skilled labor of plant fixed cost. At the same time, it must be stressed that the coefficients of the developed-country corporate fixed costs exceeded those for Brazilian corporate costs in all cases, which confirms the basic theoretical hypothesis around equation (7).

The negative coefficient of FPROT fits with our identification of the *pro-competitive effect* of international trade: higher wedges between prices and cost reduce international competitiveness and sales. Since this negative impact may express the hypothesis of misallocation of resources towards sectors with higher opportunity costs, an additional analysis is carried out in the next section to single them out. Nominal tariffs, TNOM, a less accurate measure of the effects of trade policy on firm's revenues, had a non-definite impact [columns (v) and (vi)].

The "scale effect" is to be assessed by placing PLANTBR, the plant-fixed cost of the average firm in Brazil, in lieu of PLANT, the counterfactual standing for the world without policy distortions. However, we must make sure that the differences between PLANTBR and PLANT is indeed related to variation in the firm size in the corresponding industries.

A decisive insight is provided in Figure 3, by plotting the ratio "PLANTBR_{it}/PLANT_{it}" against the equally normalized number of firms per industry, $\frac{N_{it}/\bar{N}_t}{N_{it}^*/N_t^*}$, in the vertical axis. The values are in logarithm so as to avoid a large concentration of points around zero and thus attain a more informative diagram. As shown, the cost ratio PLANTBR_{it}/PLANT_{it} is high and positively correlated with

⁸Our model rests on weak links as his, though with no reference to autarky prices.

N_{it}/\bar{N}_{it} , which supports the argument that inefficient entry in Brazil decreased (increased) industry-level economies of scale (average cost).

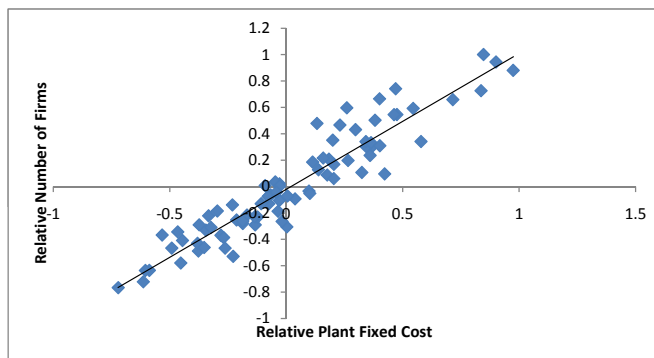


Figure 3: Number of Firms *versus* Plant Fixed Cost

Not only the variations in $PLANTBR_{it}/PLANT_{it}$ is strongly correlated with the relative level of plant-elimination (or entry), N_{it}/\bar{N}_{it} , but also the former variation is largely explained by variation in \bar{x} , the average firms size.⁹ This finding differs from previous studies (Head and Ries, 1999; Tybout, 1993; Treffer, 2001; Fernandes, 2007, ch. 5) rejecting the plant-elimination (or size) hypothesis and showing that only plant-selection can explain variation in productivity. The fact that some of these studies focuses on free-trade agreement, which compels firms to smaller sales in domestic markets, can in part explain their divergence with our findings.

It must be noticed, that Brazil’s legal apparatus for entry (and exit) had no sectorial bias, except for two or three manufacturing sectors in which foreign firms were restricted, which reinforces the role of protectionism – hardly encompassed by FPROT alone – behind the above result.

We can now move to the regression analysis, whose results are shown in Table 3. The substantial increase in the coefficients of $PLANTBR$, relatively to corresponding models with $PLANT$ in Table 2, shows a weaker linkages of fixed-plant cost to comparative (exports) in Brazil, as compared to the claimed free-trade condition. We cannot then reject the hypothesis that the Brazilian industries operated with lower economies of scale,¹⁰ given the observed empirical content of these variables, as well as the additional control for the corporate fixed cost. The statistical insignificance of $PLANTBR$

⁹As carefully show in our working paper Cinquetti and Silva (2010) and partially so in the next section.

¹⁰This is not a test about the minimum efficiency scale (MESS) because monopolistic competition prevents the MESS.

Table 3: Estimates of the Comparative Advantages with Local Scale

Dependent Variable: RCA				
Independent Variables	(i)	(ii)	(iii)	(iv)
WYEL	-0.024 0.007	-0.071 0.012	-0.028 0.009	-0.021 0.010
CPCOST	-0.060 0.057	0.020 0.069	-0.048 0.063	-0.040 0.076
DCPCOST	0.502 0.300*	0.444 0.488		0.382 0.345
SIZE	2.605 0.341	2.312 0.479	2.701 0.322	2.663 0.402
PLANTBR	-0.036 0.123	-0.095 0.136	0.059 0.098	-0.014 0.110
CORPO		0.891 0.160		
CORPBR	0.163 0.048		0.209 0.046	0.191 0.041
FPROT	-0.143 0.015	-0.126 0.023	-0.116 0.013	
TNOM				0.024 0.026
N. Observations	77	77	77	77
Adjusted R2	0.604	0.580	0.551	0.544
F statistics	23.64	21.79	23.65	19.44

Idem Table 2.

is solved in the next section.

The commented empirical analysis on the productive (scale) effect carry, apart from their larger sample, no comparative advantage relationship.¹¹ This means, moreover, having no marginal costs as control for identifying fixed costs (and the entailed economies of scales); a common practise in the empirical IO literature.

6 The Allocative and the Competitive Effects

No definitive evidence about both the allocative and pro-competitive effects has been provided yet. Regarding the former, the negative relationship between comparative cost and trade patterns in Table

¹¹Which is not intrinsic to theoretical models with heterogenous firms (see Bernard et al, 2007).

2 prevents us from stating any microeconomic inefficiency from of this protectionist experience. What remains unanswered, though, is whether unobserved-policy instruments weakened the negative partial correlation between *CPCOST* (*CPROD*) and the RCA.

Theoretically, in countries where manufacturing industries heavily rely on trade and industrial-policy instruments, a given export share of the N supported industries i draws, on average, higher input requirement compared to the same-industry exports from a country closer to free trade. Therefore, one possible answer to the above question is replacing RCA by a measure of trade performance that captures allocative efficiency more accurately, and then check the new coefficient of comparative marginal costs.¹²

More precisely, if \tilde{E} = RCA and E is the alternative vector of trade pattern, then their respective correlation to the vector of comparative opportunity cost (controlled for fixed cost and markup revenues), \mathbf{c}^p , are

$$E = b\mathbf{c}^p, \quad \tilde{E} = \mathbf{b}'\mathbf{c}^p \quad \Rightarrow \mathbf{b} > \mathbf{b}', \quad (13)$$

Since b' was negative (with the cost-dummies) but close to zero, microeconomic inefficiency pushes b up, so that the $b' < b$ expresses the weaker links of comparative advantages stemming from trade-policy distortions in Brazil – that is irrespective to the role of geography on trade.

A germane (to RCA) and more accurate measure of allocative efficiency is the *revealed comparative efficiency in the manufacturing industry*:

$$RCEM_{it} = \frac{(x_{it}^T/x_t^T) / (x_{it}/x_t)}{(x_{it}^{*T}/x_t^{*T}) / (x_{it}^*/x_t^*)}$$

where x_i^T and x_i stand for the exported and total output of i in an economy, respectively. The RCEM index combines information of both production and goods market, and in a way that resembles the efficient partition of the traded and produced output in Deardorff (1980)¹³. In fact, despite us not using *net export*, it can be argued that RCEM is superior in two senses: (i) data of imports increases

¹²Bernard et al (2003) employ a similar theoretically driven statistical experiment to obtain indirect evidence of trade costs.

¹³Though comparative costs does not rely on autarky prices. Cinquetti and Silva (2008) apply a similar variable to access the relative efficiency of manufacturing industry in a set of developing countries prior to the 1980s debt crisis.

the problem of sectorial consistency in aggregations of production (ISIS) and goods trade (SITE); and (ii) imports are as distorted by trade policy as exports.

Table 4: Estimates of the Comparative Efficiency Model

Independent Variables	Dependent Variable: RCEM					
	(i)	(ii)	(iii)	(iv)	(v)	(vi)
WYEL	-0.136 0.047	-0.076 0.035	-0.118 0.036	-0.045 0.024*	-0.036 0.025	-0.094 0.019
CPCOST	-0.031 0.038	-0.111 0.059		0.049 0.054	0.154 0.058	
DCPCOST	1.176 0.120	0.741 0.211		0.907 0.091	0.674 0.157	
CPROD			0.308 0.051			0.284 0.038
DCPROD			0.141 0.049			0.091 0.071
SIZE	-0.031 0.051	0.220 0.180	0.523 0.188	0.251 0.076	0.406 0.169	0.658 0.158
PLANT	-1.010 0.169	-0.473 0.305	-0.889 0.206			
PLANTBR				-0.148 0.050	-0.179 0.047	-0.158 0.073
CORPO						
CORPBR	0.543 0.051	0.384 0.104	0.291 0.076	0.382 0.056	0.441 0.077	0.114 0.069*
FPROT	-0.352 0.023		-0.286 0.022	-0.260 0.016		-0.189 0.015
TNOM		0.081 0.043*			0.086 0.036	
N. Observations	77	77	77	77		77
Adjusted R2	0.786	0.371	0.646	0.492	0.424	0.556
F statistics	50.78	11.82	27.46	16.60	13.64	20.21

Idem Table 2.

In Table 4 we have the result of the regression model (12) with $RCEM_{it}$ replacing RCA_{it} , aimed to test the hypothesis held in 13, or else to identify artificial (costly) export successes. As shown, the coefficient of CPCOST moved significantly upward, as compared to similar models in Tables (2) and (3). Indeed, in columns (iii)-(vi), the coefficients of comparative costs (and productivity) are positive, which can be coined an extreme microeconomic inefficiency. The significant fall in the coefficient of SIZE, in comparison to the earlier regressions, corroborates this weaker relationship

between opportunity cost and trade pattern, when the latter is adjusted for efficient specialization.

We now move to decomposing the distortions from FPROT, which would be a pure cost distortion, associated with allocative inefficiency, and a pure pricing distortions, which, by increasing the wedge from prices to costs, affects the sector's size. The former effect is partially captured by the both CPCOST and SIZE, which would allow us to associate the FPROT coefficient with the remaining pro-competitive effect, but this still does not provide a definitive identification of the impact of protection on market power.

We can use the implicit-function (to the RCA model) between FPROT and either opportunity cost or market power, which ultimately affect marginal cost, to isolate the above two effects from FPROT. Opportunity cost would be represented by both CPCOST and SIZE, whereas PLANTBR a fixed cost, which ultimately grants the pricing (market power) distortions. Naming the former by x_1 and the latter by x_2 , the sought implicit functions can be uncovered with stepwise regression (Greene, 2000), by eliminating FPROT, x_3 , from the RCA (or RCEM) model. That is, if \hat{b}_1 and $b_{1,3}$ are the partial correlation of x_1 in the restricted and in the true and unrestricted models, respectively, then $E[b_1]/E[b_{1,3}] \sim Bias[b_1]$ is the implicit relationship between x_1 and x_3 . Similar reasoning applies to \hat{b}_2 and $b_{2,3}$, regarding the implicit relationship between market power and trade protection. The latter effect cannot be rejected if the change in this parameter is expressive.

This identification can be improved by restricting our attention to the firm-size component in PLANTBR. That is, working with

$$FIRMSZBr_{it} = \frac{(x_{it} / N_{it})}{\bar{Z}_t} \quad (14)$$

where \bar{Z}_t , the year-average of the numerator, is a normalization factor. It can be told that $FIRMSZBr_{it}$ is less of plant-fixed costs measure than $PLANTBR_{it}$. To remedy any likely bias, we also test $DIFIRMSZ_{it} = \log(FIRMSZBr_{it}) - \log(USFIRMSZ_{it})$, giving the variation in $FIRMSZBr_{it}$ with respect to the average firm size of the free-trade reference.

Each pair of equations in Table (5) stands for a particular stepwise regression. As shown, the coefficient of the most relevant variables changed in the predicted direction: more positive for FIRMSZBr and SIZE, and more negative for CPCOST, meaning that the underlying productive and allocative ef-

efficiency effects are magnified in the absence of FPROT. We include TNOM in the pair (iii)-(iv), to better single out the wedge from price to cost, more consistently expressed by FPROT. Now the impact of deleting FPROT on FIRMSZBr is far greater than in the set CPCOST and SIZE. Finally, replacing FIRMSZBr by DIFIRMSZ, in columns (v)-(vi), we observe that the statistical significance of this new scale variable is maintained only in the restricted model, confirming differently its strong correlation with FPROT.

Table 5: Stepwise Estimates of the Comparative Efficiency Model

Independent Variables	Dependent Variable: RCEM					
	(i)	(ii)	(iii)	(iv)	(v)	(vi)
WYEL	-0.070	-0.083	-0.043	-0.062	-0.018	-0.031
	0.027	0.030	0.021	0.025	0.015	0.017*
CPCOST	0.081	-0.049	-0.089	-0.076	-0.138	-0.087
	0.025	0.035	0.046*	0.054	0.089	0.079
DCPCOST	0.846	0.620	0.862	0.627	0.738	0.510
	0.076	0.140	0.082	0.150	0.196	0.263
SIZE	0.179	0.218	0.167	0.187	0.501	0.726
	0.036	0.123*	0.053	0.146	0.095	0.168
FIRMSZBr	-0.156	0.026	-0.099	0.029		
	0.050	0.057	0.077	0.076		
DIFIRMSIZE					0.307	0.471
					0.076	0.092
CORPBR	0.422	0.350	0.364	0.396	0.467	0.516
	0.045	0.079	0.059	0.083	0.071	0.077
TNOM			-0.304	0.043	0.159	0.078
			0.028	0.042	0.025	0.037
FPROT	-0.297		0.083		-0.269	
	0.018		0.035		0.020	
N. Observations	77	77	77	77	77	77
Adjusted R^2	0.621	0.328	0.467	0.359	0.618	0.399
F statistics	25.11	12.42	13.36	11.43	21.41	12.74

Idem Table 2.

Hence, the weaker correlation of FPROT with the marginal-cost variables, as compared to the fixed-cost (markup revenues) variables, supports the idea of a pro-competitive effect underlying FPROT, besides a likely allocative effect.

Note, finally, that the three main policy effects, observed in our analysis, sum up to lower income (consumer expenditure) of Brazilian residents, which may be translated into an indirect utility func-

tion to study each welfare losses (see Feenstra, 1995).¹⁴ Apart from the theoretical disputes around such normative empirical normative, the several transformations we had to do in the variables prevent such welfare quantification.

7 Conclusions

This paper investigates the various efficiency impacts of the Brazilian ISI through comparative (export) advantages, where the spatial monopolistic competition allowed us drawing those policy effects from an industry-level analysis, whereas the enlarged foreign economy overcame several data difficulties, which are common to international comparison over a long (and distant) period.

The *productive (scale) inefficiency* was manifest in the exploratory graphical analysis of PLANTBR, as compared to the counterfactual *PLANT*, and by the weaker partial linkages of the former variable to Revealed Comparative Advantages. The allocative inefficiency was uncovered when Replacing RCA by a more accurate trade-pattern measure as to efficiency in resource allocation, RCEM, which drastically weakened the linkages from marginal costs, CPCOST and SIZE.

Lastly, evidence of the *anti-competitive effect* was attested when decomposing the negative impact of the ERP on RCA, which was strongly correlated with average cost terms. This finding supports the notion of a price-driven allocative distortion. The non-cost competition term showed that the country did not thrive in the globally most expanding industries, which suggests that the ISI failed to achieve a key dynamic target.

Expanding this comparative advantages with spatial monopolistic competition to a firm-level analysis and to the post openness period as well would be fruitful. However, the former entails a new theoretical model and the latter new procedures for controlling other policy reforms.

¹⁴The same applies to the ignored analysis of demand for varieties, inasmuch as their number do not increase with protection and that the possibility of income gains depends on either no entry or a constant markup (Helpman and Krugman, 1989).

A Data Appendix: Sources

RCA_{it} : UNCTAD, Handbook of International Trade and Development Statistics; United Nations, International Trade Statistics Yearbook; IBGE, Anuário Estatístico do Brasil. All in current US dollars.

$WYEL_{it}$: the same as RCA_{it} and also United Nations, *Commodity Trade Statistics Database*.

$CPROD_{it}$, $CPCOST_{it}$, $PLANT_{it}$, $PLANTBR_{it}$ and $SIZE_{it}$, $CORP_{it}$, $CORPBR_{it}$: UNIDO, *Industrial Statistics Database*; UN, *Yearbook of Industrial Statistics*; IBGE (idem), with values added deflated by the US and Brazil's GDP deflator, respectively. Industries average wages were based on UN, Statistical Yearbook and ILO, LABORSTA Labour Statistics Database, IBGE, Estatísticas Históricas do Século XX, and FIESP (São Paulo State Industry Federation), for Brazil in 1980. Lastly, number of firms in industries: Country Business Patterns for the USA, and IBGE Estatísticas Históricas for Brazil.

Brazil's series of employment and number of establishments was interrupted in 1985, at the beginning of the democratic government, and the level of the new series shifted dramatically, so that their values in 1987-1988 were interpolated according to Cinquetti and Silva (2010). Briefly, the employment data was interpolated from IBGE's special series for 1985-88, together with those ones of 1984 and 1988 (in Estatísticas Históricas), while the establishments, whose new data started only in 1986 (In "www.sidra.ibge.gov.br/bda/cempre/", collected in 11/03/2009), was interpolated from a mix of statistical forecast (up to 1986) and mathematical interpolation (from the yearly sample variation from 1986 to 1988). To forecast, we estimate a class of first-order autoregressive model, adding the actual and one-period lagged GDP as covariate, the specification of which was oriented by the structural shift (with high instability) of the Brazilian economy during the 1980s.

$FPROT_{it}$ and $TNOM_{it}$: Bergsman and Malan (1971); Neuhauss and Lobato (1978); Tyler (1985); Kume (1989).

B Statistical Appendix

We specified the unexplained constant term, α_i , as a group-specific constant (fixed effects, FE), in the regression models, based on evidence (see Table 2) that the parametric differences between cross-sections were associated with industry characteristics, which is bound to in a panel data with time and cross-section dimension like ours (Greene, 2000, p. 615). Another indication was the correlation with the regressors: the variance of the β s increase tremendously – most of them loose statistical significance – when running the baseline models either as RE. A Hausmann test yielded $\chi^2 = 4.79$ (p -value = 0.571), which does not reject the null hypothesis of the RE model, but this test is inadequate for small samples (Hsiao, 2003) like ours, so that we further applied the test of redundancy of the fixed effects, yielding $\chi^2 = 174.55$ (p -value = 0.000) that strongly rejects the null hypothesis of redundant FE.

Lastly, the sample size, the usual problems with international data, and cross-time heterogeneity of the sources (for FPROT) dictated a WLS-White estimator that corrects contemporaneous cross-equation correlation as well as different error variances in each cross-section (Arellano, 1987).

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