Import Price-Elasticities: Reconsidering the Evidence*

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Abstract:

Recent geography and trade empirical studies based on monopolistic competition [Hanson, 1998; Head and Ries, 1999; Hummels, 1999], suggest high levels of trade price elasticities (between 3 and 11). However, direct estimations of price-elasticities in trade equations, using price indexes at the aggregate or industry levels, lead to much lower values than those predicted by the theory (usually around unity). In this article, we show that these inconclusive results may be due to an econometric misspecification of these equations, measurement errors in import price indexes as well as endogeneity between prices and trade quantities. We re-estimate import price-elasticities from gravity-like equations using methods of transformed least squares and instrumental variables. Our study is based on compatible bilateral trade and activity data from the OECD and INSEE[1] for 14 import countries, 16 trading partners, 27 industries and 23 years. When suitable instrumental variables are used, we find relatively high price-elasticities, usually ranging from 1 to 7, the highest estimates corresponding to industries producing homogeneous goods. These results support recent studies on substitution elasticity estimates using monopolistic competition. Our results constitute a first step towards a reconciliation of the theory and the evidence.

Keywords : Gravity models, trade equations, trade price-elasticity, imperfect competition, market structure, product differentiation, unit value indexes of trade.

JEL classification: C2, C3 and F1.

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I Introduction

The new trade theory shows that elasticities of substitution and import price elasticities tend to be equal in industries producing large numbers of varieties [see Helpman and Krugman, 1985]. Assuming that this is the case, very recent empirical studies suggest significantly higher price-elasticities than those usually provided by the literature .

Namely, several articles based on original trade or geography frameworks [Head and Ries, 1999; Hummels, 1999; Hanson, 1998] or using new proxies of prices [Eaton and Kortum, 1997] obtain high values of substitution elasticities. Additional support for these results can be found in the field of industrial economics. In fact, low mark-up estimates or account rates of return are usually observed at industry levels[2], which may be consistent with relatively high levels of substitution elasticities, at least in the monopolistic competition type industries.

However, direct estimations of import price-elasticities at aggregate or industry levels do not generally support the theory since they lead to values that are hardly higher than unity. In this article, we suggest that these estimates might be biased due to some misspecification in traditional trade equations, price endogeneity and measurement errors in import prices.

Relying on a monopolistic competition framework, we re-estimate direct import price-elasticities from gravity-like equations on compatible bilateral trade and activity data (ISIC nomenclature). Data mainly originate from two sources: the OECD-STAN database and INSEE bilateral trade flow database (FLUBIL).We have built a database for 14 countries, 23 years and 27 industries (ISIC, 3-4 digits). When using OLS or fixed effect methods, our estimates show rather low import-price elasticities. However, when we both apply suitable instrumental variables for relative import prices and allow for cross fixed effects, we get price-elasticities around 3.5 on our pooled sample. We perform the same type of regression at the industry level and derive price-elasticities generally ranging from 1 to 7. In addition, price elasticity estimates appear to be significantly correlated with the degree of product differentiation. In fact, our estimated price-elasticities are higher in industries producing homogeneous products than in those producing differentiated ones. These results support those from previous studies on substitution elasticity estimates. Eventually, they are an attempt for reconciling the theory with the evidence.

In the following section, we review the existing studies that perform direct and indirect estimations of trade price elasticities at the industry level. In section III, we briefly present our theoretical model, as well as our estimation strategy. After describing the data (section IV), we present the results on the pooled sample, as well as on industry samples (section V).

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[2] See Schmalensee (1989) for reviewing profitability measures and Bresnahan (1989) for a survey on alternative methods of mark-ups estimates.

II- Literature review

As the new trade theory shows, price and substitution elasticities tend to be equal in industries producing large numbers of varieties. Assuming that this is the case, recent empirical studies find significantly higher price-elasticities than those usually provided in the literature. Using data on both freight charges and bilateral trade, Hummels [1999] estimates freight and trade equations from which he infers, though with some skepticism, a mean substitution elasticity of 7.6 over his all-industry-country sample. Similarly, Head and Ries [1999] get high substitution elasticities (around 8) from a border effect equation accounting for tariff and non-tariff barriers. Studying the links between bilateral trade and technology, Eaton and Kortum [1997] also find very high elasticities of substitution associated with relative wages (around 3.5), although smaller than those predicted in former studies. More striking, Hanson [1998] estimates a wage equation derived from the Krugman [1992] spatial model[3], and obtains substitution elasticities between 6 and 11. Moreover, as the Krugman model is based on a monopolistic competition framework, Hanson was able to infer mark-up estimates, evaluating them at 1.10-1.20.

The previous studies are generally consistent with industrial organization articles that focus on the estimation of degrees of market power. Following Hall's method [1986] that infers mark-ups from the Solow residual equation, Roeger [1995] finds mark-up rates ranging from 1.15 to 2.75 in the US industry. However, accounting for intermediary inputs in a multi country-study, Oliveira-Martins, Scarpetta and Pilat [1996, OMSP hereafter] get mark-ups between 1.20 and 1.30 in monopolistic industries[4]. If one beleives OMSP estimates, then price elasticities of demand can be directly inferred and, hence, should lie between 4 and 6.

Although all these studies seem to reconcile theory with observation, they prove to be inconsistent with most direct estimations of import price elasticities. Actually, direct estimates of the latter are seldom higher than unity, as is shown in table 1 in appendix, which reviews several traditional-type studies at industry level[5]. According to the related literature, the incompatibility between empirical results and theoretical frameworks can originate from two factors.

Firstly, endogenous links between prices and quantities may be responsible for relatively low price-elasticity estimates. In a competitive or a traditional oligopolistic setting, prices and quantities must adjust simultaneously, which leads to non-orthogonal price and residual vectors in a trade equation. Simultaneity problems can arise even if prices do not depend on quantities. In a monopolistic framework for instance, prices result from marginal costs

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[3] Hanson's result seems to be sensitive however to the considered period.

[4] These results concern all types of frameworks that produce monopolistic mark-ups such as monopolistic competition, monopoly or even cartels.

[5] The same levels apply to estimations on macro level data. See the survey of Goldstein and Khan [1985] in this respect.

inflated by mark-ups (see theoretical model in section III). If however some factors such as quality, technical progress, or any shock usually not accounted for by the theory enter simultaneously the residual component of the volume and price equations, then one will not be able to estimate consistent price-elasticities.

Typically, since quality is positively correlated with both prices and export quantities, omitting the quality factor in trade equations is likely to lead to downward biased price-elasticity estimates. Injecting unit value indexes and a quality indicator derived from survey data into a gravity-like equation, Crozet and Erkel Rousse [1999] show that one can get higher price-elasticities when controlling for quality effects. Besides, taking quality into account improves the statistical adjustment of the model. This result suggests that omitting this indicator from equation causes possible correlation between the price index and the residuals. However, in this study, the rise in price elasticities when including quality in trade equations reaches only 25% or so, which boosts the elasticities barely above unity. Unfortunately, this method therefore does not enable the authors to completely fill the gap between the (high) theoretical and (low) empirical levels of price-elasticities.

Secondly, insufficient geographical or industry disaggregation in the data might also cause low price-elasticities. In particular, one may obtain biased estimates when using unit values as proxies of real prices at an aggregate level. In fact, unit values of trade are expected to encompass most components of prices rather than focusing on one of them[6]. Hence, even if one accounts for quality in a trade equation, price elasticity estimates may still be biased if unit values are correlated with the residual vector.

Grossman [1982] tries to solve this potential problem by focusing on eleven homogeneous commodity groups chosen among several products at the 7-digit SITC nomenclature. Studying US imports from two groups of exporters, LDCs and industrial countries, Grossman specifies an import equation for the US that allows for heterogeneity between US price elasticities and those of foreign prices. He obtains relatively high price-elasticities with respect to US-produced goods (1 to 9), but lower ones for foreign imported goods (around unity). Several other authors performing estimations at more aggregate industry levels have tried to avoid geographical biases by using bilateral trade data. However, none of them gets fully convincing results concerning the level of price-elasticities (see table 1 in appendix).

Moreover, biases arising from aggregation or endogeneity problems might explain why one rarely gets satisfactory correlations between industry price-elasticities and the degree of product differentiation. In fact, some studies exhibit rather relatively high price-elasticities in highly differentiated and concentrated industries such as chemicals [*Cf.* Ioannidis and Schreyer, 1997] or motor vehicles [*Cf.* Anderton, 1998], or very low or statistically unsignificant price-elasticities in industries producing homogeneous goods, such as Rubber and Plastic products or Non-metallic products [*Cf.* Ioannidis and Schreyer, 1997 and Greenhalgh, Taylor and Wilson, 1994].

Hereafter, we present our theoretical model (section III). Then, we try to avoid the possible correlation between price indices and residuals that may arise from traditional trade modelling, using an original estimation method combining transformed least squares and instrumental variables (section IV).

III The theoretical model

Assume there are $I \geq 2$ countries, and K sectors producing differentiated goods. Any couple (i,k) represents a specific market (that of product k in country i). It is assumed that these markets are segmented.

III-1. Supply side:

Factor endowments and technologies may differ across countries. However, to simplify the specification of the model, factor markets are treated as exogenous. Positive fixed costs lead to increasing returns, so that *one firm produces only one variety of a given good.* Moreover, firms are supposed to produce within a given country, at conditions prevailing in the latter. In other words, within a given sector, they face the same production and cost functions.

More precisely, any firm located in country i and producing a variety v of product k $\in \{1,...K\}$maximises its profit function with respect to its prices (expressed in its national currency):

$$\text{Max } \Pi_{ikv} = \sum_{j=1}^{I} \Pi_{ijkv} = \sum_{j=1}^{I} (\tilde{p}_{vijk} - c_{ik}.\tau_{ijk}.t_{ij}).x_{vijk} - F_{ik}$$

Where $\lambda_l = 0$ represents the demand addressed to firm (v,i) on market (j,k) at a given price $\tilde{p}_{vijk}$, F_{ik} the amount of fixed costs, c_{ik} the marginal production cost, τ_{ijk} transport costs and t_{ijk} possible tariffs, both being expressed using an “ iceberg ” formulation. Transport costs and tariffs are assumed to depend on both sectors and trading partners, but not on the variety itself.

Let ε_{vijk} denote the elasticity of demand to prices:

[6] As noted by Grossman [1983, p.275], « the relationship between unit values (constructed at aggregate levels) and the true prices become distorted over time due to changes in the composition of the commodity bundles represented by the (unit values) indexes ».

$$\varepsilon_{vijk} = -\frac{\partial x_{vijk}}{\partial \tilde{p}_{vijk}} \cdot \frac{\tilde{p}_{vijk}}{x_{vijk}}$$

Maximising profit with respect to $\tilde{p}_{vijk}$ leads to the well-known result:

$$\tilde{p}_{vijk} = \frac{1}{1 - \frac{1}{\varepsilon_{vijk}}} . c_{iv} . t_{ij} . \tau_{ijk}$$

which can be expressed in terms of the currency of country *j*:

$$p_{vijk} = \frac{1}{1 - \frac{1}{\varepsilon_{vijk}}} . c_{iv} . t_{ij} . \tau_{ijk} . e_{ij} \qquad (1)$$

where e_{ij} represents the exchange rate of currency *i* with respect to currency *j*[7].

Firms sell their variety of product at a price that increases with total unit costs (consisting of marginal production costs, transport costs and tariffs), and whose mark-up rate is a decreasing function of the elasticity of demand to prices. Due to the fact that every firm located in country *i* faces the same production function and transaction costs, every variety of product *k* originating from country *i* is sold on market *j* at the same price and, consequently, faces the same demand on this market provided that consumer preferences do not differ from a variety (*v*,*i*) to the other.

III-2. Demand side:

Our demand side is inspired from Erkel-Rousse [1997] and is close to that of Head and Mayer [1999]. The representative consumer in country *j*, $j \in \{1, \dots, I\}$, maximises each of the *CES* sub-utility functions U_{jk} associated with the consumption of commodity *k*, $k \in \{1, \dots K\}$:

$$U_{jk} = \left[\sum_{i=1}^{I} \sum_{v=1}^{n_{ijk}} \alpha_{ijk} x_{vijk}^{\frac{\sigma_{jk}-1}{\sigma_{jk}}} \right]^{\frac{\sigma_{jk}}{\sigma_{jk}-1}}$$

where: x_{vijk} stands for the total demand for variety *v* addressed to its producer (in country *i*) on market (*j*,*k*) and n_{ijk} for the total number of varieties of commodity originating from country *i* available on market (*j*,*k*). Following Hickman and Lau [1973], geographic preference parameters $\left(\alpha_{ijk}\right)_{i=1,\dots,I}$ are normalised so that $\sum_{i=1}^{I} n_{ijk} \alpha_{ijk}^{\sigma_{jk}} = 1$. As in Erkel-Rousse [1997], those parameters can be viewed as relative national brand images. Finally, $\sigma_{jk} > 1$ is the elasticity of substitution between the different varieties of commodity *k*.

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[7] *i.e.* the number of units of currency *j* in one unit of currency *i*.

Maximising each sub-utility:

$$\text{Max} \quad U_{jk}$$

$$\text{subject to} \quad \sum_{i=1}^{I}\sum_{v=1}^{n_{ijk}} p_{vijk}\, x_{vijk} = R_{jk}\,,$$

where $\left(p_{vijk}\right)_{i,v}$ represent prices relative to quantities $\left(x_{vijk}\right)_{i,v}$, we obtain the consumer demand for variety (v,i) on market (j,k):

$$x_{vijk} = \left(\alpha_{ijk}^{\sigma_{jk}}\right)\left(\frac{p_{vijk}}{p_{jk}}\right)^{-\sigma_{jk}}\left(\frac{R_{jk}}{p_{jk}}\right) \qquad (2)$$

with $p_{jk} = \left[\sum_{i=1}^{I}\sum_{v=1}^{n_{ijk}} \alpha_{ijk}^{\sigma_{jk}}\, p_{vijk}^{1-\sigma_{jk}}\right]^{\frac{1}{1-\sigma_{jk}}}$ (= price of the composite product (j,k)).

From (2) and the budget constraint, we can derive the explicit formulation of the elasticity of demand to prices ε_{vijk} in (1):

$$\varepsilon_{vijk} = \sigma_{jk} - \frac{\sigma_{jk}-1}{n_{ijk}}.\left(\frac{p_{vijk}}{p_{jk}}\right)^{1-\sigma_{jk}} \qquad (3)$$

whose combination with (1) rigorously proves that the price of each variety (v,i) on market (j,k) does not depend on v itself. In other terms, since every variety of product k originating from country i is supposed to be equally appreciated by consumers in country j, profit maximisation in the supply side leads to equal prices $\left(p_{vijk}\right)_{v=1,\dots,n_{ijk}}$ (*i.e.* which do not depend on index v), and consequently to identical quantities $\left(x_{vijk}\right)_{v=1,\dots,n_{ijk}}$. Total demand X_{ijk} addressed to country i on market (j,k) is therefore equal to:

$$X_{ijk} = n_{ijk}\, x_{vijk} = \left(n_{ijk}\,\alpha_{ijk}^{\sigma_{jk}}\right)\left(\frac{p_{ijk}}{p_{jk}}\right)^{-\sigma_{jk}}\left(\frac{R_{jk}}{p_{jk}}\right) \qquad (4)$$

where p_{ijk} stands for the common price of varieties (v,i), $v \in \left\{1,\dots,n_{ijk}\right\}$, on market ($j$,$k$).

From (4), we can derive the logarithmic expression of the import demand for country i with respect to that for domestic products in country j, *i.e.* of the relative market share of country i with respect to that of country j on market (j,k):

$$Log\frac{X_{ijk}}{X_{jjk}} = -\sigma_{jk}.Log\left(\frac{p_{ijk}}{p_{jjk}}\right) + Log\left(\frac{n_{ijk}}{n_{jjk}}\right) + \sigma_{jk}\, Log\left(\frac{\alpha_{ijk}}{\alpha_{jjk}}\right) \qquad (5)$$

It is noteworthy that this demand function looks very much like an import demand *à la* Armington [1969] to which both a variety factor and a relative “ brand image ” factor would have been added.

Let $M_{ijk} = \dfrac{1}{1-1/\varepsilon_{ijk}}, \forall i$. Relative prices in (5) can be given by:

$$\frac{p_{ijk}}{p_{jjk}} = \frac{M_{ijk}}{M_{jjk}} . \frac{\tau_{ijk}}{\tau_{jjk}} . \frac{c_{ik}}{c_{jk}} .t_{ij} .e_{ij} \qquad (6)$$

III-3. Toward a testable trade equation:

Equation (5) has to be transformed into a testable equation. In this respect, several points have to be mentioned.

- The preference α terms are unobservable, so that the relative brand image factor will enter the perturbation of the trade equation. It is noteworthy that omitting this factor implies a risk of under-estimating elasticities σ_{jk} in highly vertically differentiated sectors, as is shown in Crozet and Erkel-Rousse [1999]. However, since we will include fixed and cross effects in our regressions, we will take at least part of this unobservable term into account.

- As for the number of varieties, we have decided to use a traditional *proxy* based on production. More precisely, we have replaced each n_{ijk} term with a smoothing of production in country *i* and sector *k* [8]. Note that clear theoretical foundations have been established for this kind of *proxy* by Krugman [1980] in a monopolistic competition context. To our knowledge, there is no theoretical evidence that production could correctly *proxy* the number of varieties in an oligopolistic situation. In such sectors, our *proxy* might well reflect other kinds of explanatory factors, such as size or even endogenous growth effects.

- Transport costs are usually considered to be a function of bilateral geographic distance such as $\tau_{ij} = d_{ij}^{\delta}$. When replacing transport costs with this function in equation (5) above, we introduce a distance variable and an associated ($\sigma_{jk} * \delta$) parameter. Most authors use the

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[8] A *proxy* based on current production would have rather represented short-term production capacity effects. Here, following Erkel-Rousse, Gaulier and Pajot [1999], we have assumed that the efforts made by firms in terms of horizontal differentiation at a given period have a progressive influence on import demand, more precisely an initially increasing and then slowly decreasing influence. We have annualised the quarterly weights used by these authors, so that we get annual weights of 0.3 (current year), 0.4 (year - 1) and 0.3 (year - 2). Note that this smoothing corresponds to that used by Magnier and Toujas-Bernate [1994]. However, the latter use *proxies* based on smoothed *R&D* and investment rather than production. Besides, the fact that our *proxy* does not depend on importing countries *j* is not a serious problem.

great circle distance indicator, to measure this variable. However, we opted for an alternative distance indicator *à la* Head and Mayer [1999]. (see description and computation of data below).

- Flubil database provides bilateral trade unit value indexes by trading partner and industry with respect to a year of reference but does not inform us on the levels of these unit values. In other words, Flubil series deal with price variation in time but not in cross-section, which causes an additional problem when one needs to estimate price-elasticity. One way of avoiding this problem is to decompose the price expression into a price-index component and a relative price component relating to the year of reference 1990:

$$\frac{p_{ij,t}}{p_{jj,t}} = \frac{p_{ij,t} / p_{ij,90}}{p_{jj,t} / p_{jj,90}} * \frac{p_{ij,90}}{p_{jj,90}} \qquad (7)$$

In addition, we assume that the marginal cost is a Cobb-Douglas function of factor costs:

$$c_{ik} = w_{ik}^{\eta_1} * r_i^{\eta_2} * m_i^{\eta_3} \qquad (8)$$

where w_{ik}, r_i and m_i stand for the factor prices of labour, capital and materials. Hereafter, we assume that capital and material prices are those that prevail in the whole economy, in contrast to wages, that may be specific to the industry. Moreover, we reasonably suppose that $\eta_1 + \eta_2 + \eta_3 = 1$.

Accounting for both, equation (8) and the transport costs function, equation (7) can now be expressed by:

$$\frac{p_{ijkt}}{p_{jjkt}} = \frac{p_{ijkt} / p_{ijk,90}}{p_{jjkt} / p_{jjk,90}} * \left(\frac{d_{ijk}^{k}}{d_{jjk}^{k}} \right)^{\delta} * \left(\frac{w_{ik,90}}{w_{jk,90}} \right)^{\eta_1} * \frac{\psi_i}{\psi_j} * \psi_{ij} \qquad (9)$$

with $\psi_h = r_{h,90}^{\lambda_2} * m_{h,90}^{\lambda_3}, \forall h \in \{i, j\}$ and $\psi_{ij} = e_{ij,90} * t_{ij,90}$. These variables are respectively specific to one or two given countries.

As we have chosen to work primarily on four dimension pooled data (*time*industry*importer*exporter*) we combine equations (5) and (9) and transform the resulted equation into an unrestricted empirical specification form:

$$Log\frac{X_{ijkt}}{X_{jjkt}} = -\sigma_{jk}.Log\left(\frac{p_{ijkt}/p_{ijk,90}}{p_{jjkt}/p_{jjk,90}}\right) - (\sigma_{jk} * \delta).Log\left(\frac{d_{ijk}}{d_{jjk}}\right) - (\sigma_{jk} * \eta_1).Log\left(\frac{w_{ik,90}}{w_{jk,90}}\right) \quad \ldots(10)$$

$$+ Log\left(\frac{Q_{ikt}}{Q_{jkt}}\right) + Trend + u_{ijkt}$$

with (u_{ijkt}) representing a vector of specific and cross fixed effects added to a residual random vector (v_{ijkt}). Hence, we express u_{ijkt} by:

$$u_{ijkt} = \lambda_i + \lambda_j + \lambda_k + \lambda_t + \lambda_{ij} + \lambda_{ik} + \lambda_{it} + \lambda_{jk} + \lambda_{jt} + \lambda_{kt} + \lambda_{ijk} + \lambda_{ijt} + \lambda_{jkt} + v_{ijkt}$$

For ease of manipulation, we shall note $LM_{ijkt} = Log\frac{X_{ijkt}}{X_{jjkt}}$, the log of the relative market share of country *i* with respect to that of country *j* on market (*j*,*k*)[9]. $LP_{ijkt} = Log\left(\frac{p_{ijkt}/p_{ijk,90}}{p_{jjkt}/p_{jjk,90}}\right)$ represents the ratio of the bilateral import price index to the price of domestic value added in country *j* also expressed in logarithm. $LQ_{ijkt} = Log\left(\frac{Q_{ijkt}}{Q_{jjkt}}\right)$ is the log ratio of the relative production smoothing expressed in constant 1990 prices in industry *k*. $LD_{ijk} = Log\left(\frac{d_{ijkt}}{d_{jjkt}}\right)$ stands for the Head and Mayer (*HM,* hereafter) log of weighted geographic distance and $LW_{ijk} = Log\left(\frac{w_{ik,90}}{w_{jk,90}}\right)$ represents the log of industry wage level in country *i* relative to that in *j* in 1990. We include a linear *TREND* variable to the regression, since imports have grown faster than production in our OECD countries during the estimation period (1972-1994).

Equation (10) provides four indications on what one can expect from the empirical results: 1/ the parameter of substitution associated with prices should exceed one. 2/ given that $\eta_1 < 1$, the wage effect should be lower than the price-effect. 3/ The parameter relative to the variety proxy should equal unity- *Cf.* Krugman [1980]. 4/ following Hummels findings (δ = 0.2), we expect the coefficient on the distance indicator to be smaller than the estimated elasticity of substitution, if however his estimation results still hold on our country and industry sample.

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[9] The domestic market share is based on the demand for domestic products computed as (production – exports).

In a properly specified model, the residual component u_{ijkt} should be defined, as noted above, as the sum of both specific and cross-fixed effects and the perturbation component of the model v_{ijkt}. However, international economists generally do not use this kind of econometric specification, since the latter includes too many individual dummies[10]. In fact, taking all these dummies into account makes people loose several degrees of freedom and may induce serious multicollinearity problems affecting the parameters of interest. Hence, restrictions are sometimes made on at least one of the specific fixed effect parameters indexed by $l \in \{i, j, k, t\}$: $\exists l \in \{i, j, k, t\}$ where $\lambda_l = 0$. However, restrictions are most often set on cross fixed effects, which are usually supposed to be null or to be accounted for by other variables such as bilateral distance, common language or regional dummies.

Nonetheless, since the rythm of openness of some economies or industries does not match with that of some others in the estimation period (1972-1994), one should expect cross time-industry and cross time-country effects to be significant. Moreover, prices may be correlated with industry or country specific technical progress, R&D or innovations over time. Finally and above all, the account for cross fixed effects must capture the preference term effects that are included in the theoretical equation (5) as well as the factors effects, the tariff barriers and the exchange rate effects relative to equation (9). In particular, λ_{ijk} and λ_{jk} should enclose the two terms $Log\alpha_{ijk}$ and $Log\alpha_{jjk}$, while λ_i, λ_j and λ_{ij} are more general effects than $Log\ \psi_i$, $Log\ \psi_j$ and $Log\ \psi_{ij}$.

We account for these specific effects by using an alternative method: the « deviation from mean exporter specification ». Hereafter, we define this method as a transformed least square method (TLS). More precisely, for a set of importing country, industry and year *{j,k,t}* we transform the fixed effects equation (10) as follows:

$$
\begin{aligned}
LM_{ijkt} - LM_{.jkt} &= -\sigma_{jk}.(LP_{ijkt} - LP_{.jkt}) + (LQ_{ijkt} - LQ_{.jkt}) - (\sigma_{jk} * \delta).(LD_{ij} - LD_{.j}) \\
&- (\sigma_{jk} * \eta_1).(LW_{ijk} - LW_{.jk}) + \lambda_i + \xi_{ijkt}
\end{aligned} \quad (11)
$$

where:

$$
\xi_{ijkt} = (\lambda_{ij} - \lambda_{.j}) + (\lambda_{ik} - \lambda_{.k}) + (\lambda_{it} - \lambda_{.t}) + (\lambda_{ijk} - \lambda_{.jk}) + (\lambda_{ikt} - \lambda_{..kt}) + (v_{ijkt} - v_{.jkt}) \quad (12)
$$

We assume that the deviation from the mean exporter of cross fixed effects, and thus ξ_{ijkt}, are randomly and normally distributed.

One of the advantages of this TLS specification is that it sweeps out all specific and cross-fixed effects that do no not depend on the export country *i*. Moreover, because our gravity-

[10] even though international economists often pool less than four dimension data.

like equation contains time invariant variables, this transformed least square specification is more appropriate for trade equations than the traditional within specification[11].

In order to appreciate the performance of the TLS specification (11), we compare its results to the more traditional equation (10). In a final stage, since we have stressed the endogeneity and measurement error problems relative to prices in trade equations, we instrument the import price index term in the TLS specification. Based on the theoretical equation (6), the instruments that we choose are the relative wage index and the relative exchange rate index, to which we add their respective lags. In a TLS specification, we express these instruments in terms of deviations from the mean exporter. Finally, exporter fixed effects are added to form a set of 17 instruments.

IV- The Data

We have built a panel of 14 importing countries $\times$ 16 trading partners $\times$ 27 industries $\times$ 23 years from the STAN (OECD) and FLUBIL (INSEE) databases. Tables 2 and 3 in Appendix give the list of the sectors and partner countries included in our analysis.

The *STAN* annual database from the OECD has provided us with the values of production, total imports and exports, as well as value added in current and constant prices from 1972 to 1994[12]. Note that the 27 elementary industries of *STAN* are aggregated *ISIC* sectors at the 3 or 4 digit levels - *Cf.* Table 2 in Appendix. *STAN* supplies data that are compatible with OECD industry surveys such as ISDB and national accounts. Actually, OECD surveys are made at a more disaggregated level, but they are not exhaustive. For instance, they usually collect information on firms of more than 20 employees. *STAN* adjusts these data with national accounts which are exhaustive but at more aggregated level. However, as for the trade with self indicator, exports exceed production in some cases for three main reasons reported from the STAN documentation[13]: 1/ Exports include re-exports; 2/ Production data are based on industrial surveys that record establishment *primary activities*. 3/ A bias is introduced by the conversion from product-based trade statistics to activity-based industry statistics for some industries. Finally we have kept only countries and industries that did not show apparent problems when calculating the trade with self indicator[14].

- Very few databases contain bilateral data in current and constant prices for a large number of countries and industries. We have used the *FLUBIL* database of the French Statistical Institute INSEE, which provides such annual series at very detailed country and product levels from 1960 to 1994. *FLUBIL* contains bilateral trade flows calculated on the basis of

[11] The traditional within specification only allows for inter-temporal variations since it deals with deviations from the mean variable across time.

[12] Price-indexes $p_{jjk} / p_{jjk,90}$ have been approximated with value added indexes.

[13] Stan Database for Industrial Analysis, *ed.* by OECD, 1998.

[14] Belgium, Denmark and Netherlands have been removed from the importer sample because their exports exceed their production in most of their industries, probably because they are big re-exporters.

several sources, among which *Series C* of the *OECD*[15]. Like the *Series C*, *FLUBIL* provides trade data for about 5,000 products classified in the *SITC* product nomenclature. We drew up conversion tables between *SITC* (product) and *ISIC* (sector) nomenclatures to get bilateral trade values and prices for the *STAN* 27 industries and 14 countries. The sum of bilateral values proved to be quasi identical to *STAN* total trade values (imports as well as exports), which is quite reassuring. Note that we have calculated imports and unit value indexes on the basis of import declarations rather than on that of export declarations. In fact, we are interested in quantifying the degree of competition between countries at the *entry* of each market, rather than at the departure of commodities from their producing countries.

We performed a number of internal and external consistency controls on our data from *STAN* and *FLUBIL* (among which macroeconomic comparisons with trade series from the *OECD Economic Outlook*), which proved to be rather satisfactory for most countries and industries[16]. However, we had to deal with a number of systematic missing data or consistency problems in some countries or sectors, that we estimated[17] or eliminated from the analysis, depending on the frequency of the problems. Tables 2 and 3 in Appendix list the set of 17 countries and 31 sectors that have finally been included into our analysis. Note that Belgium trade encompasses that of Belgium and Luxembourg, while corresponding production data are that of Belgium only. Besides, German data are relative to West Germany during the whole estimation period.

The transport cost *proxy* has been obtained from Head and Mayer (1999) for 10 European countries. We have applied the same calculation method for the rest of the countries in our sample. Following *HM* and indexing the region of exporting country i (importing country j) by $h_i\ (h_j)$, the weighted distance can be expressed as:

$$d_{ij} = \sum_{h_i \in i} \left(\sum_{h_j \in j} s_{h_i} \cdot d_{h_i h_j} \right) s_{h_i}$$

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[15] As we focus on *OECD* countries, this source is the only “ raw ” input from which the *INSEE* derives its decomposition between trade prices and flows in constant prices.

[16] Programs and tables are available upon request in SAS format.

[17] For instance, value added in constant prices was systematically missing for the only 4 digit *ISIC* sectors kept in *STAN*, namely: 3522, 3529, 3829, 3832 and 3839 (see Appendix for a literal interpretation of these sectors). We chose to estimate these missing values by applying the 4 digit structure of value added in current price to the 3 digit corresponding aggregates (352, 382 and 383) in constant prices This method implicitly assumes that prices rise in the 4-digit sectors as in the corresponding 3-digit aggregate, which is obviously a very strong approximation. As for *FLUBIL*, we had to estimate a small number of trade prices, on the basis of mirror trade flows, when there were some, or (if there was none) on that of close aggregates (total trade flows of the two trading partners in the corresponding sector, or bilateral trade flows in an close aggregated sector...). The sectors in which this sort of estimation was most often performed were, again, some 4-digit sectors: 3112, 3529, 3829 and 3839.

where $d_{h_i h_j}$ stands for the distance between the centres of regions h_i and h_j, and s_{h_i} for the population weight of region h_i in country *i*[18]. We obtained Japanese 1990 regional population data (by prefecture) from the Japanese statistics bureau and statistics center, those of US (by state) from the US Census Bureau and those of Canada (by province) from Statistics Canada[19]. Regional population are not available for Sweden, Austria, Norway and Finland. Concerning Sweden and Austria, we used the 1990 population data of their main cities that we classed into group of cities geographically close from one another (above 150 miles), each group of cities was treated as a region. Norway and Finland have been considered to be sufficiently small countries with respect to the other countries of the sample to be represented respectively by their main cities.

V The results

V-1. Pooled estimations

Table 6 in Appendix presents alternative estimation methods for the trade equation on pooled data. Great circle distance was chosen to *proxy* trade costs in the first two equations in order to compare with the *HM* relative weighted-distance, alternatively included in the rest of the equations.

The first OLS equation *(1.a)* is similar to most gravity equations that can be found in the literature in the sense that it includes regional free trade agreement dummies (EU, NAFTA) without accounting for fixed effects. Although the estimated coefficients of these dummies have a positive sign, Matyaz (1998) shows that regional dummies may not express what they are expected to, since they are linear combinations of fixed effects. Moreover as Matyaz suggests, omitting fixed effects from a gravity equation may bias the estimates. In fact, when comparing our OLS estimation (equation *1.a)* with the fixed effects equation (*1.b)*, we find significantly different results for most of the parameters of interest[20]. Note however, that the coefficient on the intercept, possibly interpreted as the border effect in other similar studies, must not be qualified as such in our equations *(1.a)* and *(1.b).* Actually, the intercept is very sensitive to the choice of the distance parameter as well as to the introduction of the fixed effect parameters. When the distance variable does not take into account the country internal distance it biases automatically upward the coefficient on the intercept.

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[18] Head and Mayer used industry-level employment for origin weights and GDP for destination weights. As we were not provided by these kind of data we used the population weights.
[19] All these statistic sources provide data on line.
[20] This evidence holds as well when we replace the traditional distance indicator by the HM-distance.

Replacing traditional distance with the *HM* weighted distance improves the distance effect on trade, thus increasing the associated elasticity from 1.2 to 1.6 (equation *1.c).* The only estimates that are affected by the change of the distance indicator are the intercept and the fixed effects[21]. However, in the previous equations the distance effect does not confirm our expectations, since it appears to be higher than the price effect. In particular, price-elasticities in the two alternative equations *(1.b)* and *(1.c)* hardly reach 0.85[22]. On the contrary, the coefficient on the relative wage indicator reaches 0.25 which is compatible with the theory. Nevertheless, the wage effect might capture a quality or productivity effect that is not taken into account by the theory.

When comparing the traditional fixed effect specification with that of the transformed least squares based on equation (11), we find rather different estimates for the parameters. Hence, equation (*2a)* shows a price-elasticity above unity (1.15) but still smaller than that of the distance. In addition, the production and wage parameters are higher than those estimated using the prior specifications. Although theory predicts a unity elasticity, the production effect is however smaller than that estimated by Harrigan [1996] which reaches 1.20[23].

Finally, we perform an instrumental variable specification based on the transformed least squares model by instrumenting prices. In order to verify whether it is consistent or not to instrument the unit value index, we have run a Durbin-Hu-Hausman (DWH) test. The latter rejects the null hypothesis (*i.e.* the exogeneity of this indicator)[24]. We obtain a price-elasticity estimate close to 3.7 - 3.8 (see equation *2c*). Note that the other coefficients are unchanged with respect to those relative to the simple TLS method (equation *2b*). Here, the coefficient on the distance is no longer higher than the elasticity of substitution. An estimate of the elasticity of distance to transport costs can be inferred: $\hat{\delta} = 1.61/3.75 = 0.43$. The main difference between our method and that of Hummels is that he estimates δ from a direct freight equation and then infers the level of the elasticity of substitution from a gravity equation. Instead, we estimate the elasticity of substitution and that of distance simultaneously.

V-2. Industry level estimations

In the prior sub-section, we have performed estimations on pooled data, assuming that price-elasticity, as well as production and distance elasticities, are homogeneous across industries. Here, we relax this hypothesis and hence, estimate the same kind of equations on each industry individually. Following the theory, price-elasticity levels should depend on the

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[21] The fixed effect parameters are not shown in the table, but are available upon request. Moreover, the intercept appears with the same sign although taking a smaller value than the one relative to Head and Mayer's result.

[22] This result is however similar to or roughly smaller than those provided in most traditional empirical work. See the survey of Goldstein and Khan [1985] for measures of price-elasticities at the macro level and table 1 for estimates at the industry level.

[23] As is the case in this article, Harrigan tests a bilateral trade equation on OECD countries based on a monopolistic framework.

[24] For a clear exposition of this test, see Davidson and Mc. Keenon [1993], p.237-239.

degree of both product differentiation and industry fragmentation (see for exemple Krugman, 1979). However, since the fragmentation effect is controlled by the variety *proxy*, we only examine the extent to which the sensitivity to prices is related to the degree of differentiation in the commodities produced by each industry.

Table 7 in appendix presents results relative to trade price-elasticity estimates for each industry of our sample[25]. First, it should be noted that the estimates of price-elasticities at the industry level using the traditional fixed effect method are similar to those given in the literature. They are relatively low. In fact, 14 out of 27 industries are associated with price-elasticities roughly higher than one, with a maximum value for the Paper Industry, Iron and Steel, Non-ferrous metals and Motor Vehicles reaching 1.2.

Price-elasticities that we derive from our TLS estimates are a little higher than those resulting from the traditional estimations in 22 industries. This result, similar to that obtained from pooled estimation, suggests that cross-fixed effects have to be controlled for when studying the sensitivity of bilateral trade to prices. Moreover, the latter results are consistent with the assumption that brand images effects represent a part of cross specific effects.

Finally we perform estimations based on the combined TLS-I.V specification, with prices instrumented in the same way as in the equivalent specification on pooled data. In order to obtain robust estimates, we check whether our usual instruments remained good ones for prices at the industry level. In this respect, two conditions has to be met. These instruments have to be both correlated with prices and independent from the residuals. In addition, we check the necessity of instrumenting the price indicator by running further DWH tests. Seventeen industries pass this tests, most of them known as homogenous good industries (see table7). Actually, the available instrumental variables are not really adapted to prices in differentiated product industries mainly because wages and exchange rates usually reflect a smaller proportion of the price in these industries, more intensive in capital.

Price-elasticity estimates are found to be significantly higher than those resulting from the two prior specifications, except for 5 industries, three of which presenting non-significant estimates: Paper products, Machinery and equipments and Railroad industries. Actually, in these industries, the chosen instruments are not highly correlated to prices (R-squared below 0.05), which explains their poor performance.

As for the remaining industries, the price-elasticity levels that we get seem to match the prediction of the theory. To prove this result, we compare our price-elasticity levels with the degree of product differentiation in each industry provided by two alternative classifications. The first one is derived from Rauch [1996] calculations (see Table 4). The second

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[25] For ease of discussion, we just present the parameter estimates associated with relative prices, since they are our primer interest. Thorough results for each of the presented specifications are available upon request from the authors. Note that the 1990 relative wage vector has been removed from the industry regression as it showed multicollinearity with the fixed effects in the regressions. This is not surprising since this indicator is industry and country specific.

classification is due to OMSP [1996][26]. Table 7 shows that the industries producing relatively low differentiated goods in both classifications, such as Textiles, Wood, Furniture, Rubber, Iron and Steel, Non-metallic products, and Pottery are associated with high price-elasticities (roughly 3.5 to 6.5). In addition, when the instrumental variable method is appropriate, and provided that our instruments are sufficiently correlated to prices, highly differentiated good industries such as Motor Vehicles or Other Chemicals, show price-elasticities around 3.5 to 4.

VI Conclusion

In this article, we showed that direct estimates of price elasticities can be reconciled with both elasticities of substitution estimates and theoretical predictions. Hence, once they are derived from proper econometric specifications, and when one controls for price measurement errors and endogeneity, these estimates are found to be much higher than those found in traditional empirical work. We show that the price elasticity reaches 3.7 over the pooled sample, and ranges from 1 to 7 when estimations are performed at the industry level. Moreover, unlike differentiated good industries, homogeneous good ones are associated with high price elasticities, which corroborates the theory.

Do these findings necessarily imply that trade policies, at least in terms of tariffs barriers, are more effective than it is usually assumed? Put differently, is protection really profitable for the domestic country? Actually, our estimates are based on a monopolistic behaviour framework as each representative firm in an exporting country benefits from a rent due to the specificity of its exported variety. Therefore, an increase in tariffs might only reduce domestic producers' relative market share, without necessarily affecting the level of their production. Hence, if one believes our theoretical framework, then the resulting high price elasticities suggest that a high level of protection, especially on homogeneous products, reduces consumers welfare and that the induced tariff revenues might not be as profitable as expected.

[26] The Oliveira-Martins-Scarpetta and Pilat [1996] classification is inspired from that of Oliveira-Martins [1994]. See table 5 in appendix.

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Appendix

Tables:

Table 1: Previous papers that estimate price elasticities at the industry level

Authors	Level of aggregation	Period	Export ing Countries	Importing Countries	Trade Flows	Type of equation	Import Price indicator	Level of desaggregation	Price-elasticities levels
Grossman [1982]	11 'homogeneous' commodity groups selected from 7-digit SITC data	1968-1978 (quarterly)	Less Developed or Industrial Countries	USA	Multilateral	Import equation with cross-price elasticities	multilateral unit values	by group of commodities	US price elasticities: 1 to 9 ; Non-US price-elasticities take more usual values.
Marquez & McNeilly [1988]	3 commodity groups: Food, Raw Materials and Manufactures	1973-1984 (quarterly)	Less Developed Countries	Canada, Germany, Japan, UK, US	Bilateral	Bilateral import equation	multilateral import prices	by country and industry	More (less) than unity for manufactures (food and raw materials)
Bergstrand [1989]	1-Digit SITC data	1965, 1966, 1967	16 OECD countries and Switzeland	16 OECD countries and Switzeland	Bilateral	Gravity equation model	aggregate wholesales price index for importers and exporters	by industry	Large range of coefficients (from 0.1 to 11). Most parameters are statistically insignificant
Greenhalgh, Taylor and Wilson [1994]	36 industries, Cambridge Econometric database (CE)	1954-1985 (annual)	UK	Industrial and part of LDCs	Multilateral	Import share equation	aggregate import price index	by industry	Between 0.0 and 2.5
Ioannidis & Schreyer [1997]	2-digit ISIC data	1975-1994 (annual)	10 exporting OECD countries	Industrial and part of LDCs	Bilateral	Mean bilateral export share equation	mean import bilateral prices	by industry	Between 0.0 and 1.8
Anderton [1998]	2-digit ISIC data	1970-1987 (annual)	UK and Germany	Industrial and part of LDCs	Bilateral	Bilateral import equation	bilateral import prices	by industry and importing country	UK: around unity; Germany: less than unity
Head & Mayer [1999]	2-digit Eurostat database	1986-1995 (annual)	12 EC countries	12 EC countries	Bilateral	Gravity equation model	price index at industry level	by industry	average price-elasticity around unity.
Crozet & Erkel-Rousse [1999]	2 categories : consumer goods and *other* goods	1994-1997 (annual)	4 EC countries	4 EC countries	Bilateral	Gravity-like eq. model, including a quality *proxy*	bilateral unit values	by group of commodities (consumer or *other*)	average price-elasticity above unity.
This study	3-4 digit ISIC data	1972-1994 (annual)	17 or 12 OECD countries (depending on specification)	17 or 12 OECD countries (depending on specification)	Bilateral	Gravity-like eq. model	bilateral unit values	pooled and by industry	Between 1 and 7, depending on degree of differentiation of goods produced in the industry

Table 2: Sectors of *STAN* included in the analysis

ISIC	*Description*	ISIC	*Description*
3112	Food	361	Pottery and China
313	Beverages	362	Glass and products
321	Textiles	369	Non-metallic products, nec.
322	Wearing Apparel	371	Iron and Steel
323	Leather and Products	372	Non-ferrous metals
324	Footwear	381	Metal products
331	Wood products	3829	Machinery and equipment, nec.
332	Furniture and fixtures	3832	Radio, TV and communication equip.
341	Paper Products	3839	Electrical Apparatus
342	Printing and Publishing	3842	Railroad equipment
351	Industrial Chemicals	3843	Motor Vehicles
3522	Drugs and Medicines	39	Other manufacturing
3529	Chemical products, nec.		
355	Rubber products		
356	Plastic products, nec.		

Table 3: Importing and exporting countries included in the analysis

17 Exporting countries	*14 Importing Countries*	*Mnemonic*
Japan	Japan	JPN
United States	United States	USA
Canada	Canada	CAN
France	France	FRA
Germany	Germany	DEU
Italy	Italy	ITA
Spain	Spain	ESP
Portugal	Portugal	PRT
Norway	Norway	NOR
Finland	Finland	FIN
The Netherlands		NLD
United Kingdom	United Kingdom	GBR
Belgium		BEL
Austria	Austria	AUT
Denmark		DNK
Sweden	Sweden	SWE
Greece		GRC

Table 4: Classification of *STAN* sectors derived from Rauch's calculations [1996]

ISIC	Description of the sector	Share of industry producing homogeneous goods (Rauch's calculations)	Classification based on Rauch's calculations
3112	Food	0.9133	HOM
313	Beverages	0.5394	HOM
321	Textiles	0.2639	DIF
322	Wearing Apparel	0	DIF
323	Leather and Products	0	DIF
324	Footwear	0.023	DIF
331	Wood products	0.492	HOM
332	Furniture and fixtures	0	DIF
341	Paper Products	0.5079	HOM
342	Printing and Publishing	0	DIF
351	Industrial Chemicals	0.5348	HOM
3522	Drugs and Medicines	0.050	DIF
3529	Chemical products, nec.	0.1164	DIF
355	Rubber products	0	DIF
356	Plastic products, nec.	0	DIF
361	Pottery and China	0	DIF
362	Glass and products	0.0792	DIF
369	Non-metallic products, nec.	0.5403	HOM
371	Iron and Steel	0.4729	HOM
372	Non-ferrous metals	0.6583	HOM
381	Metal products	0.1540	DIF
3825	Office and computing equip. nec.	0	DIF
3829	Machinery and equipment, nec.	0	DIF
3832	Radio, TV and communication equip.	0.0458	DIF
3839	Electrical Apparatus	0.012	DIF
3841	Shipbuilding and repairing	0	DIF
3842	Railroad equipment	0	DIF
3843	Motor Vehicles	0.0056	DIF
3844	Motorcycles and bicycles	0	DIF
385	Professional goods	0	DIF
39	Other manufacturing	0	DIF

Table 5: The Oliveira-Martins-Scarpetta and Pilat (1996) classification of *STAN* sectors

Degree of product Differentiation	*Market structure in terms of number of firms* Fragmented (high number of firms)	Concentrated (low number of firms)
Low (Homogeneous products)	Food Textiles Wearing Apparel Leather and Products Footwear Wood products Furniture and Fixture Printing and publishing Plastic products, nec Non-metallic products Metal products	Beverages Tobacco Paper products Rubber products Pottery and china Glass and products Iron and Steel Non-ferrous metals Shipbuilding and repairing
High (Differentiated products)	Chemical products, nec Machinery and equipment Motorcycles and bicycles Professional goods	Industrial chemicals Drugs and medicines Office and computing equip. Radio, TV and communication Electrical apparatus Railroad equipment Motor vehicles

Table 6: Bilateral Trade Equation: All-Industry-Country Sample

	Equation 1.a		Equation 1.b		Equation 1.c		Equation 2.a		Equation 2.b	
Method	OLS		Fixed Effects		Fixed Effects		Transformed Least Sqaures TLS (a)		I.V on TLS	
Intercept	1.696	***	3.05	***	-2.045	***	_		_	
	(0.048)		(0.056)		(0.026)					
TREND	0.044	***	0.054	***	0.054	***	_		_	
	(0.001)		(0.001)		(0.001)					
EU	0.318	***	_		_		_		_	
	(0.015)									
NAFTA	1.273	***	_		_		_		_	
	(0.090)									
Rel. Great Circle Distance	-1.047	***	-1.228	***	_		_		_	
	(0.0006)		(0.007)							
Rel. Weighted Distance					-1.586	***	-1.595	***	-1.611	***
					(0.009)		(0.009)		(0.010)	
Rel. Production	0.703	***	1.014	***	1.014	***	1.128	***	1.153	***
	(0.002)		(0.005)		(0.005)		(0.007)		(0.008)	
Rel. Prices	**-1.202**	***	**-0.841**	***	**-0.844**	***	**-1.14**	***	**-3.753**	***
	(0.013)		(0.0010)		(0.010)		(0.014)		(0.229)	
Wage90	-0.248	***	-0.256	***	-0.256	***	-0.351	***	-0.336	***
	(0.010)		(0.016)		(0.016)		(0.024)		(0.0028)	
Exporter Fixed Effects	No		Yes		Yes		Yes		Yes	
Importer Fixed Effects	No		Yes		Yes		Implicit		Implicit	
Industry Fixed Effects	No		Yes		Yes		Implicit		Implicit	
Cross fixed effects	No		No		No		Implicit		Implicit	
Number of countries	14		14		14		14		14	
R2	0,521		0,718		0,726		0,616		0.623	
Nb. of observations	130190		130190		130190		130190		130190	
Period	1972-1994		1972-1994		1972-1994		1972-1994		1972-1994	
DWH test									244,933	***

***Significant at the 1% level
(a) deviation from mean exporter for a given year , industry and import country
values between brackets express the standard error of the estimates.

Table 7: Price-elasticities derived from bilateral trade equations, by industry sample

Label	PDT_RA	PDT_OM	F.E		TLS		IV		DWH tests predicting consistency
Manuf, nec.	DIF	DIF	-0.872	***	-0.985	***	**-1.117**		**yes**
s.e			*0.029*		*0.032*		*0.678*		
Beverages	HOM	HOM	-0.776	***	**-0.896**	***	-1.703	***	no
s.e			*0.049*		*0.069*		*0.476*		
Textiles	DIF	HOM	-1.134	***	-1.239	***	**-4.253**	***	**yes**
s.e			*0.062*		*0.082*		*0.454*		
Apparel	DIF	HOM	-0.956	***	-0.85	***	**2.115**	**	**yes**
s.e			*0.052*		*0.072*		*0.842*		
Leather	DIF	HOM	-0.967	***	**-1.116**	***	-0.821		no
s.e			*0.042*		*0.053*		*0.796*		
Footwear	DIF	HOM	-1.007	***	**-0.625**	***	-2.364	***	no
s.e			*0.058*		*0.092*		*0.895*		
Wood	HOM	HOM	-0.943	***	-0.898	***	**-3.129**	***	**yes**
s.e			*0.047*		*0.064*		*0.735*		
Furniture	DIF	HOM	-1.114	***	-1.227	***	**-3.898**	***	**yes**
s.e			*0.036*		*0.056*		*0.429*		
Paper	HOM	HOM	-1.243	***	**-1.518**	***	-0.099		**yes**
s.e			*0.063*		*0.088*		*0.736*		
Print/Publish.	DIF	HOM	-1.055	***	-1.194	***	**-1.462**	***	**yes**
s.e			*0.04*		*0.051*		*0.464*		
Chemicals	HOM	DIF	-1.085	***	**-1.315**	***	-0.859	***	no
s.e			*0.038*		*0.056*		*0.312*		
Rubber	DIF	HOM	-0.984	***	-0.891	***	**-6.482**	***	**yes**
s.e			*0.054*		*0.084*		*1.282*		
Plastic	DIF	HOM	-0.815	***	-0.989	***	**-1.448**	***	**yes**
s.e			*0.037*		*0.047*		*0.312*		
Pottery/China	DIF	HOM	-0.764	***	-0.854	***	**-3.782**	***	**yes**
s.e			*0.041*		*0.052*		*0.543*		
Glass	DIF	HOM	-1.033	***	**-1.035**	***	-1.056	**	no
s.e			*0.043*		*0.056*		*0.52*		
Non-metallic	HOM	HOM	-1,000	***	-1.047	***	**-6.619**	***	**yes**
s.e			*0.044*		*0.053*		*0.743*		
Iron/Steel	HOM	DIF	-1.245	***	-1.356	***	**-3.225**	***	**yes**
s.e			*0.055*		*0.075*		*1.032*		
Non-ferrous	HOM	HOM	-1.226	***	**-1.521**	***	-0.828		no
s.e			*0.055*		*0.084*		*1.118*		
Metal	DIF	HOM	-0.924	***	**-1.098**	***	-1.444	***	no
s.e			*0.047*		*0.06*		*0.286*		
Food	HOM	HOM	-1.064	***	**-1.195**	***	-0.95	**	no
s.e			*0.036*		*0.043*		*0.398*		
Drugs/Med.	DIF	DIF	-0.981	***	**-1.002**	***	-2.018	**	no
s.e			*0.03*		*0.037*		*1.015*		
Chemical, nec.	DIF	DIF	-1.13	***	-1.265	***	**-4.163**	***	**yes**
s.e			*0.051*		*0.058*		*1.337*		
Machin/Equip.	DIF	DIF	-0.803	***	**-1.291**	***	1.079		**yes**
s.e			*0.04*		*0.057*		*0.596*		
Radio, TV,Tel	DIF	DIF	-1.096	***	**-1.192**	***	-0.484		no
s.e			*0.039*		*0.049*		*0.649*		
Electric	DIF	DIF	-0.776	***	-0.9	***	**3.063**	***	**yes**
s.e			*0.039*		*0.045*		*1.138*		
Railroad	DIF	DIF	-0.794	***	**-0.921**	***	3.689		**yes**
s.e			*0.04*		*0.062*		*4.282*		
Vehicles	DIF	DIF	-1.201	***	-1.562	***	**-3.32**	***	**yes**
s.e			*0.058*		*0.077*		*0.68*		

Notes: 1/ PDT_OM and PDT_RA refer respectively to Oliveira Martins and Rauch's adapted classifications of industries producing relatively differentiated (DIF) or Homogeneous (HOM) products.

2/ *** significant at 1%; ** at 5%, * at 10%