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Trade performances, product quality perceptions and the estimation of trade price-elasticities [+]

*Matthieu Crozet and Hélène Erkel-Rousse**
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LRH: Matthieu Crozet and Hélène Erkel-Rousse

Abstract: Traditional trade models ignoring the dimension of product quality generally lead to excessively low trade price elasticities. In this paper, we estimate import market-share equations including a quality image proxy derived from survey data. Our estimation results, based on panel data for the four main EU member States, confirm the part played by product quality perceptions in the estimation of trade price elasticities, at least for highly differentiated products. Introducing the quality image proxy into the models leads to a significant increase in the price elasticities, which thus become superior to unity, i.e. in conformity with theoretical elasticities of substitution.

* Crozet: TEAM – CNRS, Paris I Panthéon-Sorbonne University, 106-112 Boulevard de l'Hôpital, 75657 Paris Cedex 13, France; Tel/Fax: (33) 1 44 07 82 67. E-mail: crozet@univ-paris1.fr. Erkel-Rousse: TEAM, CNRS, Paris I Panthéon-Sorbonne University, Paris, France and Direction de la Prévision (Forecasting Directorate), French Ministry of the Economy, Finance and Industry at the moment when the study was performed. E-mail : helene@erkel.easynetbox.net. We thank, without implication, the French Ministry of the Economy, Finance and Industry (Direction de la Prévision - Forecasting Directorate), which financed this study, as well as the COE-CCIP, and more especially Françoise Précicaud, for having kindly provided us with data from the « Image of European Products » survey. We are also grateful to an anonymous referee for his or her fruitful remarks and suggestions on the paper.

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Abbreviations:
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Address of Contact Author: Matthieu Crozet, TEAM – CNRS, Paris I Panthéon-Sorbonne University, 106-112 Boulevard de l'Hôpital, 75657 Paris Cedex 13, France; Tel/Fax: (33) 1 44 07 82 67. E-mail: crozet@univ-paris1.fr.

1 Introduction

Most operational trade models do not take into account the "new" theory of trade and stick to the traditional Armington (1969) framework. Such trade equations, however, often suffer from serious estimation difficulties such as excessively low or unstable trade price elasticities, notably suggesting specification problems (Orcutt, 1950; Harberger, 1953; Goldstein and Khan, 1985; Madsen, 1999; Deyak, Sawyer, and Sprinkle, 1997). More generally, apart from few recent valuable contributions (Hummels, 2000; Eaton and Kortum, 2002; Head and Ries, 2001; Baier and Bergstrand, 2001; Clausing, 2001; Erkel-Rousse and Mirza, 2002), many estimations of trade price elasticities lead to relatively low values in the literature.

This problem is crucial from both a theoretical and empirical point of view. Firstly, the "new" trade theory shows that elasticities of substitution and import price elasticities are superior to unity and tend to be equal in industries producing large numbers of varieties (Helpman and Krugman, 1985). Secondly, several authors such as Cox and Harris (1985), Brown (1987) or Shiells and Reinert (1993) point out that the values of trade price elasticities have a crucial incidence on the quantitative and qualitative results of analyses performed on the basis of multinational models. In particular, these values condition both welfare effects of trade and the consequences of exchange rate policies, as well as the macroeconomic regulation of open economies in general.

Traditionally, this problem has been addressed by stressing the crucial part of both imperfect measurement of trade prices and potential endogeneity problems on trade price elasticity estimation[1]. In this paper, we suggest that these unquestionable causes for low price elasticities might not be exclusive. More precisely, we show that ignoring vertical product differentiation in trade equations also leads to under-estimated price elasticities. The underlying intuition is easy to understand. In fact, adding a quality variable into a market share equation would enable one to suppress the (positive) indirect effect of product quality through prices from the (negative) overall relative price effect. The relative price contribution would thus become a pure price effect, which has an unambiguous negative impact on market shares, while the overall positive influence of product quality would be captured by the quality factor.

In practice, the quantitative impact of ignoring product quality might be relatively high in vertically differentiated industries (Shaked and Sutton, 1984; Falvey and Kierzkowski, 1987), in such a way that it be significant at macroeconomic level. In this respect, an empirical study performed by Fontagné, Freudenberg and Péridy (1998) confirms the increasing share of trade in vertically differentiated products in total trade, especially within the European Union (EU).

Unfortunately, product quality is usually unobservable, so that introducing such a variable into trade equations requires the definition of a proxy. Many authors use proxies based on R&D expenses or human capital variables (Greenhalgh, Taylor, and Wilson, 1994; Eaton and Kortum, 2002; Ioannidis and Schreyer, 1997; also see other references below). Such indirect measures of product quality may a priori differ from what they are supposed to capture or, at least, focus on a specific dimension of product quality, namely technological differentiation. As far as the present study is concerned, we have had the opportunity to use a direct measure of quality perceptions or « images » relating to products originating from the four main EU countries and derived from a survey performed by the Centre d'Observation Economique (COE) of the Chambre de Commerce et d'Industrie de Paris on a European basis. Our estimations lead to results in conformity with intuition. Controlling a market share equation for product quality perceptions enables us to obtain a significant increase in trade price elasticity; the price elasticity thus becomes superior to unity. We obtain this result without changing our relative price index. In this respect our approach differs from that of Feenstra (1994), who modifies his trade price index to take product variety into account. We then replace our quality proxy with an innovation perception proxy derived from the same survey, for comparison purpose. Results of models using the former or the latter proxies prove to be very much alike. This comparison provides one with a significant even though fragile ex post argument in favour of using indirect quality image proxies based on innovation variables as a second best solution.

2. The theoretical model

In this section we present a trade model derived from Erkel-Rousse (1997, 2002), which leads to the determination of a testable market-share equation.

Assume there are $I \geq 2$ trading countries, producing and exchanging K differentiated products k = 1,...,K. The representative consumer of country j, $j \in \{1,...,I\}$, maximises a Spence-Dixit-Stiglitz sub-utility function U_{kj} subject to his or her budget constraint:

$$U_{kj} = \left[\sum_{i=1}^{I} \sum_{v=1}^{n_{ki}} \alpha_{kij} y_{vij}^{(\sigma-1)/\sigma} \right]^{\sigma/(\sigma-1)} \tag{1}$$

where y_{vij} stands for total demand of variety v addressed to its producer in country i and n_{ki} for the number of varieties originated from country i. σ is the elasticity of substitution between domestic and imported goods from different origins. Preference parameters $(\alpha_{kij})_{i=1,\dots,I}$ can be viewed as the quality perceptions or images of product varieties originating from country i. They depend on the intrinsic quality of varieties produced in country i, but also on the representative consumer in country j's purely subjective perceptions (as regards product desirability), which may differ from one importing country to another. Note that our reading of preference weights therefore encompasses the alternative interpretations given by Feenstra (1994) (in terms of product quality) and Head and Mayer (2000) (in reference to home bias or typical aversion to foreign products).

Preference parameters are equal for all varieties of product k originating from the same country i. This property stems from the fact that preference parameters are supposed to essentially derive from national differences in terms of production technologies and know-how. Conversely, we assume that firms from a given country face the same production conditions.

Total production of variety (v,i) is broken down between markets (k,j), $j = 1,...,I$:

$$y_{vi} = \sum_{j=1}^{I}(1+t_{kij})y_{vij}$$

where y_{vij} stands for the production share of variety (v,i) sold on market (k,j), which is identical to the demand expressed on this market at equilibrium. The combination of transport and transaction costs is viewed as the destruction of a part ($t_{kij}y_{vij}$) of the production shipped towards market (k,j) during the transportation from country i to country j ("iceberg" representation).

Production and transport conditions being identical for every variety (v,i), the latter are sold in equal quantities $y_{vij} \equiv y_{kij}$ and at the same price $p_{vij} \equiv p_{kij}$ on market (k,j). From the producer profit maximisation, one derives the well known price expression:

$$p_{kij} = c_{ki}(1+t_{ij})\,\varepsilon_{kij}\big/(\varepsilon_{kij}-1) \tag{2}$$

where c_{ki} denotes the unit cost of producing a variety of product k in country i and $\varepsilon_{kij} \equiv -\left(\partial y_{kij} / y_{kij}\right)\big/\left(\partial p_{kij} / p_{kij}\right)$ the price elasticity of demand for variety (v,i) in country j. The expression of ε_{kij} is calculated on the basis of demand functions. The latter are derived from the first-order conditions associated with maximisation of (1) under the consumer budget constraint:

$$y_{kij} = \left(p_{kij}\big/p_{kj}\right)^{-\sigma}\left(\alpha_{kij}^{\sigma}\Big/\sum_{i'=1}^{I} n_{ki'}\alpha_{ki'j}^{\sigma}\right)\left(E_{kj}\big/p_{kj}\right) \tag{3}$$

where, as is shown in Hickman and Lau (1973), $p_{kj} = \left[\sum_{i=1}^{I} n_{ki}\alpha_{kij}^{\sigma}p_{kij}^{1-\sigma}\Big/\sum_{i=1}^{I} n_{ki}\alpha_{kij}^{\sigma}\right]^{1/(1-\sigma)}$ stands for the average price of product k on market (k,j) and E_{kj} for the share of country j's national revenue allocated to the consumption of product k. The higher the elasticity of substitution σ, the more sensitive demands to relative prices and quality images. Partial derivations of (3) lead to:

$$\varepsilon_{kij} = 1 + (\sigma - 1)\left(1 - \alpha_{kij}^{\sigma} p_{kij}^{1-\sigma} \Big/ \sum_{i'=1}^{I} n_{ki'} \alpha_{ki'j}^{\sigma} p_{ki'j}^{1-\sigma}\right) \quad (4)$$

As (2) expresses, obtaining positive and finite price-cost margin ratios supposes that $\varepsilon_{kij} > 1$. Due to (4), this inequality is equivalent to condition $\sigma > 1$. Moreover, (4) implies that ε_{kij} tends towards σ when the number of firms tends towards infinity (case of monopolistic competition with atomistic markets). By combining (2) and (4), it can also be shown that prices (p_{kij}) are increasing functions of quality perceptions (α_{kij})[2] and decreasing functions of the number of varieties (n_{ki}) ceteris paribus. The former result is due to a differentiation effect, the latter to a pro-competitive effect, as in Krugman (1979).

Besides, in the case when firms can produce several varieties provided that they accept some increase in their fixed costs, it can be shown that the optimal number of varieties is an increasing relationship of production at the firm level, even if the number of firms is low (Erkel-Rousse, 2002). This result suggests the definition of a proxy of the number of varieties based on general theoretical foundations (in terms of number of firms) for empirical work.

Country *j*'s bilateral imports of product *k* originating from country *i* derives from (3):

$$M_{kij} = n_{ki} p_{kij} y_{kij} = \left(p_{kij} / p_{kj}\right)^{1-\sigma} \left(n_{ki} \alpha_{kij}^{\sigma} \Big/ \sum_{i'=1}^{I} n_{ki'} \alpha_{ki'j}^{\sigma}\right) E_{kj}$$

which, expressed with respect to country *j*'s imports from a competitor *i'*, leads to:

$$M_{kij} / M_{ki'j} = \left(p_{kij} / p_{ki'j}\right)^{1-\sigma} \left(n_{ki} / n_{ki'}\right) \left(\alpha_{kij} / \alpha_{ki'j}\right)^{\sigma} \quad (5)$$

The value of country *i*'s market share in country *j* with respect to that of a set of trading competitors $I' \subset \{1,...,I\} \setminus \{i, j\}$ is derived from (5) by shifting to growth rates:

$$\dot{M}_{kij} - \overset{\cdot}{\overbrace{\sum_{i' \in I'} M_{ki'j}}} = \sum_{i' \in I'} a_{ki'j} \left(\dot{M}_{kij} - \dot{M}_{ki'j}\right)$$

$$\text{with } a_{ki'j} = M_{ki'j} \Big/ \sum_{l \in I'} M_{klj} \quad , \quad \sum_{i' \in I'} a_{ki'j} = 1 ,$$

where time index t is implicit (as is the case in the whole section). Note that, if reasoning in discrete time, we would have to replace the $(a_{ki'j})$ coefficients with their lagged values.

Transforming (5) into growth rates and then integrating (which introduces an invariant factor $c_{kijI'}$), we obtain:

$$Log\left(mshare_{kijI'}\right) = -(\sigma - 1)Log\left(price_{kijI'}\right) + Log\left(variety_{kijI'}\right) + \sigma Log\left(imageq_{kijI'}\right) + c_{kijI'} \qquad (6)$$

where: $mshare_{kijI'} = M_{kij} \Big/ \sum_{i' \in I'} M_{ki'j}$, $price_{kijI'} = p_{kij} / \overline{p}_{kI'j}$, $variety_{kijI'} = n_{ki} / \overline{n}_{kI'}$, and $imageq_{kijI'} = \alpha_{kij} / \overline{\alpha}_{kI'j}$ ($\overline{p}_{kI'j} = \prod_{i \in I'} p_{ki'j}{}^{a_{ki'j}}$, $\overline{n}_{kI'} = \prod_{i \in I'} n_{ki'}{}^{a_{ki'j}}$, and $\overline{\alpha}_{kI'j} = \prod_{i \in I'} \alpha_{ki'j}{}^{a_{ki'j}}$ standing for, respectively, the average import price, number and quality image of varieties originating from the set of i's competitors I').

Consequently, in this model, relative market shares depend on relative prices as well as two differentiation terms, the relative quality image and number of varieties, plus an invariant factor. Therefore, exporters can increase their market shares by lowering their prices with respect to those of their foreign competitors, or by reinforcing their relative differentiation effort in order to raise their relative number of varieties or modify their relative image of quality to their advantage. Note that the coefficient of the price factor in (6) is strictly negative, due to condition $\sigma > 1$.

3 Toward a testable trade equation.

Equation (6) has to be transformed into a testable equation. In this respect, two crucial points have to be mentioned.

Firstly, trade prices are measured with error through unit values of trade, which may cause correlations between this imperfectly measured explanatory variable and the perturbation of the model. Note however that we have calculated unit values of trade not only in time variations but also in cross-section. In other terms, we expect to capture at least part of the spatial dimension of prices. This is a consequential point due to the potentially important influence of spatial effects in panel modelling. Besides, we have considered that import unit values would be a more convincing approximation for bilateral prices than export unit values, as the former take into account price competition between exporters at the entry of market *j* (transport and other transaction costs from any exporting country to market *j* being included in CIF import unit values)[3]. For the same reason, import declarations have been chosen as a theoretically more satisfactory measurement of bilateral trade flows than export declarations in the context of our model. Note that the under-estimation of intra-European import declarations since the creation of the INTRASTAT system of measurement of intra-EU trade flows in 1993 should be notably limited by the definition of the dependent variable of our trade equation, which is based on a ratio of imports (rather than on import levels). All trade variables have been calculated on the basis of six-digit data originating from the COMEXT data base of Eurostat. On this basis, we have calculated more aggregated trade flows corresponding to the nomenclature of the COE survey.

Secondly, differentiation terms (n_{kit}) and (α_{kijt}) - where time index t is now explicit - are unobservable. Omitting these factors or pretending to capture them through fixed effects might lead to a biased price elasticity and consequently to a bad evaluation of the elasticity of substitution σ, especially in highly differentiated industries. Therefore, we have decided to build proxies of these explanatory factors. As will be explained below, our focusing on the four biggest EU countries (France, Germany, Italy, United Kingdom) results from the availability of quality image measures in the case of these member States. Consequently, for each couple of countries (*i*,*j*) considered among this set of four member States, the sub-set of competitors *I'* of the theoretical model will be restricted to the two other EU countries.

The quality image proxy derives from the "Image of European products" annual survey of the COE. This survey consists in interviewing a panel of importers from different EU member States concerning their perceptions of the relative characteristic features of products originating from other EU countries in their sectors of activity. Specific questions notably deal with product quality, notoriety, degree of innovation, price, and ratio of quality to price, depending on geographic origin. This exceptional source therefore provides us with a purely exogenous piece of information on perceived quality with respect to trade data. Moreover, quality images collected from this survey depend on both the intrinsic quality of products and importers' subjective perceptions, which may differ significantly from one importing country to the other. In this respect, bilateral quality images which result from the COE survey are perfectly consistent with the theoretical preference parameters (α_{kijt}).

Survey data refer to years 1992 to 1997 for different kinds of products, namely "consumer goods" (split up into four sub-sectors: food, hygiene, lodging and clothing), and "other goods" (consisting of three sub-sectors: raw intermediate products, mechanical goods and electric goods other than consumer products) – Cf. Appendix 1. As was mentioned above, we have focused on the four main EU countries (as producing and purchasing countries), for which the COE survey results are available on a sufficiently long period and can be considered to be enough robust. Our quality image proxy is calculated on the basis of the answers to the four questions: "*In terms of quality levels, do you think that French / German / British / Italian products are: - the most competitive ones (mark = 1) - as competitive as those from other countries (mark = 2) - less competitive than those from other countries (mark = 3) - not competitive at all (mark = 4)?*".

Each interviewed importer is supposed to answer this question for products from every geographic origin, except from his or her own country. Let $imqual_{kijt}$ denote the percentage of interviewed importers from country j in sector k answering 1 or 2 to the question relating to country i in the survey performed in year $t-1$. We define the proxy for relative quality image ($imageq_{kijI't} = \alpha_{kijt} / \overline{\alpha}_{kI'jt}$) as:

$$im\widetilde{a}geq_{kijI't} \equiv imqual_{kijt} \Big/ \overline{imqual}_{kI'jt}$$

using the formula defining every theoretical explanatory variable in (6), including $imageq_{kijI't}$, with $\overline{imqual}_{kI'jt} \equiv \prod_{i \in I'} imqual_{ki'jt}{}^{a_{ki'j}}$ and $I' = \{France, Germany, Italy, United\ Kingdom\} \setminus \{i, j\}$. Weights $(a_{ki'j})$, defined in section 2, are relating to 1991 in order to avoid endogeneity problems. Although $Log(im\widetilde{a}geq_{kijI't})$ does not derive from a quantitative measurement of quality images, we expect it to be at least positively correlated with $Log(imageq_{kijI't})$. However, several technical points deserve some comments.

The reason for defining $imqual_{kijt}$ on the basis of the survey relating to year $t-1$ originates from the fact that COE surveys are performed in October, i.e. late each year. Therefore, we consider that a survey performed in October of year t might reveal quality images which are closer to those operating in year $t+1$ than in year t in terms of consumer choices. Above all, this choice may protect us from possible endogeneity problems during the estimation process.

It might seem more intuitive to define $imqual_{kijt}$ on the basis of responses to mark 1 only, or at least to over-weight percentages of marks 1 with respect to those of marks 2 in the calculus of $imqual_{kijt}$. We experimented all these alternatives, using several possible weights. However, the indicator that we finally chose proved to lead to more convincing results, due without doubt to its higher robustness. In fact, cumulated percentages of marks 1 and 2 prove to be much more regular over time than percentages of marks 1 (respectively 2) alone, which are more volatile. We find the same result concerning cumulated percentages of marks 3 and 4 compared to those of marks 3 and 4 considered separately. This property of the COE survey results suggests that the interviewed have a clearly positive or negative opinion concerning products from other EU countries, but not as qualified as to choose without doubt between the two positive (or two negative) modalities of questions asked in the COE survey. Our definition of the quality image proxy takes this characteristic feature of the survey data into account.

Ignored by the theoretical model, multinational companies, intra-firm trade and vertical integration may loosen the link between national product quality perceptions and geographic origins of trade flows. The fact that the surveyed are import professionals (who can be viewed as "experts" in their field) does limit the difficulty, but may not suppress it. This constitutes a possible limitation to the present study.

COE surveys successively deal with consumer goods (1992, 1994, 1996) and "other" goods (1993, 1995, 1997). Therefore, we have had to reconstitute annual indicators from biennial ($imqual_{kijt}$). Assuming that national quality images are relatively stable structural variables (which is confirmed by the evolution of our proxy), we have filled missing years with the simple arithmetic means of two successive biennial indicators.

As for the number of varieties (n_{kit}), the theoretical model suggests to build a proxy on the basis of sectoral production[4] or, preferably, GDP, due to the existence of intermediate consumption (the latter was neglected in the theoretical model, with no practical incidence on our trade equation expression and estimation results). Unfortunately, we do not have GDP data corresponding to COE sectors at our disposal. Therefore, we have introduced two factors:

- a relative GDP factor at macro level: $GDP_{ijI't} = GDP_{it} \big/ \overline{GDP_{I'jt}}$,

- a relative specialisation factor: $spe_{kijI'} = (X_{ki.} \,/\, X_{.i.}) \,/\, (\overline{X^{j}_{kI'.} \,/\, X^{j}_{.I'.}})$,

where, for consistency purpose with respect to the denominator of the theoretical explanatory variable ($variety_{kijI't} = n_{kit} / \overline{n}_{kI't}$) in (6):

$$\overline{GDP}_{I'jt} = \prod_{i' \in I'} GDP_{i't}^{\;a_{ki'j}} \text{ and } (\overline{X^j_{kI'.} / X^j_{.I'.}}) = \prod_{i' \in I'} \left(X_{ki'.} / X_{.i'.} \right)^{a_{ki'j}}.$$

The relative specialisation factor proxies the unobserved sector structure of GDP through that of exports. The closer and the more stable the structures of national GDPs in terms of tradable and non tradable products, the more satisfactory this approximation. As the countries taken into account are highly integrated European member States of relatively comparable sizes, we have assumed that this approximation was acceptable. In this case, the product of the two preceding factors constitutes a proxy for relative GDP (and consequently for the relative number of varieties $variety_{kijI't}$) at sector level. However, due to the approximation made, as well as to the fact that the specialisation factor has been calculated on a reference year (1991) rather than on a current basis[5], we have tested a trade equation in which the logarithms of these two factors are introduced separately. Therefore, it is possible to check whether the estimated coefficients of these two explanatory variables are similar (in conformity to intuition) or not.

National GDP data, expressed in 1991 prices, derive from the CHELEM data base of the CEPII. They have been smoothed over three years (t, $t-1$ and $t-2$) in order to capture the structural dimension of product variety rather than short-term economic fluctuations: the weights used in the smoothing are respectively 0.3, 0.4 and 0.3 for t, $t-1$ and $t-2$. Overall export data by sector ($X_{kl.}$) and at macro level ($X_{.l.}$) in 1991 originate from export declarations of the COMEXT data base.

Finally, we have replaced the invariant factor $c_{kijI'}$ with a linear combination of miscellaneous fixed effects, an intercept and a relative distance effect ($dist_{ijI'} = dist_{ij} / \overline{dist_{I'j}}$) based on a ratio of absolute distances, the absolute distance between two countries *i'* and *j* (noted $dist_{i'j}$, $i' \in I' \cup \{i\}$) being defined as that between the latter countries' two capital towns. Note that, as relative unit values of trade are not indices, distance is not supposed to capture part of relative transport costs, unlike in many gravity models. Moreover, as our proxy for relative quality image encompasses importers' subjective perceptions, distance is in principle not needed either to capture part of the latter. Nonetheless, as is suggested by Anderson and Marcouiller (1999) or Rauch (1999), introducing relative distance may enable us to capture other kinds of obstacles to trade than those taken into account through relative prices, or more probably (within a EU country sample) to limit the estimation effects of imperfect price and quality image measurements.

To sum up, the non restricted equation to be estimated is:

$$\begin{aligned} Log\left(mshare_{kijI't}\right) = -(e_p - 1)Log\left(pr\tilde{i}ce_{kijI't}\right) + e_g Log\left(GDP_{ijI't}\right) + e_s Log(spe_{kijI'}) \\ + e_q Log\left(im\tilde{a}geq_{kijI't}\right) - e_d Log\left(dist_{ijI'}\right) + fixed\ \ effects + intercept + u_{kijI't} \end{aligned} \quad (7)$$

The parameters of interest (referred to by *e*) may not be equal to the theoretical coefficients as all explanatory variables are proxies. However, the theoretical model gives some indications on the expected values for the estimated coefficients that should be satisfied so as to be considered to be convincing. The price elasticity e_p should be homogeneous to an elasticity of substitution σ, i.e. be strictly superior to unity. The coefficients associated with the variety proxies (e_g and e_s) should be of the same order of magnitude. They should also be either equal to unity (in accordance with Krugman, 1980) or, at least, inferior or equal to unity (in accordance with a modified version of our theoretical framework allowing for multi-variety producing firms[6]). However, the fact that relative GDP varies across time while the specialisation variable does not may induce some coefficient asymmetry. That is why we have tested the non restricted equation (7) in which e_g and e_s are not a priori set to be equal. The coefficient of the quality image proxy (e_q) should be strictly positive. Its value depends notably on the elasticity of substitution between varieties as well as the correlation between the proxy and the true relative quality image, which may significantly differ from unity, due to the qualitative foundation of the ($imqual_{kijt}$)

percentages. Finally, the coefficient of relative distance (e_d) should be negative and probably close to zero.

The perturbation $u_{kijI't}$ originates from the difference between theoretical variables and proxies. $u_{kijI't}$ also takes into account possible exceptional events and the parts of potential missing variables that are orthogonal to our explanatory factors. Time index *t* corresponds to years 1993 (or 1994) to 1997. Note that we « loose » the first year when COE survey data are available, namely 1992 (consumer goods) or 1993 (other goods), due to the construction of the quality image proxy (the $imqual_{kijt}$ percentages being set to be relating to the COE survey performed in year $t-1$). Country indices i, j are relating to France, Germany, Italy, or the UK, while sector index *k* represents either food, clothing, hygiene, lodging, raw intermediate products, mechanical goods, or electrical goods other than consumer goods. In sum, we have $5 \times 4 \times 3 \times 4 = 240$ observations for consumer goods and $4 \times 4 \times 3 \times 3 = 144$ for other goods.

We have performed two sets of estimations: one on consumer goods, on the 1993-1997 estimation period, and the other pooling all goods together, from 1994 to 1997 (i.e. using $240 - 48 + 144 = 336$ observations). Each set of estimations has been compared with the results derived from a more traditional sub-model excluding the quality image variable (e_q restricted to zero), or both dimensions of product differentiation (e_q, e_g and e_s restricted to zero). The interesting aspect of such comparisons is that the latter enable us to study how estimated price elasticities are modified when the adding of relative quality image (respectively differentiation variables) into the model suppresses at least part of the quality (respectively product differentiation) dimension contained in the relative price effect.

Finally, we have estimated equation (7) using three alternative sets of econometric methods, to test the robustness of the results. Firstly, we have performed ordinary least squares (OLS). Secondly, we have tested different instrumental variable estimation methods (hereafter referred to as 2SLS, for 2 Stage Least Squares). In fact, the differentiation and price variables are measured with error. Moreover, their exogenous status is highly questionable. Therefore, they have been instrumented by the first lag of the relative price and GDP variables, as well as the unvarying factors of the model (the intercept, specialisation, distance, and the fixed effects which are included in the model, when there are some), plus an instrument for quality image defined in the same way as the quality image proxy, but where the $imqual_{klj}$ percentages are calculated on the basis of the results derived from the first available survey (namely 1992 for consumer goods and 1993 for other goods). In order to limit the risk of correlation between this variable and the perturbation of the model, we have performed this instrumental variable method on 1995-1997 for consumer goods, and 1996-1997 for the whole sample. Opting for these restricted estimation periods have enabled us to suppress any reference to the 1992 and 1993 surveys in $im\hat{a}geq_{kijI'}$ and, consequently, in the current perturbation $u_{kijI't}$ (remember that percentages $imqual_{kijt}$ are set to be relating to the COE survey performed in year *t*-1). Thirdly, to suppress heteroskedasticity and correlation from our estimation residuals, we have performed two quasi-generalised least square alternative methods, referred to as QGLS 1 and QGLS 2 and presented in Appendix 2. All these estimation methods lead to very similar results, which can be viewed as a sign of robustness.

4 Main estimation results

Our main estimation results are summarised in Tables 1 (consumer goods alone), 2 (all goods pooled together), and 3 (models without fixed effects). Coefficients' estimates are little affected by the econometric methods. However, when QGLS methods are applied, variances estimations drop sharply, which may modify the results of the significance and inequality tests. More interesting, results of models including quality image appear to be much more satisfactory than those of models excluding this variable.

Firstly, in equations taking quality image into account, the coefficient relating to the quality image proxy appears to be clearly significant and of the expected positive sign[7].

Secondly, models including quality image show a very positive property: their estimated price elasticities are strictly superior to unity, thus taking values in conformity with the underlying theoretical model. Note that these estimated price elasticities (around 1.2) are close to those obtained by authors using other quality proxies (Erkel-Rousse and Le Gallo; 2002), and of the same order of magnitude as those obtained by authors using innovation variables such as R&D expenses or numbers of patents (Greenhalgh, Taylor and Wilson, 1994; Magnier and Toujas-Bernate, 1994; Ioannidis and Schreyer, 1997; Anderton, 1999). Above all, they are superior to those obtained in models excluding quality image (see below). The increase in the price elasticity when adding quality image is easy to interpret. In models excluding quality images, the coefficient relating to the price factor takes into account a pure price effect (which is negative) plus the indirect positive incidence of product quality on market shares through prices ceteris paribus (therefore, the sum of the two effects is less negative than the pure price effect). When quality image is taken into account, its coefficient captures this indirect effect, which disappears from the price coefficient. The latter then becomes a « pure » price elasticity, which is homogeneous to an elasticity of substitution (and therefore must be superior to unity).

Besides, the coefficients associated with relative GDP and specialisation are systematically inferior to unity (between 0.3 and 0.6, depending on the model), which is consistent with a version of our theoretical framework allowing for multi-variety producing firms, as was explained above. In most models with fixed effects, the coefficient relating to GDP (around 0.6 or 0.7) is higher than that relating to the specialisation variable (around 0.3). As was stressed before, the asymmetric definition of the GDP and specialisation variables (the former being allowed to vary across time, contrarily to the latter) may break the expected similarity of their two coefficients. However, in most models without fixed effects (table 3), the two coefficients are very much alike (between 0.4 and 0.5), especially when all products are pooled together within the panel (for instance, the two coefficients derived from QGLS 1 are equal to, respectively, 0.47 and 0.49).

When the proxy for quality image is excluded from the model, as is the case in most empirical work on trade equations, several problems show up.

The estimated price elasticity drops significantly below unity, whatever the model and the econometric technique. As was explained above, the price elasticity here captures part of the positive quality effect on market shares, which explains its lower absolute value (of around 0.8 or 0.9, depending on the model). This is a crucial problem as the price coefficient can no longer be interpreted as an elasticity of substitution. Nor can it be easily linked to any other parameter of interest derived from the underlying theoretical model. If one had obtained such an estimated equation on the basis of a theoretical framework excluding vertical differentiation, the (apparent) inconsistency between the value taken by the estimated price elasticity and its expected order of magnitude (which should in any case be superior to unity) would have (rightly) suggested a specification problem.

In the models including fixed effects and the two variety proxies, the coefficient associated with relative GDP is now higher than expected. First, it is significantly superior to unity (around 1.2 or 1.3, depending on sectors taken into account), which is not consistent with theory (see above). This result should shed doubt on the validity of the estimations, as the two alternative theoretical foundations for the variety proxy used require that the elasticity of variety to production be either equal or inferior to unity. Moreover, the coefficient of relative GDP is often about twice as high as that relating to specialisation (which amounts to 0.7), while it should be of the same order of magnitude. In this respect, the coefficient associated with specialisation proves to be in conformity with the theoretical model. The problem definitely lies in the excessive value of the GDP coefficient. Remember, however, that the estimated value of this coefficient is notably affected by the presence of fixed effects (contrarily to the other parameters of interest). In fact, when fixed effects are excluded from the model, it becomes slightly inferior to unity (or even to 0.9 in the consumer good model) and closer to the specialisation coefficient (the latter then being close to 0.8) (table 3).

Last but not least, the way the price elasticity is modified when the variety proxies are excluded from the model may seem somewhat puzzling, at least at first sight. In fact, it can be shown that excluding variety from the market share equation leads to a modification of the theoretical price elasticity, the latter now encompassing both the price effect corresponding to the price elasticity in the model including variety plus the indirect incidence of variety on market shares through prices everything else being unchanged in the model. One can establish that the sign of this indirect effect depends on that of the partial correlation between prices and variety, ceteris paribus in the trade equation. This partial correlation can be seen as the linear equivalent of the (non linear) partial derivative of prices with respect to variety in the underlying theoretical model[8]. According to the latter, a rise in the number of varieties ceteris paribus should imply a drop in prices due to higher competition: in other terms, the partial derivative of prices with respect to variety is negative. Therefore, we can reasonably[9] expect the partial correlation between prices and variety to be negative, at least if the model is correctly specified[10]. Such a negative correlation implies that the price elasticity should be higher in equations excluding the variety proxies than in equations including them. Unfortunately, this property is not satisfied in the case of the equations excluding quality image (when the variety proxies are removed from the equation, the price elasticity decreases by more than 0.1). This counterintuitive result again suggests a specification error. Here again, the adding of the quality image proxy in the model solves this problem. In fact, in equations taking quality image into account, the excluding of the variety proxies leads to a slight, but significant, decrease in the price elasticity (of 0.05 or more, depending on the model), which is consistent with intuition. In reality, the intuition does not hold in equations excluding quality image because the price and variety effects encompass part of the quality image effect, instead of limiting themselves to pure price and variety effects.

Note that adding quality image induces a significant drop in the variety coefficients (by more than 40%), whatever the specification of the model. In fact, in equations excluding quality image, the coefficients of the variety proxies encompass a direct variety effect plus the indirect incidence of quality on market shares through variety everything else being unchanged in the model. Adding quality image in the equation leads to the suppression of this indirect effect from the variety coefficients. It can be shown that this indirect effect is of the sign of the partial correlation between variety and quality image, ceteris paribus in the equation (the argument being the same as that used above for prices and variety). The drop in the coefficients relating to the variety proxies implies that this partial correlation is negative, which might originate from the trade-off between variety and quality often suggested in theoretical models of product differentiation[11].

Table 4 illustrates the specific feature of raw intermediate products. In this sector of almost homogeneous products, price elasticity (which reaches values between 2.1 and 2.6, depending on the model) proves to be notably higher than in other sectors. Moreover, export performances seem to be essentially driven by low relative prices, contrarily to what happens for consumer goods, as well as for equipment and other intermediate goods.

5. Quality image versus innovation image

Responses to the following question of the COE survey: "*In terms of innovation, do you think that French / German / British / Italian products are: - the most competitive ones (mark = 1) - as competitive as those from other countries (mark = 2) - less competitive than those from other countries (mark = 3) - not competitive at all (mark = 4)?"* enable us to build an innovation image proxy on the same basis as the quality image proxy. This innovation image proxy proves to be highly correlated with that of quality image. This result is interesting as many theoretical models derived from the new trade theory establish a tight link between quality and innovation.

Moreover, several authors use R&D, the number of patents, or other innovation variables as proxies of quality in trade equations (Greenhalgh, Taylor, and Wilson, 1994; Magnier and Toujas-Bernate, 1994; Amable and Verspagen, 1995; Anderton, 1999; Carlin, Glyn and Van Reenen 1997; Eaton and Kortum, 2002; Ioannidis and Schreyer, 1997). Conversely, Fontagné, Freudenberg and Ünal-Kesenci (1998) suggest that trade specialisation in quality does not coincide with that in technological products. However, the approach of COE survey (like that of the formerly mentioned literature) differs radically from that of Fontagné et alii. In fact, the survey examines the innovation image of any set of products, while Fontagné and alii focus on technological products.

We have performed a set of estimations using image proxies based on innovation instead of quality, for comparison purpose. Whatever the panel (consumer goods or all products pooled together), we find the same qualitative results when replacing quality image with the innovation proxy. As far as consumer goods are concerned, the econometric adjustment seems to be slightly better when using the quality variable (table 5). Besides, an attempt to include both quality and innovation into the model suggests that the quality variable might dominate as an explanatory variable for market shares. However, an ambiguous collinearity diagnosis between quality and innovation sheds doubt on the robustness of this result. In addition, estimations performed on the whole sample no longer suggest any superiority of the quality criterion over that relating to innovation. As the degree of collinearity between quality and innovation is much higher for raw intermediate goods and for capital goods than for consumer goods, it becomes impossible to evaluate which of the two criteria might be preferable.

The product-cycle hypothesis might play a role in these results (Feenstra and Rose, 2000). In any case, these estimations suggest that, had the quality variable not been available, the choice of a proxy based on the innovation criterion would have led to very similar results, which argues in favour of the empirical literature mentioned above. However, we must be cautious as regards the general incidence of our results. In fact, our innovation criterion might well be much closer to our quality criterion than any other innovation indicator based on R&D expenses or the number of patents...

6 Conclusion

In this paper, we have aimed at showing that more convincing estimated trade price elasticities can be obtained by controlling product quality in trade equations. In this purpose, we have estimated trade equations including a quality image proxy derived from survey data. Our estimation results, based on panel data for the four main EU Member States, confirm our initial intuition. A contrario, these results suggest that traditional models (especially macro-econometric ones) ignoring the dimension of product quality lead to under-estimated trade price elasticities and thus to incorrect evaluations of economic policy implications in open countries.

One might expect true price elasticities to be even higher than those derived from our estimations, at least for the most competitive industries. In fact, our approach has not led to as high price elasticities as those (estimated using different methodologies and kinds of data) by Hummels (2000), Head and Ries (2001), Eaton and Kortum (2002), Baier and Bergstrand (2001), Clausing (2001), or Erkel-Rousse and Mirza (2002) for instance. Moreover, low mark-up estimates or account rates of return are usually observed at industry levels (see Schmalensee, 1989, and Bresnahan, 1989), which may be consistent with relatively high levels of substitution elasticities, at least in industries characterized by monopolistic competition. We could probably obtain higher price elasticities, had we both more accurate proxies of quality perceptions and better measures of relative prices at our disposal, or at least could we build more sophisticated instruments for the latter. However, even in such an ideal context, we might need more broken-up data as well. Now, we have had to stick to the relatively aggregated product classification of the COE survey in this respect, which, besides, has prevented us from studying industry and country heterogeneity thoroughly.

Results obtained when using the economic approach consisting in taking product quality into account or, alternatively, an econometric method based on the definition of sophisticated instrumental variables for trade prices suggest that each approach succeeds in correcting part of the under-estimation of price elasticities, but not the whole of it. Unfortunately, up to now (at least to our knowledge), there has not been available direct broken-up quality measures which would enable one to mix the two methodologies.

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Appendix 1: The product classification of the COE survey

A. CONSUMER GOODS

Food: Meat, Fish, Milk, eggs, Vegetables, Fruit, Coffee, tea, spices..., Cereals, Flours, Fats and food oils, Meat-based preparations, Sugar, sweets, Cocoa, Flour-based preparations, Fruit and vegetable-based preparations, Various food preparations, Fruit juice...

Clothing: Leather goods, Hosiery, Clothes other than hosiery, Footwear, Watches and other accessories in precious metals, Wrist watches...

Hygiene: Drugs and medicine, Essential oils, perfumes, make-up, Soaps, Cleaning products, Polishes and creams for shoes, wax polish...

Lodging and other consumer goods: Contact lenses, glasses, Lenses, prisms, Binoculars, Cameras, Movie cameras and projectors for films <16mn and super 8, Slide projectors, Photograph films, Carpets, floorings, Linen of household, Non medical furniture, Household refrigerators, Dish washers, Washing machines, Other electrical household goods, Electric razors, Water-heaters, ovens, radiators household cooking stoves, Equipment and electronic and hi-fi consumer accessories, Radios and TV, Jewellery, toys, sport articles (other than fair carousels), Dishes..., Candles, Tapestries, Umbrellas, parasols...

B. OTHER GOODS

Raw intermediates goods: Marbles and other chalk stones, Gypsum, plaster, Stones for lime and for cement, Lime, Cements, Inorganic chemical products, Organic chemical products, Fertilisers, Tannins, pigments, paints, varnishes, Powders and explosives, Various products of chemical industry, Plastic and plastic products, Rubber and rubber products, Woods (other than sawed wood and furniture) and cork, Wood pulp and other fibrous pulps, Paper and cardboard, Wool, Cotton, Other vegetal fibre textiles, Artificial or synthetic filaments and fibres, Felt, special threads, strings and ropes, Special material (laces, velvet), covered, or laminated material, Materials of hosiery, Stone, plaster cement products, Ceramics, Glass and works in glass, Cast iron, iron and steel, Works in iron or steel, Copper and brass and works in copper, Nickel and works in nickel, Aluminium and works in aluminium, Lead and works in lead, Other common metals and works in these matters, Tools and kits, Works in common metals, Ball bearings ...

Electric goods (other than consumer goods): Industrial ovens, Professional typewriters, Calculators, electric and electronic cash registers, Computers and office machines and parts of these machines, Electric engines and generators, Parts of machines, Electrical transformers, Electromagnets, Electrical batteries and accumulators, Electrical starters and other parts of engines, Lamps, Electrical welders and brazers, Electrical signalling, Various electronic and electrical equipment, Optical fibres ...

Mechanical goods (other than consumer goods): Elements of railway tracks, Reservoirs, casks, vats, Turbines, engines and industrial machines, Specialised machines for particular industries, Parts of machines, Electromechanical hand tools, Tanks, Measuring devices, Arms, munitions and their parts and accessories ...

<u>Appendix 2</u>: Two alternative Quasi-Generalized Least Squares methods

The presence of heteroskedasticity and autocorrelation in models estimated using OLS and to a lesser extent 2*SLS* introduces a bias in the calculation of the *T*-statistic derived from these estimations. In our multi-dimensional panel data, it is likely that the variance-covariance matrix of residuals contains not only variable residual variances but also some non-zero, non-diagonal elements. In fact, for given sector *k* and time *t*, one can expect the trade performances of a given country *i* on different export markets to be correlated. Similarly, on a given importing market (k, j) at time *t*, relative market shares of exporters *i* = 1 to 3 can be expected to be negatively correlated. Our specification of the dependent variable therefore prevents us from simply correcting heteroskedasticity using weighed least squares estimators. Whatever the calculation method of the estimated variance-covariance matrix, we have assumed that the essential source of correlation within OLS residuals came from cross-section autocorrelations.

1) First method (referred to as QGLS1):

Here, we aim at taking into account correlations between relative market shares of each exporter *i* on its three export markets (k, j), *j* = 1 to 3. Let C^1_{ki} be the square matrix of $(I-1)\times(I-1)$ elements $\left(C^1_{jj'ki}\right)$ defined as $C^1_{jj'ki} = (1/T)\sum_{t=1}^{T}\hat{u}_{kijt}.\hat{u}_{kij't}$ where $(\hat{u}_{kijt})$ denotes the vector of OLS residuals, *T* the number of years in the panel, and *j* and *j'* two importing countries. Let:

$$A^1_k = \begin{bmatrix} C^1_{k1} & 0 & 0 & 0 \\ 0 & . & 0 & 0 \\ 0 & 0 & . & 0 \\ 0 & 0 & 0 & C^1_{kI} \end{bmatrix},\quad B_1 = \begin{bmatrix} A^1_1 & 0 & 0 & 0 \\ 0 & . & 0 & 0 \\ 0 & 0 & . & 0 \\ 0 & 0 & 0 & A^1_K \end{bmatrix}, \text{ and } \quad \Omega_1 = \begin{bmatrix} B_1 & 0 & 0 & 0 \\ 0 & . & 0 & 0 \\ 0 & 0 & . & 0 \\ 0 & 0 & 0 & B_1 \end{bmatrix} \begin{matrix} t=1 \\ . \\ . \\ t=T \end{matrix}$$

The QGLS1 estimator of the vector of coefficients β in the model: $Y_1 = X_1\beta + u_1$, where observations are classified by increasing (t,k,i,j) (the last index of the quadruplet being the first to move, then the third index, then the sector index, and finally the time index), is $\hat{\beta}_1 = \left(X_1'\Omega_1^{-1}X_1\right)^{-1}\left(X_1'\Omega_1^{-1}Y_1\right)$.

2) Second method (referred to as QGLS2):

Here, we aim at taking into account correlations between relative market shares of exporters *i* = 1 to 3 on each market (k, j). We define the QGLS2 method in the same way as previously, but on the basis of the square matrix C^2_{kj} of $(I-1)\times(I-1)$ elements $\left(C^2_{ii'kj}\right)$ defined as: $C^2_{ii'kj} = (1/T)\sum_{t=1}^{T}\hat{u}_{kijt}.\hat{u}_{ki'jt}$ where $(\hat{u}_{kijt})$ denotes the vector of OLS residuals and *i* and *i'* two exporting countries, observations being this time classified by increasing (t,k,j,i)

	Quality image excluded					Quality image included						
Estimation method[#]	*QGLS* 1	*QGLS* 1	*OLS*	*QGLS* 1	*QGLS* 2	2 *SLS*	*QGLS* 1	*QGLS* 1	*OLS*	2 *SLS*	*QGLS* 1	*QGLS* 2
Quality image (e_q)	-	-	-	-	-	0.33	0.33	0.33	0.25	0.27	0.25	0.24
						(16.71)	(77.43)	(56.54)	(11.90)	(9.78)	(41.08)	(43.18)
Price [-($e_p - 1$)]	0.20	0.21	0.08	0.08	0.07	-0.25	-0.24	-0.23	-0.18	-0.20	-0.18	-0.17
	(85.31)	(91.31)	(4.58)	(26.51)	(26.32)	(-8.05)	(-35.85)	(-26.01)	(-7.14)	(-6.19)	(-22.78)	(-27.66)
GDP (e_g)	-	0.84	1.29	1.28	1.28	-	-	0.33	0.64	0.40^a	0.62	0.60
		(20.20)	(6.81)	(35.00)	(41.89)			(5.83)	(4.03)	(1.85)	(11.82)	(15.71)
Specialisation (e_s)	-	-	0.71	0.71	0.71	-	-	-	0.29	0.24	0.28	0.30
			(14.59)	(56.69)	(77.31)				(5.50)	(3.49)	(27.86)	(22.06)
Distance (e_d)	-0.23	-0.21	-0.13	-0.13	-0.13	-0.18	-0.18	-0.17	-0.15	-0.16	-0.15	-0.15
	(-22.35)	(-18.90)	(-3.11)	(-14.55)	(-23.24)	(8.04)	(-14.17)	(-13.17)	(-4.45)	(-3.61)	(-22.67)	(-22.44)
Number of obs.	240	240	240	240	240	144	240	240	240	240	240	240
R^2	0.994*	0.992*	0.765	0.996*	1.000*	0.834	0.996*	0.996	0.854	0.850*	0.999*	0.999*
Root *MSE*	1.012	1.014	0.365	1.016	1.014	0.298	1.013	1.016	0.288	0.287	1.017	1.006
F-Stat	$0.6\ 10^4$	$0.4\ 10^4$	108	$0.6\ 10^4$	$0.2\ 10^4$	116	$0.8\ 10^4$	$0.7\ 10^4$	169	96	$1.8\ 10^4$	$4.0\ 10^4$
Multicollinearity**	No	No	No	No	No	No	No	No	No	No	No	No
(max. cond. index)	(2)	(6)	(7)	(6)	(12)	(5)	(10)	(15)	(8)	(9)	(19)	(23)
Heteroskedasticity	No	No	Yes	No	No	?	No	No	Yes	Yes	No	No
(*P*-value)	(1.000)	(1.000)	(0.004)	(1.000)	(1.000)	(0.012)	(0.999)	(1.000)	(0.000)	(0.004)	(0.994)	(1.000)

The models are estimated with an intercept plus 3 crossed fixed effects of the "exporting country × sector" type (namely: Italy × clothing, Germany × hygiene, Germany × other consumer goods), which encompass all the potential other fixed effects. Full results are available upon request.
Numbers in parentheses below each estimated coefficient are *T*-statistics. n = non significant at 10%; a (respectively b, c, no exponent) = significant at 10% (respectively 5%, 1%, any usual level: *P*-value < 0.01). * = Corrected R^2 (in models having a « corrected » intercept, due to *QGLS*).
** The positive multicollinearity diagnosis, when it occurs, affects specialisation, the intercept, the quality image proxy, and distance.

<u>Table 1</u>: Estimation Results for Consumer Goods.

	Quality image excluded				Quality image included					
Estimation method[#]	*QGLS* 1	2 *SLS*	*QGLS* 1	*QGLS* 2	*QGLS* 1	*QGLS* 1	*OLS*	2 *SLS*	*QGLS* 1	*QGLS* 2
Quality image (e_q)	-	-	-	-	0.31	0.29	0.22	0.23	0.22	0.22
					(169.98)	(117.12)	(13.52)	(9.05)	(62.30)	(58.3)
Price [-($e_p - 1$)]	0.22	0.06	0.06	0.06	-0.19	-0.17	-0.15	-0.16	-0.14	-0.15
	(110.16)	(2.69)	(38.66)	(29.03)	(-56.15)	(-44.97)	(-7.64)	(-5.26)	(-40.15)	(-34.45)
GDP (e_g)	-	1.16	1.19	1.20	-	0.43	0.68	0.62	0.68	0.68
		(8.45)	(390.56)	(92.40)		(38.45)	(8.21)	(5.02)	(49.47)	(38.86)
Specialisation (e_s)	-	0.71	0.72	0.72	-	-	0.34	0.31	0.34	0.34
		(9.99)	(135.41)	(75.84)			(7.07)	(4.41)	(35.64)	(33.45)
Distance (e_d)	-0.26	-0.14	-0.15	-0.15	-0.21	-0.16	-0.15	-0.15	-0.15	-0.15
	(-94.90)	(-3.02)	(-47.80)	(-55.44)	(-49.69)	(-71.56)	(-6.02)	(-4.09)	(-33.84)	(-30.31)
Number of obs.	336	168	336	336	336	336	336	168	336	336
R^2	1.000*	0.759	1.000*	0.999*	1.000*	1.000*	0.859	0.847	1.000*	0.999*
Root *MSE*	1.014	0.348	1.017	1.017	1.014	1.016	0.267	0.277	1.017	1.018
F Statistic	$10.2\ 10^5$	45	$1.6\ 10^7$	$2.3\ 10^4$	$8.5\ 10^5$	$1.2\ 10^5$	164	71	$1.0\ 10^5$	$1.9\ 10^5$
Multicollinearity**	No	No	Yes*	No	No	No	No	No	No	No
(max. condition index)	(14)	(4)	(38)	(10)	(6)	(6)	(7)	(7)	(12)	(18)
Heteroskedasticity	No	No	No	No	No	No	Yes	No	No	No
(*P*-value)	(1.000)	(0.869)	(1.000)	(1.000)	(1.000)	(1.000)	(0.002)	(0.300)	(1.000)	(1.000)

The models are estimated with an intercept, 1 simple fixed effect (capital goods) and 6 crossed fixed effects, among which the 3 used in the models for consumer goods (see legend of table 1), plus 2 other fixed effects of the "exporting country × sector" type (namely: France × capital goods, Germany × raw intermediate products), plus a fixed effect of the "importing country × sector" type (France × capital goods), which encompass all the potential other fixed effects. Full results are available upon request.

Numbers in parentheses below each estimated coefficient are *T*-statistics. n = non significant at 10%. a (respectively b, c, no exponent) = significant at 10% (respectively 5%, 1%, any usual level: *P*-value < 0.01). * = Corrected R^2 (in models having a « corrected » intercept, due to *QGLS*).

** The positive multicollinearity diagnosis, when it occurs, possibly affects the intercept and 2 fixed effects (Germany × other consumer goods and Germany × raw intermediate products).

Table 2: Estimation Results for All Products Together.

	Consumer goods				All Products Together			
	Quality excluded		Quality included		Quality excluded		Quality included	
Estimation method#	*QGLS* 1	*QGLS* 1	*QGLS* 1	*QGLS* 1	*QGLS* 1	*QGLS* 1	*QGLS* 1	*QGLS* 1
Quality image (e_q)	-	-	0.34 (88.44)	0.21 (42.94)	-	-	0.32 (213.64)	0.20 (84.37)
Price [-($e_p - 1$)]	0.24 (64.70)	0.06 (25.32)	-0.24 (-43.09)	-0.15 (-26.7)	0.26 (217.99)	0.06 (44.79)	-0.17 (-69.27)	-0.12 (-47.07)
GDP (e_g)	-	0.87 (33.20)	-	0.39 (12.86)	-	0.98 (314.98)	-	0.47 (65.53)
Specialisation (e_s)	-	0.79 (80.01)	-	0.47 (37.87)	-	0.81 (200.76)	-	0.49 (98.15)
Distance (e_d)	-0.07 (-8.85)	-0.08 (-9.29)	-0.02[n] (-1.55)	-0.06 (-8.00)	-0.17 (-85.60)	-0.14 (-95.69)	-0.10 (-49.42)	-0.13 (-64.81)
Number of obs.	240	240	240	240	336	336	336	336
R^2	0.993*	0.995*	0.996*	0.998*	1.000	1.000*	1.000	1.000*
Root *MSE*	1.005	1.010	1.008	1.010	1.004	1.007	1.005	1.007
F-Stat	$1.1\ 10^4$	$0.9\ 10^4$	$1.5\ 10^4$	$1.8\ 10^4$	$2.7\ 10^6$	$4.5\ 10^6$	$4.3\ 10^6$	$1.3\ 10^6$
Multicollinearity (max. condit. index)	No (1)	No (6)	No (8)	No (14)	No (2)	No (3)	No (6)	No (10)
Heteroskedasticity (*P*-value)	No (0.979)	No (1.000)	No (1.000)	No (1.000)	No (1.000)	No (1.000)	No (0.916)	No (1.000)

The models are estimated with an intercept (but without any fixed effects). Numbers in parentheses below each estimated coefficient are *T*-statistics. n = non significant at 10%. a (respectively b, c, no exponent) = significant at 10% (respectively 5%, 1%, any usual level: *P*-value < 0.01).
* = Corrected R^2 (in models having a « corrected » intercept, due to *QGLS*).

<u>Table 3</u>: Estimation Results – Models without Fixed Effects.

Estimation method#	Quality image excluded *QGLS* 1	2 *SLS*	*QGLS* 1	*QGLS* 2	Quality image included *QGLS* 1	*QGLS* 1	*OLS*	2 *SLS*	*QGLS* 1	*QGLS* 2
Quality image - non raw intermediate goods (e_q)	-	-	-	-	0.30 (84.50)	0.25 (63.70)	0.18 (10.19)	0.19 (6.59)	0.18 (46.24)	0.18 (33.25)
Price - raw intermediate goods [-($e_p - 1$)]	-1.41 (-68.58)	-1.20 (-3.77)	-1.15 (-43.78)	-1.16 (-50.53)	-1.61 (-64.58)	-1.53 (-62.98)	-1.38 (-7.97)	-1.42 (-5.09)	-1.37 (-63.14)	-1.36 (-35.80)
Price - other products [-($1 - e_p$)]	0.23 (119.22)	0.07 (3.40)	0.07 (42.53)	0.07 (33.01)	-0.17 (-30.90)	-0.15 (-27.41)	-0.12 (-5.19)	-0.12 (-3.51)	-0.11 (-22.60)	-0.12 (-17.31)
GDP (e_g)	-	1.13 (8.62)	1.16 (100.61)	1.16 (86.71)	-	-	0.84 (10.28)	0.80 (6.47)	0.84 (55.36)	0.84 (44.17)
Specialisation (e_s)	-	0.68 (10.02)	0.69 (128.78)	0.70 (72.95)	-	0.23 (31.82)	0.41 (8.50)	0.39 (5.42)	0.41 (49.41)	0.41 (33.15)
Distance (e_d)	-0.20 (-70.52)	-0.10^{b} (-2.24)	-0.11 (-21.91)	-0.11 (-16.48)	-0.19 (-41.49)	-0.20 (-44.02)	-0.12 (-4.60)	-0.12 (-3.07)	-0.12 (-25.04)	-0.13 (-16.52)
Number of observations	336	168	336	336	336	336	336	168	336	336
R^2	1.000	0.781	1.000*	0.998*	1.000	1.000	0.851	0.834	1.000*	0.999*
Root *MSE*	1.016	0.333	1.017	1.019	1.016	1.015	0.275	0.290	1.018	1.019
F Statistic	$2.9\ 10^{5}$	46	$4.4\ 10^{6}$	$1.5\ 10^{4}$	$2.1\ 10^{5}$	$1.4\ 10^{5}$	141	59	$8.9\ 10^{4}$	$2.2\ 10^{4}$
Multicollinearity (maximal condition index)	No (10)	No (4)	Yes** (62)	No (8)	No (11)	No (13)	No (7)	No (7)	No (13)	Yes** (30)
Heteroskedasticity (*P*-value)	No (1.000)	No (0.945)	No (1.000)	No (1.000)	No (1.000)	No (1.000)	Yes (0.000)	No (0.217)	No (1.000)	No (1.000)

The models are estimated with an intercept plus the 7 fixed effects defined in the legend of table 2. Numbers in parentheses below each estimated coefficient are *T*-statistics. n = non significant at 10%. a (respectively b, c, no exponent) = significant at 10% (respectively 5%, 1%, any usual level: *P*-value < 0.01). * = Corrected R^2 (in models having a « corrected » intercept, due to *QGLS*). The surveyed seem to find it difficult to answer the question concerning quality for raw intermediate goods, probably because the latter are little differentiated. Therefore, we have performed estimations with quality image for non raw intermediate goods only. However, the estimation results are little modified for the other sets of products - compare with tables 2 and 3.
** The multicollinearity possibly affects the coefficients relating to price (other products) and specialisation.

Table 4: Estimation Results for Models with Heterogeneous Price and Quality Image Effects (All Products Together).

	-	Quality	Quality + Innovation			Innovation		
Estimation method[#]	*QGLS* 1	*QGLS* 1	*OLS*	2 *SLS*	*QGLS* 1	*OLS*	2 *SLS*	*QGLS* 1
Quality image (e_q)	-	0.25	0.20	0.23^{n}	0.20	-	-	-
		(41.08)	(4.99)	(0.61)	(18.21)			
Innovation image (e_i)	-	-	0.05^{n}	0.04^{n}	0.05	0.19	0.25	0.19
			(1.36)	(0.11)	(4.41)	(10.38)	(8.29)	(24.14)
Relative price [-($e_p - 1$)]	0.08	-0.18	-0.18	-0.20	-0.19	-0.14	-0.19	-0.14
	(26.51)	(-22.78)	(-7.26)	(-6.18)	(-23.30)	(-5.62)	(-5.33)	(-13.75)
Relative *GDP* (e_g)	1.28	0.62	0.62	0.39^{a}	0.58	0.68	0.42^{a}	0.67
	(35.00)	(11.82)	(3.86)	(1.84)	(10.94)	(4.07)	(1.94)	(12.46)
Specialisation (e_s)	0.71	0.28	0.28	0.23	0.28	0.37	0.25	0.37
	(56.69)	(27.86)	(5.48)	(3.45)	(28.27)	(7.18)	(3.42)	(26.32)
Distance (e_d)	-0.13	-0.15	-0.16	-0.16	-0.16	-0.19	-0.20	-0.18
	(-14.55)	(-22.67)	(-4.65)	(-2.07)	(-22.27)	(-5.26)	(-4.56)	(-19.61)
Number of obs.	240	240	240	144	240	240	144	240
R^2	0.996*	0.999*	0.855	0.851	0.999*	0.840	0.846	0.997*
Root *MSE*	1.016	1.017	0.288	0.287	1.018	0.302	0.287	1.016
F Stat	$0.7\ 10^4$	$1.8\ 10^4$	151	85	$1.5\ 10^4$	151	93	$0.8\ 10^4$
Multicollinearity**	No	No	?	Yes	Yes	No	No	?
(max. condition index)	(6)	(19)	(17)	(134)	(31)	(8)	(9)	(24)
Heteroskedasticity	No	No	Yes	?	No	Yes	?	No
(*P*-value)	(1.000)	(0.994)	(0.004)	(0.010)	(0.926)	(0.001)	(0.033)	(0.997)

The models are estimated with an intercept plus the 3 crossed fixed effects defined in the legend of table 1.

Numbers in parentheses below each estimated coefficient are *T*-statistics. n = non significant at 10%; a (respectively b, c, no exponent) = significant at 10% (respectively 5%, 1%, any usual level: *P*-value < 0.01). * = Corrected R^2 (in models having a « corrected » intercept, due to *QGLS*).

** The positive multicollinearity diagnosis, when it occurs, possibly affects the coefficients relating to quality and innovation images, or innovation image and price (last column). The model with a condition index of 17 might be affected with multicollinearity, due to the ambiguous maximal condition index in the "intercept adjusted" model (16).

<u>Table 5</u>: Quality or Innovation? The Case of Consumer Goods.

[1] More recently, Feenstra (1994) addressed the issue of trade price measurement errors in terms of product differentiation, arguing that ignoring changes in the number of varieties supplied from each country leads to incorrect price evaluation. Taking these changes into account enabled him to obtain better price indices.

[2] This result is established ceteris paribus, as well as under the assumption of an implicit increasing link between q_{ki} and c_{ki}. This assumption expresses that high quality may be more expensive to produce than low quality. Assuming a link of this kind anticipates the endogeneisation of quality. However, this point is not developed in the present paper.

[3] Besides, import unit values have been smoothed so as to correspond to what is generally observed in terms of the progressive influence of prices on trade values, as well as to limit potential endogeneity problems. If t is the current year, the smoothing uses weights 0.3, 0.7 for respectively t and t-1. These weights derive from impulse functions resulting from dynamic models in time series econometrics.

[4] This approach supposes transposing at sector level a theoretical property established at firm level, which may imply some aggregation difficulties. Note that Krugman (1980) established a positive link between production and the number of varieties at sector (or even macroeconomic) levels, but in a less general framework (monopolistic competition with a high number of firms).

[5] We opted for this reference year to avoid possible endogeneity and simultaneity problems due to the approximation of national GDP structures by those of total exports, as well as to limit the effects of potential divergence between the evolution of export structures and that of GDP structures on the estimation period. Moreover, the 1991 constant export structure might better capture that of horizontal differentiation than its current equivalent during an estimation period marked by notable monetary shocks between France, Germany, Italy and the United-Kingdom.

[6] In such a model, values of e_g and e_s higher than unity would be incompatible with the existence of an optimal number of varieties in the producer optimum (Erkel-Rousse, 1997 and 2002).

[7] Admittedly, with values of around 0.2 or 0.3, it is notably lower than expected on the basis of the theoretical model. However, as was stressed above, many approximations have been performed to elaborate an annual proxy of quality images, which have undoubtedly prevented us from obtaining a precise quantitative estimate of market share elasticity to quality image.

[8] For a mathematical definition of the partial correlation between prices and variety as well as a thorough demonstration, see Erkel-Rousse (2002).

[9] If not automatically, due to the possible occurrence of exceptional configurations in which linearisation induces changes in the direction of variation of some variables with respect to some others.

[10] If an explanatory variable is omitted from the equation, the partial correlation between prices and variety may be biased in such a way that its sign be modified.

[11] See Neven and Thisse (1989). It could be shown that the endogeneisation of quality image in our theoretical model would also lead to the occurrence of such a trade-off between variety and quality under some conditions involving production and utility functions' parameters (more details available on request to the authors).