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Introduction

The current debate on protectionism *versus* trade liberalisation is undoubtedly one of the core topics in International Economics not only in terms of the policy implications of both regimes but also in order to attain a better theoretical understanding of the observed phenomena. As is the case with the expected benefits of globalisation processes for national welfare levels, the optimum structure of international trade and its derived policies as predicted by the principle of comparative advantages are put into question by the accumulated empirical evidence that contradicts the theoretical results (Krugman, 1995; Rodrik, 1997).

The validity of the Law of One Price (LOP) and the assumption of consumers considering goods produced in different countries as homogenous - and hence perfect substitutes – are two key assumptions required by the traditional trade models. However, the empirical evidence obtained up to now strongly suggests their validity is confined only to some specific cases, being the growing share of intra-industry trade one of the more outstanding non-explained issues.

The New Trade Theories have contributed, since the pioneering work of Helpman and Krugman (1985), to the explanation of many of these international Trade Theory puzzles, especially within a partial equilibrium framework. The applied research performed under this approach has used diverse instruments for evaluating the impact of policies, being econometric models one of the most popular choices. However, an insurmountable restriction that might hinder the robustness of the results so obtained is precisely them being framed within a partial equilibrium context, thus unable to capture all indirect effects.

A recently widespread alternative strategy is the use of Computable General Equilibrium Models (CGEM). Given they are designed so as to account for both the direct and indirect effects of policies they are potentially the most powerful available instrument that may overcome the above stated shortcoming of econometric modelling. However, their empirical implementation still faces many obstacles that limit the precision of the quantitative results of the simulations with them performed. One of these major restrictions is their need of a huge number of parameters that are not always available, and even at times impossible to estimate.

Focusing on CGEMs applied to trade, one of the above mentioned practical difficulties refers to the inclusion of monopolistic competition and endogenous differentiated goods in the models, so as to account for the stylised facts on the matter as reported in the empirical literature.

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A widely used strategy that allows for partially considering the above behaviour has been to assume exogenous national differentiation of goods. As a consequence, a renewed interest on the approach proposed by Armington (1969) has emerged.

The use of an Armington structure within the models implies that one key set of parameters in terms of the overall results of simulations is that of the elasticities of substitution between national and foreign varieties of each good. As such, the adequacy of the magnitude of these parameters is crucial for guaranteeing the robustness of the results derived from CGEMs simulations when performed for a particular case study.

While in the past the common practice among CGEMs builders has been to impose these values, either at will or by using those stemming from other existing applied research, it is currently agreed that this is not the best strategy, although at times it is the only feasible one. Furthermore, even if econometrically estimated parameters are available, they are frequently considered to be underestimates of both the real and theoretically expected elasticities, and at times even useless for performing reliable simulations with CGEMs (Riedel, 1988; Athukorala and Riedel, 1994).

However, according to the literature on international real business cycle (IRBC) the expected magnitude of Armington elasticities should not only be small but also quite in line with those obtained using econometric time series models.

The dichotomy might be explained in terms of the transitory/permanent character of shocks being modelled in IRBC and CGE models, assuming that economic agents have a differentiated reaction to each type of shock. IRBC models proxy the evolution of the terms of trade and its relation to trade balances using high frequency data, thus always modelling transitory shocks. Comparative statics performed with CGEMs, on the contrary, are used to compare two different structures arising from a change in the overall economic framework, as would be the case when a trade liberalisation takes place, or after a change in policies takes place, thus incorporating the effect of permanent shocks to the system (Ruhl, 2005).

Low estimated Armington elasticities may also arise due to ignoring the supply side behaviour, so that the so obtained estimators are actually an average of demand and supply elasticities. Further, such low values would imply that even small economies have enough market power to set international prices.

Consequently, a growing concern on the reasons underlying the smaller than expected estimated values of Armington elasticities has emerged. A strand of the literature has focused on the eventual role played by methodological and empirical issues (Gan, 2006; McKitrick, 1998; Saito, 2004; Shiells, Stern and Deardorff 1986; Welsch, 2006, among many others). Aiming at finding clues by means of comparing international studies, McDaniel and Balistreri (2003) perform a survey on the applied literature on Armington models that is taken as supporting evidence for stating several conclusions that have gained quite a general consensus among researchers.

According to the authors, high estimated Armington elasticities are more likely to be obtained when using cross section and pool cross section-time series datasets than when using time series information; long-run estimated elasticities are higher than short run elasticities; and estimates are higher the greater is the level of disaggregation of economic sectors used as proxies of goods and markets.

Although the above assessments are most sensible, we here state that it would be highly misleading to consider them as supporting evidence for choosing datasets with specific characteristics when estimating Armington elasticities by means of econometric models.

We found our argument on the fact that the above allegedly robust findings are based on the comparison of results obtained from econometric models that are in some cases estimated for unrelated spatial/temporal units of observation. Worse still, some of the estimators that are being compared stem from statistically misspecified models. Since none of the conclusions or results derived from misspecified models can be considered as robust, it is highly recommended not to include them when general comparisons are being made.

A further issue that has been analysed by some authors, although not addressed in the above mentioned survey, refers to the impact of the temporal frequency of the data on the estimated mean values (Gallaway, McDaniel and Rivera, 2003; Hertel, Ianchovichina and McDonald, 1997). However, no clear-cut results were found.

We focus on four particular issues - the temporal and/or spatial character of the data; the temporal frequency chosen; the aggregation level used to define a good; and the estimation methodologies. We critically review the current debate on these topics, using Uruguayan data as a means of testing for the validity of the different hypotheses discussed. Our conclusion points at the type of data not being the underlying cause of the small magnitude of estimates. Instead, the focus should be set on attaining statistically correct specifications of those models giving rise to the estimated elasticities and to building better quality datasets so as to guarantee a correct approximation to the theoretical concepts as well as consistency between variables.

Cross section versus time series data

One topic reviewed in the literature regarding the expected relative magnitude of Armington elasticities refers to the eventual role of using time series and/or cross section data. Most of the existent research has been done using temporal observations of different frequency or else pooling time series data with cross section information. Exceptions are Hummels (1999) and Bilgic *et al.* (2002), who estimate substitution elasticities at one point in time using data for several countries and US geographical regions, respectively.

The widespread conclusion in terms of the estimated relative values of the elasticities of substitution is that they are higher when using cross section data, so that results obtained from the latter are considered to be improved given the magnitudes required in performing simulations with CGEMs.

McDaniel and Balistreri (2003) arrive at the above conclusion by means of comparing 5 studies performed for the US using time series uniequational models with 2 multi-country gravity models[1]. However, it is worth noting that the estimators stemming from these studies are not always strictly comparable.

A first comment refers to the time series studies included in the comparison being referred to quite different time periods -two refer to the 60s-70s, two refer to the 80s and the fifth is performed for the 90s. As it has been already discussed in the literature (Gan, 2006; Hernández, 1998), the elasticity of substitution need not be stable in time, so that care should be put when comparing the corresponding estimates. Further, given that the estimated elasticity also captures the impact of trade barriers, it should be expected that it notably increases immediately after a liberalisation process to decline again afterwards (Welsch, 2006).

Secondly, time series analyses refer to the ease of substituting national by imported varieties of a single good, no matter the level of aggregation used for defining the good. As such, they cannot be directly compared to the value of estimates that are obtained by pooling sectors, as the latter refers to an average of the individual goods elasticities. However, the papers specifying gravity models also provide estimates by sector, the values of which are also in line with the conclusions of the survey. Nonetheless, we believe there is no evidence supporting that the result derives from the use of cross section data. In fact, Bilgic *et al.* (2002) use cross section data to model US regional trade flows, obtaining estimated elasticities of substitution of a lower magnitude than those reported by the papers cited in the review.

Gravity models make use of bilateral trade flows, while the time series studies surveyed are multi-country analyses. As such, comparisons should not be done in terms of the type of data used but in terms of the specification of the models themselves.

It is further argued that possible underlying causes for obtaining higher elasticities by means of using cross section information relate to the fact that time series models are generally specified as the reduced form of supply and demand equations and thus the estimates are an average of supply and demand elasticities. However, the statement is not true in absolute terms, but only when the specification of the models is not adequate or else the estimation method used is unsuitable, so that endogeneity is not properly taken into account. This is the case of 4 of the 5 papers reviewed. Shiells *et al.* (1986) does account for endogeneity of prices, also specifying the models in order to capture other relevant properties, such as simultaneous determination of relative demands for the different goods and the existence of observed and unobservable dynamic structures. It should be of no surprise that the estimated elasticities are indeed quite high for many goods, their values ranging from to 0.45 to 32.1.

Alternatively, it is argued that time series studies implicitly ignore the eventually persistent long run distortions that affect the supply side. The same above counterargument applies here, as a proper econometric treatment of the models would allow for surmounting the issue. Particularly, the analysis of the order of integration of the stochastic processes involved in the models would be most useful

[1] The papers included are: Stern, Francis and Schumacher (1976), Shiells, Stern and Deardorff (1986), Reinert and Roland-Holst (1992), Shiells and Reinert (1993), and Gallaway, McDaniel and Rivera (2003).

for an appropriate treatment of the persistent effects of shocks within the specification of the models. The topic has not been paid attention as frequently as it should have after the mid-eighties, least of all has the existence of cointegration among the processes been analysed in depth.

Regarding the 5 reviewed papers, only Gallaway *et al.* (2003) addresses the issue, dividing the processes in stationary and non stationary. Whenever both quantities and prices are stationary, the authors estimate first order dynamic models in levels (72 sectors out of 312) while when cointegrated processes are involved the models are estimated as first order ECMs (28 cases). Although the magnitude of the estimated elasticities is not high in most cases, no report is included in the paper on the statistical misspecification analysis performed to the models. Consequently, the robustness of the estimates cannot be assessed. Regarding the rest of the cases analysed (65%), given no cointegration relation was found the authors specify a model in first differences. Although the strategy allows for a correct application of standard estimation methods, the models so specified lack any economic meaning given the results of the cointegration analyses performed. The same strategy is also found in many other papers related to the estimation of Armington elasticities, and at times it is pursued even when the order of integration of the processes differs so that no balanced regression can be asserted and hence no sensible relation is being proposed (Fontes *et al.*, 2007; Gallaway *et al*., 2003; Gan, 2006; Gibson, 2003; González and Wong, 2005; among others). A notable exception is Welsch (2006).

It is worth noting that the suggestion on the alleged differences in estimators being due to the fact that in using cross section data the main source of variation stems only from permanent shocks to trade flows while when using time series data the impact of transitory shocks is the leading source, may apply only to those models estimated in differences, as in doing so the long run information in the data is explicitly being excluded from them.

Finally, the assessment on long run elasticities being higher than their short run counterparts is quite surprising, given no other result should be expected unless the processes involved are explosive or whenever a greater than 1 lag structure is modelled and overshooting processes take place. However, this is not the case discussed in the survey since the lag structure of the models reviewed is never of order greater than 1.

The above discussion suggests that it is not the temporal/spatial nature of the data used that determines higher or smaller estimators of the Armington elasticity, but the correct specification of the models and the use of suitable estimation methods instead. Further, the accumulated applied literature is still insufficient for providing evidence on the matter, because there are not enough analyses performed using both types of datasets for the same temporal-spatial case study.

The temporal frequency of the data

The highest frequency of the data used in the empirical literature is diverse and generally not chosen by the researcher but imposed by data availability. The issue should not be neglected as irrelevant due to the fact that some responses to changes in prices or to exogenous shocks may occur in a very short period of time, thus they may not be captured when using low frequency data.

Further, a significant difference among the short and long run response of relative demand to shocks or to changes in prices may also exist. Consequently, the relevant elasticity to be taken into account depends on the analysis to be performed based on its magnitude. In the particular case of CGEMs, the only meaningful estimate is that corresponding to the long run.

The best scenario when long run elasticities are being sought for would be given by the availability of annual data information sets covering a century, but the odds of being in such a world are yet too low. However, currently most researchers do have access to quarterly data, a frequency that generally provides with an adequate number of observations for performing cointegration analysis (around 80, equivalent to 20 years). This frequency further allows for identification of the lags registered between the instant in which a shock takes place and the moment when its effects are completely absorbed by the phenomenon under study.

The use of monthly data is at times necessary when the number of quarterly observations is not sufficient to perform certain analyses. Its main disadvantage is, however, that they include many atypical observations, frequently due to the way some data are generated. For example, imports are usually registered at the moment of entrance to a country and not when sold to consumers, so that the relative monthly demand of the two varieties of each good would be misrepresented by the proxy variable used. This is probably not so when a quarter is taken into account, at least for most final consumption goods and particularly for perishable goods.

The above mentioned aspect is reflected in a higher variability of estimates arising from models using monthly data with respect to those using quarterly information. Moreover, in many of the former cases models have to be estimated including a large set of binary variables in order to purge the estimation process from the perverse effect of these atypical observations.

According to Hertel, Ianchovichina and McDonald (1997) more inelastic elasticities are to be expected the lower the frequency of the data. This would be the case when the temporal aggregation hides adjustment patterns, such as overshooting effects or J-curve like effects, taking place in the very short run.

The above mentioned overshooting effects in the estimation of Armington elasticities are found in the case of several Uruguayan manufacturing sectors (Flores, 2008). Further, evidence is also there found relative to the existence of differentiated effects of changes in relative prices on relative demand depending on which of its components varies – international to local prices, exchange rate or tariffs. Although these differential effects can be clearly appreciated while working with monthly data, the difference in the parameters vanishes when quarterly information is used. However, no straightforward patterns are identified for the Uruguayan manufacturing sector regarding the magnitude of the estimated Armington elasticities when using monthly and quarterly data. As such, the result does not conform either to the non significant differences reported by Gallaway, McDaniel and Rivera (2003).

The level of aggregation for defining goods

The debate on the effects of aggregating data in Economics is long dated and is encompassed by the larger discussion on the microeconomic foundations of macroeconomic relations. The debate is also extended to the methodological implications of making inferences on the individual behaviour of agents by means of the estimated parameters obtained from models that use aggregated data (Theil, 1954). There is no consensus on the precision of estimates, since Theil's original conclusion stating a higher precision is to be expected from the use of aggregated data was afterwards contradicted by the results presented by Grunfeld and Griliches (1960). More recently, Pesaran (2003) has discussed the topic within what he denominated "*model selection problem*".

Regarding the literature on Armington elasticities, the topic includes additional dimensions. A major facet is that of the definition of 'goods' that may sustain the separability of preferences, in turn linked to the fact that whatever is defined as a good should have the property of being homogeneous except for its production origin.

The debate on aggregation within the applied literature on International Trade has been extensively accounted for, as well as its effects on the magnitude of the estimated elasticity of substitution (an early reference is Alaouze, 1977). The same can be said regarding the literature on CGEMs (Hertel, 1999). However, what cannot be discussed at this stage is the relevance of matching the level of aggregation used in the CGEMs to that of the data used in the econometric models from which the parameters needed by CGEMs are estimated (Hertel *et al.*, 2004).

Some authors have suggested that aggregating data results in underestimates of the elasticities (Hummels, 1999; Erkel-Rousse and Mirza, 2002; Gibson, 2003; McDaniel and Balistreri, 2003; among others).

Hummels (1999) provide robust evidence on the existence of a negative aggregation bias in the gravity models estimated, concluding that it is originated in the heterogeneity of goods included in aggregated categories. Hertel *et al.* (2004) suggest that the biases may also be linked to aggregating origins of imports while Erkel-Rousse and Mirza (2002) signals at unobservable differences in quality as possible sources of these negative biases that would be then associated to unmodelled endogeneity.

Finally, McDaniel and Balistreri (2003) arrive at the same conclusion by comparing the results reported by Gallaway, McDaniel and Rivera (2003) that define goods as equivalent to the result of production at a 4-digit ISIC industry level, with those in the paper by Reinert and Roland-Holst (1992) that use a 3-digit classification. However, it is worth noting that these case studies are performed using different time periods, so that the previously mentioned instability of the true elasticities may be playing a role, while the methodologies used differ substantially, thus casting doubts on the validity of the results obtained from their comparison[2].

[2] Reinert and Roland-Holst (1992) use distributed lags models with quarterly data not subject to analysis of the properties of the stochastic processes involved, while Gallaway *et al.* specify different models depending on the results obtained from the cointegration analysis previously performed for each sector.

The early studies by Alaouze, Marsden and Zeitsch (1977) and Alaouze (1977) provide additional evidence in favour of the existence of negative aggregation biases, since the models used are estimated using very similar methodologies and the same time periods for different sectoral aggregations. The magnitude of the estimators obtained is lower when using the most aggregated definition of goods. On the contrary, Gan (2006) shows that a negative aggregation bias should not always be expected, by means of analysing a single US industry using data at diverse levels of aggregation.

The fact that homogeneity is more easily guaranteed the more disaggregated the definition of goods used is unquestionable as the ease with which goods may be substituted should be larger the more homogeneous goods are. However, the current state of the art does not allow for assuring that just by using further disaggregated data much higher Armington elasticities are to be expected. Nonetheless, the evidence does signal at the key role of attaining consistency between the aggregation levels of all the variables included in the models.

Econometric Methodology

The above discussed methodological issues may be considered as less relevant when compared to the shortcomings found in most of the econometric analyses on Armington elasticities. Since the early 80s, most applied academics have stopped considering econometric modelling as a tool for quantifying theoretical relationships and started viewing the discipline as a different approach to economic analysis instead. As such, the information stemming from estimation, misspecification analysis and evaluation testing is taken as evidence signalling at the need to rethink both the estimable and the theoretical models proposed to approximate the phenomena of interest.

Models that have not been subject to exhaustive statistical testing are currently considered as non reliable and hence incapable of providing robust evidence supporting any hypothesis of interest. Further, whenever time series data are involved, the analysis of the order of integration of the stochastic processes and their cointegration is of utmost importance, since it has been shown that inference stemming from models including integrated but not cointegrated variables is not only non robust but generally misleading (Granger and Newbold, 1974).

Since the early 80s a vast literature has emerged providing tests and estimation procedures that do take into account the non stationarity of data and allow for a correct econometric analysis of these processes, while a large number of additional tests have been proposed for performing an exhaustive testing of the statistical assumptions underlying the different types of econometric models. Most of them are even included in any standard econometric software. Hence, although the applied work on Armington elasticities that was developed prior to the 90s may suffer from statistical misspecification, work that has been done afterwards should not.

In Table 1 we summarise the characteristics of many of the papers that perform time series analyses in the last two decades without properly accounting for the dynamic properties of the processes included in the models specified. We also report the statistical tests performed in each case in order to validate the models and the inference based on them.

A first striking observation refers to the lack of an exhaustive statistical evaluation of the underlying assumptions in most models. Assumptions such as normality or functional form are rarely tested for, while autocorrelation and heteroskedasticity are generally poorly tested for in the papers that do report a few misspecification tests.

Most striking still is the fact that many papers do discuss the crucial role of using a reduced form model but they do not test for the endogeneity of prices, with the notable exceptions of Ivanova (2005) and Saito (2004). A further exception is Shiells, *et al.* (1986) who specify a simultaneous supply and demand model from the very start and thus estimate the reduced form by 2-Stages Least Squares.

Interesting enough is Ivanova's result related to the magnitude of the estimated Armington elasticities being higher when using Instrumental Variables methods than when estimating by Least Squares, revealing the eventually key role of performing misspecification tests.

A related topic that has not been widely analysed refers to the existence of contemporaneous correlation among unobservable components of the models specified for several single goods. Whenever exogenous shocks to one market have also effects on other related markets, there are gains in precision if estimating jointly the relative demand of the diverse goods. The paper by Shiells, *et al.* (1986) constitutes again an exception to the rule.

Regarding the by now unavoidable analysis of the order of integration of the stochastic processes included in the models, the existence of a balanced relationship and of cointegration when not all variables are stationary, the review is also discouraging. Many papers do not even address the issue. All the other papers here reviewed that do test for stationarity, once integrated variables are found pursue specification and estimation strategies that are at best non optimum while they are senseless at times. The only exceptions are Saito (2004) and Welsch (2006).

It is well known that inference performed from estimated models involving integrated variables is misleading due to the standard test statistics not having a known distribution while the power of tests is lower and the size unknown (Granger, 1981).

One possible strategy is to differentiate the integrated time series in order to achieve stationarity and thus use the standard estimation and testing techniques. However, in doing so the long run information contained in the data is lost. Needless to say that the so transformed model will be equivalent to the originally specified only whenever all variables are differentiated an equal number of times, although this may in turn generate a unit root in the MA representation of the series due to over-differentiation, a consequence that is not analysed nor mentioned in any of the papers following this path.

Table 1: Main characteristics of econometric models estimated in the applied literature

Author/s	Year	Model/s	Estimation method	Misspecification tests
Alaouze, Marsden & Zeitsch	1977	Four and three linear squares sliding trend projection, with Partial Adjustment Model (PAM) and Rapid Adjustment Model (RAM)	OLS	Durbin-Watson/ h-Durbin in PAM
Alaouze	1977	Four and three linear squares sliding trend projections, with PAM and RAM	OLS	Durbin-Watson/ h-Durbin in PAM
Shiells, Stern and Deardorff	1986	PAM with stock adjustment mechanism for imports.	2SLS/ 2SLS + Cochrane-Orcutt/ 3SLS	Omitted variables
Corado & de Melo	1986	Static nested model, with and without lags in the price variable	OLS	Durbin-Watson
Reinert & Shiells	1991	Nested and non nested Armington models	Restricted SUR corrected for AR	Non reported
Reinert & Roland-Holst	1992	Static Armington model. PAM and DLM (non reported)	OLS + Cochrane-Orcutt	Durbin-Watson
Shiells & Reinert	1993	Nested and non nested Armington models	ML with AR(1) error term	Non reported
Hernández	1998	Static Armington model. PAM	OLS	Chow/ Breusch/Godfrey/ RESET
Blonigen & Wilson	1999	Static Armington model	OLS + Cochrane-Orcutt	Non reported
Kapuscinski & Warr	1999	Static Armington model. PAM. ECM	OLS	ML AR(1) test/ RESET
Le Roux Burrows	1999	Static Armington model	OLS with AR error term	Non reported
Gallaway, McDaniel & Rivera	,2003	Dynamic Armington models in levels or differences according to: I(0)-I(0): DGLM in levels/C(1,1): ECM/ Non C(m,n)/Unbalanced: Static models in differences	OLS + White HS correction	Durbin-Watson
Bilgic, King, Lusby & Schreiner	2002	Cross section Armington model	OLS or GLS	Non reported
López & Pagoulatos	2002	Static Armington model	OLS	Durbin-Watson
Gibson	2003	Dynamic Armington models in levels or differences according to: I(0)-I(0): DGLM in levels/ C(1,1): ECM/ Non C(m,n)/Unbalanced: Static models in differences	OLS/GLS-HCSE estimator/GLS-Cochrane-Orcutt	Normality Shapiro W Test Wilk / Cook-Weisberg Heterosckedasticity Test/Durbin-Watson/ Non parametric RUNS Test of Geary/ Breusch-Godfrey Test/RESET
Lozano	2004	Armington ARDL(m,n). The order (m,n) is at the most 4 and selected with Akaike criterion. Cointegration is found in all the cases.	OLS	ML Serial Correlation/Un specified tests for functional form; normality and heterosckedasticity
Saito	2004	Armington bilateral and multilateral models	Fully Modified OLS	Non reported
González and Wong	2005	Dynamic Armington models in levels or differences according to: I(0)-I(0): DGLM in levels/ C(1,1): ECM/ Non C(m,n)/Unbalanced: Static models in differences	OLS	Unspecified tests for heteroskedasticity and serial correlation
Ivanova	2005	LA-AIDS	2SLS+HCSE/G2SLS/IV+Fixed Effects	Exogeneity/Durbin-Wu-Hausman/ Overidentifying restrictions Test
Gan	2006	Dynamic Armington models in levels or differences according to: I(0)-I(0): DGLM in levels/ C(1,1): ECM/ Non C(m,n)/Unbalanced: Static models in differences/ Linear variable parameters model	OLS/ FGLS	Durbin-Watson/ h-Durbin/ Quandt - structural breaks.
Welsch	2006	Dynamic Armington models in levels or differences according to: C(1,1): Johansen/ Non C(m,n) or unbalanced: Discarded	Johansen	Non reported
Fontes, Kume and de Souza Pedroso	2007	Armington according to dynamic properties of the series: I(0)-I(0):Static model in levels/ C(1,1): VEC/ Unbalanced: if q I(1) and p I(0):First differences static model. If q is I(0) - p is I(1): Static model in levels + lags of prices	OLS/ ML with AR error term	Durbin-Watson

An even more worrisome finding in several papers here reviewed is them neglecting the information that the order of integration analyses provide relative to the specification of the models themselves. Once the order of integration of the processes involved has been determined, the existence of balance in the proposed models should be verified so as not to estimate statistically nonsensical relations – as would be the case when an I(1) stochastic process is assumed to be explained by I(0) processes or *vice versa*. However, unbalanced models are estimated and analysed in many of the papers in Table 1. In spite of the fact that the variables are transformed into stationary processes before estimation, in doing so the economic information stemming from the unbalanced order of integration of the phenomena is not being considered as a sign of models being in need of respecification.

Finally, not all the here reviewed papers that study the order of integration of the variables included in the models analyse the existence of cointegration for those involving integrated variables. Even with a balanced model, the only sensible strategy that has to be followed whenever cointegration is rejected is the respecification of the original relation. Although differentiating the time series would allow for the use of the statistical tools suitable for stationary processes, the rejection of the existence of cointegration implies that the economic relation that is being proposed lacks sense. It may otherwise be suggesting that there are other variables that need to be taken into account for a proper understanding of the phenomenon of interest. These variables may be related to institutional features that have a role only in the particular case study under analysis, as would be the case of non tariff trade barriers, or even aspects such as the degree of unionisation. Moreover, it may also be the case that the inexistence of cointegration is signalling at the inadequate maintenance of assumptions, such as a unity income elasticity of demand that would imply the incorrect omission of a variable accounting for the activity level.

Disregarding this key information prevents attaining a proper understanding of the relations under analysis. Consequently, the parameters that are being estimated cannot be taken as representative of theoretically meaningful concepts. This is the case of many of the above reported models. In other cases, the use of a proper procedure is dimmed by the fact that observations are not enough or rejection regions are not well defined (Kapuscinski and Warr, 1999; Lozano, 2004; Warr, 2005). The last two papers cited do not test for cointegration directly but estimate static and error corrections versions of the models and argue that the resulting estimates support its existence. Looking at the reported results, however, the correctness of the procedures followed is not guaranteed, due both to the use of inadequate critical values (taken from DF tables instead of MacKinnon (1991) cointegration response surfaces) and to the misspecification of the models, as stemming from the results of the tests reported.

All the other papers listed in Table 1 do not address the issue at all while some of them are unable to do it due to the scarce temporal observations used (Ivanova, 2005, e.g.) or else do not need to account for the issue (Hummels, 1999; Bilgic *et al.*, 2002).

The brief review of the econometric analyses performed in the recent applied literature on Armington elasticities reveals that there is a poor and insufficient statistical evaluation of the estimated models. As such, many of the conclusions from them drawn are subject to discussion due to the inference based on them lacking enough robustness. Particularly, discussing the estimated magnitude of the models' parameters is out of place since even their statistical significance cannot be guaranteed.

Further, regarding models using time series data, it is absolutely necessary to perform meaningful analyses on the properties of the stochastic processes involved, something that is feasible only if the information provided by the results of the testing procedures performed is understood and taken into account for a correct respecification of the models.

Consequently, it is highly recommended to revise the debate on which are the most recommended strategies to obtain adequate estimated values of Armington elasticities in light of results obtained from models that should be previously proven to be adequate statistical representations of the relationship between relative demands and prices.

Estimations for Uruguay

As a means of giving some empirical support to some of the statements here sustained, we perform an exercise using Uruguayan data for the period 1989 to 2001.

In order to test for the implications of using cross section and/or time series data on the value of estimated Armington elasticities, we estimate uniequational time series models for manufacturing industries in the *Food and Beverages* division and we also estimate a multivariate model including the same industries. We perform the multivariate analysis both imposing an identical elasticity for all sectors – equivalent to using pooled cross section-time series data - and the unrestricted version as well, so as to allow for the existence of correlation among contemporaneous unobservable components.

We further estimate two versions of the unrestricted model, allowing for correlations to be present among 4-digit ISIC industries belonging to the same 3-digit grouping, as well as for all industries in the 2-digit division[3]. The results are summarised in Table 2.

As it is readily seen, the unrestricted estimates from the multivariate model are very similar if not identical to those stemming from the uniequational models, while precision is significantly gained (20% on average).

The magnitudes of the estimated elasticities using pooled cross section-time series data are indeed different, as they are in a way equivalent to a weighted average of the parameters estimated by the uniequational models. As such, they are smaller than the largest elasticities but larger than the smallest estimates. The obtained results are consistent with those suggested by Theil (1954).

[3] A list of the industries included in the analysis is provided in Annex 2.

Table 2 - Armington elasticities using cross section and time series data

	Uniequational Models		Unrestricted Multivariate Models		Restricted 3-digit Multivariate Models		Restricted 2-digit Multivariate Model	
Industry	Armington Elasticity	SD	Armington Elasticity	SD	Armington Elasticity	SD	Armington Elasticity	SD
3111	**1.53**	0.22	**1.53**	0.18	**0.88**	0.05	**0.97**	0.04
3112	**0.76**	0.24	**0.76**	0.19				
3113	**1.96**	0.25	**1.96**	0.20				
3115	**0.45**	0.12	**0.43**	0.09				
3116	**0.51**	0.23	**0.51**	0.18				
3117	**0.98**	0.18	**0.98**	0.14				
3118	**0.49**	0.18	**0.49**	0.13				
3119	**0.96**	0.16	**0.96**	0.12				
3121	**0.92**	0.17	**0.92**	0.13	**1.00**	0.07		
3122	**1.02**	0.09	**1.03**	0.08				
3131	**0.83**	0.14	**0.83**	0.11	**2.00**	0.24		
3132	**1.17**	0.57	**1.17**	0.45				
3133	**2.46**	0.34	**2.46**	0.26				
3134	**1.13**	0.45	**1.13**	0.39				

Note: In Annex 1 misspecification tests are reported. The full results of estimation are available upon request.

No clear-cut effects of the temporal frequency of data on the magnitude of Armington elasticities are here identified. For some industries differences are not significant while they do for others. However, eventual biases would both be positive and negative (Table 3).

Table 3 - Armington elasticities using monthly and quarterly data

	Monthly data			Quarterly data		
Industry	Starting date	Armington Elasticity	SD	Starting date	Armington Elasticity	SD
3111	1989.02	**1.95**	0.3	1992.1	**1.53**	0.22
3112	1991.01	**0.46**	0.21	1992.1	**0.76**	0.24
3113	1989.02	**1.46**	0.6	1992.1	**1.96**	0.25
3115	1990.06	**1.36**	0.24	1992.1	**0.45**	0.12
3116	1994.01	**1.56**	0.4	1992.1	**0.51**	0.23
3117	1994.01	**0.64**	0.08	1992.1	**0.98**	0.18
3118	1993.01	**0.96**	0.26	1992.1	**0.49**	0.18
3119	1993.01	**1**	0.1	1992.1	**0.96**	0.16
3121	1995.01	**0.37**	0.07	1992.1	**0.92**	0.17
3122	1989.03	**0.95**	0.2	1992.1	**1.02**	0.09
3131	1989.05	**1.25**	0.21	1992.1	**0.83**	0.14
3132	1989.03	**0.71**	0.21	1992.1	**1.17**	0.57
3133	1989.03	**2.98**	0.46	1992.1	**2.46**	0.34
3134	1992.01	**1.13**	0.36	1992.1	**1.13**	0.45

Note: In Annex 1 misspecification tests are reported. The full results of estimation are available upon request.

The third issue relates to the impact that using different levels of aggregation of the data has on the estimated values of the elasticities of substitution. The exercise was performed estimating models using data at 4, 3 and 2 digits level of the ISIC. In doing so we are assuming that there is homogeneity among national varieties of goods belonging either within industries such as *Slaughtering, preparing and preserving meat*; *Manufacture of dairy products*; *Canning and preserving of fruits and vegetables;* etc. or between groupings such as *Manufacture of Food products*; or *Beverage industries* ; or among all goods included in the division *Food and Beverages*.

The assumption is thus quite reliable for most goods in industries defined at the 4-digit level while quite unsustainable when using 2-digit level economic sectors. In spite of this fact, the results obtained are not as straightforward as to assess they fully support the statement, as shown in Tables 4 and 5 below.

In Table 4 the results of the uniequational models estimated for each 4-digit level industry within the division *Food and Beverages* are reported. Similarly, elasticities estimated at the 3-digit aggregation level – including *Manufacture of Food products; Other Food products;* and *Beverage industries* – and for the whole division are also detailed.

Table 4 - Armington elasticities using different levels of aggregation of data
Uniequational models

Industry	Armington Elasticity	SD	Industry	Armington Elasticity	SD	Industry	Armington Elasticity	SD
3111	**1.53**	0.22	**311**	**1.17**	0.46	**31**	**0.85**	0.22
3112	**0.76**	0.24						
3113	**1.96**	0.25						
3115	**0.45**	0.12						
3116	**0.51**	0.23						
3117	**0.98**	0.18						
3118	**0.49**	0.18						
3119	**0.96**	0.16						
3121	**0.92**	0.17	**312**	**0.63**	0.14			
3122	**1.02**	0.09						
3131	**0.83**	0.14	**313**	**1.10**	0.18			
3132	**1.17**	0.57						
3133	**2.46**	0.34						
3134	**1.13**	0.45						

Note: In Annex 1 misspecification tests are reported. The full results of estimation are available upon request.

In Table 5 the results of the analogous exercise using a multivariate model are shown. Contemporaneous correlation is found to be significant among the 4-digit aggregation level industries included in each 3-digit group, as well as among the economic sectors defined at a 3-digit level of aggregation belonging to the division (2-digit aggregation level).

Table 5 - Armington elasticities using different levels of aggregation of data
Multivariate models

Industry	Armington Elasticity	SD	Industry	Armington Elasticity	SD
3111	**1.53**	0.18	**311**	**1.17**	0.29
3112	**0.76**	0.19			
3113	**1.96**	0.20			
3115	**0.43**	0.09			
3116	**0.51**	0.18			
3117	**0.98**	0.14			
3118	**0.49**	0.13			
3119	**0.96**	0.12			
3121	**0.92**	0.13	**312**	**0.63**	0.10
3122	**1.03**	0.08			
3131	**0.83**	0.11	**313**	**1.10**	0.15
3132	**1.17**	0.45			
3133	**2.46**	0.26			
3134	**1.13**	0.39			

Note: In Annex 1 misspecification tests are reported.
The full results of estimation are available upon request.

The magnitude of the estimated elasticity for *Manufacture of Food products* (group 311) is apparently just reproducing the average behaviour of the 4-digit industries belonging to the group (3111 to 3119) if considering the relative weight in total production of those 4-digit level industries with the highest elasticities (more than 40%). The same can be said for the relative magnitude of the estimates at the 3-digit level as compared with the estimated elasticity using 2-digit level data.

However, the above is not observed in *Other Food products* for which the aggregated elasticity is lower than the disaggregated estimates, while the result is unclear in the case of *Beverages*.

The above figures are probably reflecting two facts. First, it is more plausible to observe the theoretical prediction when aggregating goods of a very heterogeneous character. This is the case of the 312 grouping, conformed by unspecified food products and prepared animal feeds, for which the aggregated elasticity is lower than its disaggregated counterparts. This may be so as it is most likely that the ease with which consumers substitute national by foreign varieties of animal feeds is different from that with which they substitute unspecified food of different origin of production.

On the contrary, when looking at *Manufacture of Food products*, the aggregated estimate is statistically equal to most of the disaggregated elasticities. Although the included goods are not homogeneous there exists a certain degree of substitutability among them, as consumers may well substitute red meat by fish, or else chocolate by bakery products. As such, to assume that they substitute foreign by national varieties of each of these goods with a similar ease is quite sensible. Consequently, aggregation may lose relevance in such a framework.

Second, the results also suggest that there may be other dimensions that are being disregarded in the analyses. One of them refers to the role played by the presence of multinational firms that may be underlying the mixed results obtained for the *Beverages* industry. The aggregated estimate is smaller in value for three of the four disaggregated elasticities but it is statistically equal to the fourth. Actually, the aggregate estimate is statistically smaller only to that obtained for the Beer industry (3133). It may be argued that the point estimate result is in line with the role of heterogeneity among the included merchandises above stated, as consumers need not substitute with ease sodas by wine. However, it is also likely that although goods are not homogeneous, the ease with which consumers may substitute imported varieties by national products is similar among all beverages with the exception of beers, so that aggregation *per se* should not be expected to reduce the elasticity except in that sole case.

The above hypothesis might be linked to the concentration process taking place worldwide in the industry, so that in the 90s multinational trusts increasingly acquired Uruguayan local firms producing alcoholic and non alcoholic beverages. Consequently, coordination in marketing strategies allowed for increasingly ignoring the preferences of local consumers in each submarket, and hence to partially eliminate eventual differences in the ease of substitution between national and foreign varieties within each type of beverage. In this context it is quite sensible to obtain similar Armington elasticities using distinct levels of aggregation although the goods are intrinsically differentiated.

Despite the wine industry partially escapes this pattern of consumption, its share is too small to have an incidence in the average estimates. The concentration process in the Beer industry, on the contrary, started only by the end of the here analysed period. The fidelity of local consumers to brands has never been high in Uruguay so that even for imported beers, demanded only by a small portion of the market, the reaction to changes in relative prices is high. As a consequence, a reduction of the elasticity of substitution in this particular sub-market should be expected when pooled with other beverages.

A final issue relates to the econometric analyses themselves. The estimated elasticities above reported were obtained after a long respecification process, sustained in an exhaustive statistical evaluation of the models. The results of the misspecification tests performed signalled at diverse unaccounted properties of the joint conditional distribution as initially specified.

The dynamic structure was defined first by means of diverse tests, although in many cases it was further reduced when other omitted variables were included. A second step was the analysis of strong exogeneity and the consequent change of the estimation method whenever rejected. The existence of atypical observations, generally confounded with the rejection of normality of the conditional distribution or of the constancy of the variance, was accounted for with binary variables, thus avoiding an unnecessary respecification of higher conditional moments. It further pointed at the relevance of enduring the unavoidable task of building better quality and consistent datasets, so as to guarantee that the observable variables are indeed good proxies of the theoretical concepts.

The thorough analysis of the order of integration and cointegration whenever faced to balanced models shed light on diverse characteristics of the phenomena under analysis, such as them reacting differently when subject to shocks. The result may imply, e.g., different mechanisms of price determination exist, a fact that may help to identify reliable candidates for omitted variables when so diagnosed.

Finally, a significant number of equilibrium relations were not rejected at a low level of confidence (90%), a fact that also may be reflecting the need of including other phenomena in the initially stated long run behavioural relationships, or else that the relations themselves have changed at some point in time. This was indeed the case in a few of industries analysed although their results are not here reported (see Flores, 2008 for a review of all results).

Conclusions

We here argue that the debate focused on identifying the best strategy within the econometrics methodology that would allow for obtaining larger estimated values of the Armington elasticities than those currently available is misdirected. Some suggestions of generalised acceptance in the applied literature are here questioned due to them lacking enough robustness, and in some cases because of them being inadequate and/or misleading.

Issues such as the temporal frequency of the data and their spatial or temporal character cannot be considered *per se* as responsible of the so considered underestimation of the elasticity. Instead, they are revealing the key role that the available proxy variables are playing due to their measurement exhibiting major shortcomings. They are also signalling at the incorrect specification of models in terms of what they are intended to capture.

Regarding aggregation levels, the discussion could be a lot more fertile if redirected towards building more sophisticated models and/or relaxing some restrictive assumptions, while trying to identify crucial features explaining some of the differences in the dynamic structure of the processes involved. Further, the temporal non constancy of the Armington elasticities needs further study, a fact that is rarely tested for in the existent models.

Our main conclusion though refers to the need of reversing the generalised disregard of the information provided by each necessary stage undergone in the econometric analyses, at least as reported in the reviewed literature on Armington elasticities. This is most notable when time series data are involved. Only by studying in depth the characteristics of the included stochastic processes, both individually and jointly, and by understanding the economic meaning of the results stemming from misspecification tests, will the specified models be guaranteed enough robustness so as to use their outcomes in policy analyses with the necessary confidence levels.

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Annex 1 - Misspecificaton tests
Multivariate models 4 & 3-digit level – Uniequational model 2-digit level

Industry	**RESET** (General Error in Conditional Mean)	**Jarque-Bera** (Normality)	**Breusch y Godfrey** (Autocorrelation)	**White** (Heteroskedasticity)	**ARCH** (Autocorrelated Conditional Variance)	**Incorrect Omission:** Exchange rate	**Incorrect Omission:** Tariffs
3111	0.49	2.08	1.83	1.80	0.05	0.87	0.15
3112	0.04	0.74	1.92	0.64	0.81	0.25	-0.15
3113	0.60	0.13	1.03	0.85	0.03	0.00	0.05
3115	1.38	1.25	-0.15	0.71	0.88	0.99	0.66
3116	-0.06	1.43	-1.05	0.40	0.69	-0.12	0.81
3117	0.31	0.85	0.16	1.14	0.44	0.05	0.29
3118	-0.51	0.99	0.11	1.70	0.50	1.79	-0.87
3119	0.54	0.59	0.23	0.85	0.54	0.04	0.59
3121	0.15	0.08	-0.41	0.45	2.51	-0.52	-0.83
3122	1.24	1.33	2.64	0.90	0.00	0.85	0.06
3131	0.19	0.32	0.20	1.05	1.48	0.14	0.51
3132	-0.13	5.11 /1	0.04	0.78	1.37	0.80	0.77
3133	0.74	0.03	-1.50	1.37	2.17	-1.68	-1.68
3134	-0.37	0.62	0.07	1.26	0.72	-1.95	NA
311	-0.56	0.59	1.17	0.39	1.49	-0.21	-0.31
312	0.46	2.3	0.54	0.83	0.18	-0.48	-1.23
313	-0.83	0.82	-1.24	0.46	0.26	0.31	0.24
31	1.47	0.65	0.69	1.07	4.17	0.17	0.20

Notes: /1 Normality is rejected due to the presence of too many atypical observations. NA: Not Applicable.

Exogeneity – Hausman Test

Industry	Variable	b_LS	b_IV	q	m	
3111	**R.PRICE**	1.202	1.174	-0.028	**8.044**	***
3112	**R.PRICE**	0.441	0.411	-0.029	**62.269**	***
3113	**R.PRICE**	1.402	1.392	-0.010	**1.270**	
3115	**R.PRICE**	0.278	0.286	0.008	**5.253**	**
3116	**R.PRICE**	0.428	0.406	-0.023	**9.824**	***
3117	**R.PRICE**	0.451	0.455	0.004	**1.264**	
3118	**R.PRICE**	0.465	0.430	-0.036	**45.374**	***
3119	**R.PRICE**	0.625	0.614	-0.012	**2.544**	
3121	**R.PRICE**	0.687	0.704	0.017	**6.404**	**
3122	**R.PRICE**	0.852	0.852	0.000	**0.000**	
3131	**R.PRICE**	0.785	0.732	-0.053	**17.608**	***
3132	**R.PRICE**	0.572	0.610	0.038	**66.842**	***
3133	**R.PRICE**	1.409	1.430	0.021	**8.434**	***
3134	**R.PRICE**	0.651	0.693	0.042	**17.369**	***
	TARIFF	-10.180	-14.553	-4.373	**115.84**	***
311	**R.PRICE**	0.188	0.219	0.032	**59.710**	***
312	**R.PRICE**	0.447	0.462	0.015	**6.928**	***
313	**R.PRICE**	0.688	0.744	0.056	**76.682**	***
31	**R.PRICE**	0.219	0.216	-0.003	**0.302**	

Note: b_LS is the Least Square estímate; b_IV is the Instrumental Variables estimator; q is the difference between b_LS and b_IV; m is the Hausman statistic. * 10% level of significance; ** 5% level of significance; *** 1% level of significance.

Annex 2 - International Standard Industrial Classification (Rev. 2)

Divisions, Major Groups and Groups

|

31 Manufacture of Food, Beverages and Tobacco

311 Food manufacturing

- **3111** Slaughtering, preparing and preserving meat
- **3112** Manufacture of dairy products
- **3113** Canning and preserving of fruits and vegetables
- **3114** Canning, preserving and processing of fish, crustacea and similar food
- **3115** Manufacture of vegetable and animal oils and fats
- **3116** Grain mill products
- **3117** Manufacture of bakery products
- **3118** Ingenios y refinerías de azúcar
- **3119** Manufacture of cocoa, chocolate and sugar confectionery

312 Other food products

- **3121** Manufacture of food products not elsewhere classified
- **3122** Manufacture of prepared animal feeds

313 Beverage industries

- **3131** Distilling, rectifying and blending spirits
- **3132** Wine industries
- **3133** Malt liquors and malt
- **3134** Soft drinks and carbonated waters industries