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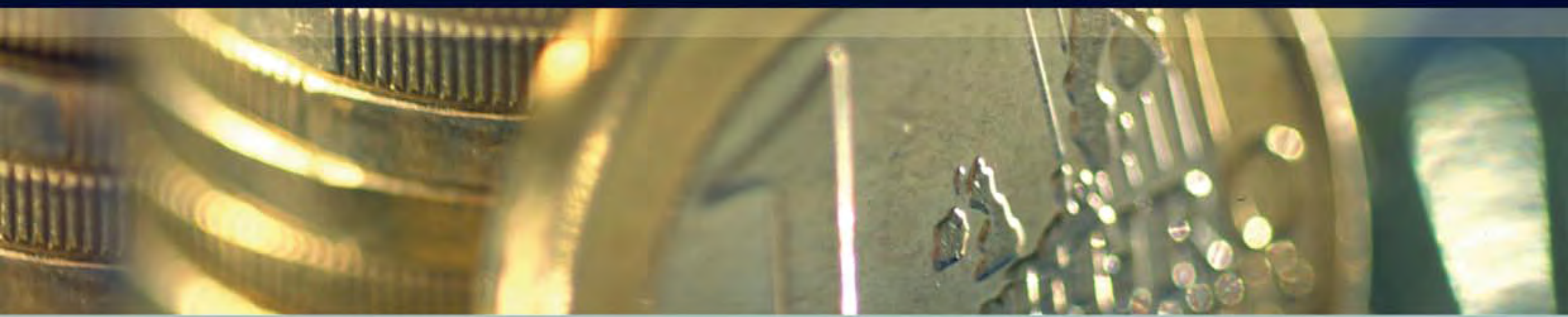

Trade Elasticities
A Final Report for the European Commission

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Trade Elasticities

A Final Report for the European Commission*

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Abstract

In a demand system with conventional CES preferences, the price elasticitites of aggregate trade flows are weighted averages of sector-specific elasticities of substitution. We describe a methodology that can be used to estimate country-specific values for the price elasticities of aggregate imports and exports. We first use disaggregated trade data to compute structural estimates of international substitutability for a large cross section of countries. We aggregate up the estimates using model-implied, country-specific weights. We obtain structural estimates of the price elasticities of aggregate exports and imports for more than 30 countries, including most developed and developing economies.

JEL Classification Numbers: F32, F02, F15, F41

Keywords: Price Elasticity of Exports, Price Elasticity of Imports.

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1 Introduction

How responsive are traded quantities to a shift in international prices? The answer is of direct relevance to a broad range of past and current issues in international economics. Export elasticities are often invoked to illustrate the relative resilience of certain exporters in the face of a sudden deterioration in their competitive position. The price elasticity of imports, in turn, reflects consumers' fidelity to domestic or foreign goods. And the price elasticity of net exports determines directly whether the venerable Marshall-Lerner condition is verified, and favorable shifts in relative prices have positive end effects on the trade balance. In fact, the price elasticities of trade govern the dynamics of the trade balance, the J-curve discussed in Backus, Kehoe and Kydland (1994).

In a conventional demand system with Constant Elasticity of Substitution (CES) utility, there is a close mapping between substitutability in preferences and trade elasticities. The price elasticity of imports depends linearly on the preference parameters of the importing representative consumer. The price elasticity of exports, in turn, is given by a weighted average of preference parameters across exports destination markets. At least in a conventional CES demand system, there is a tight link between stuctural parameters and reduced form price elasticites.

In this paper, we describe and implement a structural methodology to estimate the price elasticities of imports and exports. The approach builds on a multi-country demand system with nested CES preferences. For each country, import price elasticities depend on the domestic willingness to substitute domestic and foreign varieties, aggregated across sectors. Export price elasticities are given by a similarly weighted average of substitutability, now aggregated across both destination markets and sectors. All that is needed to pin down trade elasticities are cross-country estimates of the elasticities of substitution - and the weights used in aggregation. In Imbs and Méjean (2009), we adapted Feenstra's (1994) technology to obtain structural estimates of the elasticity of substitution in the US. Here we show how this can be extended to a multi-country framework. We use the setup to compute structural estimates of trade elasticities for a cross-section of more than 30 countries, including most developed and developing economies.

The rest of the paper is structured as follows. Section 2 develops the model that relates trade and substitution elasticities. Section 3 discusses our structural estimation of substitution elasticities across countries, along with the data we use. Our results are presented in Section 4, and Section 5 concludes.

2 Theory

We build on a Constant Elasticity of Substitution (CES) demand system, with two layers of aggregation. Aggregate consumption is a CES aggregate of sectors indexed by $k = 1, ..., K$. Each sector, in turn, is a CES index of varieties $j \in I_{kj}$ that can be produced either at home or abroad. Consumption in country j is given by

$$C_j = \left[\sum_{k \in K_j} (\alpha_{kj} C_{kj})^{\frac{\gamma_j - 1}{\gamma_j}} \right]^{\frac{\gamma_j}{\gamma_j - 1}}$$

where α_{kj} denotes an exogenous preference parameter and γ_j the elasticity of substitution between sectors in country j. Consumption in each sector is derived from a range of varieties of good k, that may be imported or not, as in

$$C_{kj} = \left[\sum_{i \in I_{kj}} \left(\beta_{kij} \ C_{kij}\right)^{\frac{\sigma_{kj} - 1}{\sigma_{kj}}} \right]^{\frac{\sigma_{kj}}{\sigma_{kj} - 1}}$$

Here $i \in I_{kj}$ indexes varieties of good k, produced in country i and consumed by country j. We let the elasticity of substitution σ_{kj} be heterogeneous across industries and importing countries. β_{kij} lets preferences vary exogenously across varieties, reflecting for instance differences in quality or a home bias in consumption.

The representative maximizing agent chooses consumption keeping in mind that all varieties incur a transport cost τ_{kij}. In the trade literature, the transport cost is usually assumed to be zero for domestically produced goods, i.e. $\tau_{kjj} = 1$. Utility maximization implies that demand for variety i in each sector k is given by

$$C_{kij} = \beta_{kij}^{\sigma_{kj} - 1} \left(\frac{P_{kij}}{P_{kj}} \right)^{1 - \sigma_{kj}} \frac{1}{P_{kij}} \alpha_{kj}^{\gamma_j - 1} \left(\frac{P_{kj}}{P_j} \right)^{1 - \gamma_j} P_j C_j \qquad (1)$$

with

$$
\begin{aligned}
P_{kij} &= \tau_{kij} P_{kij}^{fob} \\
P_{kj} &= \left[\sum_{i \in I_{kj}} \left(\frac{P_{kij}}{\beta_{kij}} \right)^{1-\sigma_{kj}} \right]^{\frac{1}{1-\sigma_{kj}}} \\
P_j &= \left[\sum_{k \in K_j} \left(\frac{P_{kj}}{\alpha_{kj}} \right)^{1-\gamma_j} \right]^{\frac{1}{1-\gamma_j}}
\end{aligned}
$$

where P_{kij}^{fob} is the Free On Board (FOB) price of variety i. Without loss of generality, we assume FOB prices are expressed in the importer's currency.

We now ask our model how aggregate quantities respond to changes in aggregate international relative prices. Our aim here is to replicate the assumptions that underly conventional estimations of trade elasticities. In macroeconomics, the conventional estimated regressions write

$$
\begin{aligned}
\ln M_{jt} &= \eta_j^M \ln \left(\frac{P_{jt}^M}{P_{jt}} \right) + Y_{jt} \\
\ln X_{jt} &= \eta_j^X \ln \left(\frac{P_{jt}^X}{P_t^*} \right) + Z_{jt}
\end{aligned}
$$

where M_{jt} and X_{jt} measure aggregate imports and exports by country j, $\frac{P_{jt}^M}{P_{jt}}$ is the price of imports relative to domestic prices in country j, $\frac{P_{jt}^X}{P_t^*}$ is the price of exports relative to prices in the destination country, denoted with a *, and Y_{jt} and Z_{jt} are country-specific controls. η_j^M (η_j^X) measures the response of imports (exports) to a shock in the price of imports (exports), relative to the price of the competing goods. Both are expected to be negative.

In what follows we use these definitions of trade elasticities and compute the response of trade to a shock affecting all relative prices in country j, across all sectors k. Let η_{kj}^M (η_{kj}^X) denote the response of country j's sectoral imports (exports) to the shock, while η_j^M (η_j^X) is the aggregate response of imports (exports). By definition:

$$
\begin{aligned}
\eta^M_{kj} &= \frac{\partial \ln \sum_{i\neq j} \ P_{kij}C_{kij}}{\partial \ln\{P_{kij}/P_{kjj}\}_{\forall k, i\neq j}} \\
\eta^X_{kj} &= \frac{\partial \ln \sum_{i\neq j} \ P_{kji}C_{kji}}{\partial \ln\{P_{kji}/P_{kii}\}_{\forall k, i\neq j}} \\
\eta^M_{j} &= \frac{\partial \ln \sum_k \sum_{i\neq j} \ P_{kij}C_{kij}}{\partial \ln\{P_{kij}/P_{kjj}\}_{\forall k, i\neq j}} = \sum_k m_{kj}\eta^M_{kj} \\
\eta^X_{j} &= \frac{\partial \ln \sum_k \sum_{i\neq j} \ P_{kji}C_{kji}}{\partial \ln\{P_{kji}/P_{kii}\}_{\forall k, i\neq j}} = \sum_k x_{kj}\eta^X_{kj}
\end{aligned}
$$

Aggregate trade elasticities are weighted averages of the corresponding sectoral elasticities, with weights given by

$$
m_{kj} = \frac{\sum_{i\neq j} P_{kij}C_{kij}}{\sum_k \sum_{i\neq j} P_{kij}C_{kij}}
$$

the value share of sector k in j's aggregate imports and

$$
x_{kj} = \frac{\sum_{i\neq j} P_{kji}C_{kji}}{\sum_k \sum_{i\neq j} P_{kji}C_{kji}}
$$

the value share of k in j's aggregate exports. These definitions capture the response of trade in *value*. By definition, the corresponding response of traded *volumes* is obtained by substracting 1 to η^M_{kj} and η^X_{kj}. This abstracts from the direct impact of the price shock on the value of trade.

Consider first the response of sectoral imports. Using equation (1), simple algebra implies

$$
\begin{aligned}
\eta^M_{jk} &= \sum_{i\neq j} m_{kij}\left[(1-\sigma_{kj})\frac{\partial \ln P_{kij}/P_{kjj}}{\partial \ln P_{kij}/P_{kjj}} + (\sigma_{kj}-\gamma_j)\frac{\partial \ln P_{kj}/P_{kjj}}{\partial \ln P_{kij}/P_{kjj}} + (\gamma_j - 1)\frac{\partial \ln P_{j}/P_{kjj}}{\partial \ln P_{kij}/P_{kjj}}\right] \\
&= (1-\sigma_{kj}) + (1-w_{kjj})(\sigma_{kj}-\gamma_j) + (\gamma_j-1)\sum_k w_{kj}(1-w_{kjj}) \qquad (2)
\end{aligned}
$$

with

$$
m_{kij} \equiv \frac{P_{kij}\ C_{kij}}{\sum_{i\neq j} P_{kij}\ C_{kij}}
$$

the share of variety i in country j's imports of product k,

$$w_{kjj} \equiv \frac{P_{kjj}\ C_{kjj}}{\sum_i P_{kij}\ C_{kij}}$$

the share of domestic goods in country j's nominal consumption of products k and

$$w_{kj} \equiv \frac{P_{kj}\ C_{kj}}{P_j\ C_j}$$

the share of good k in country j's nominal consumption.

Using equation (2), the aggregate elasticity of imports in country j becomes:

$$\eta_j^M = \sum_k m_{kj}(1-\sigma_{kj}) + \sum_k m_{kj}(1-w_{kjj})(\sigma_{kj}-\gamma_j) + (\gamma_j - 1)\sum_k w_{kj}(1-w_{kjj}) \tag{3}$$

The response of aggregate imports to a shock in international prices is given by an adequately weighted average of σ_{kj}, the elasticity of substitution between varieties of good k in country j. With structural estimates of σ_{kj}, and calibrated values for m_{kj}, w_{kjj} and w_{kj}, equation (3) implies a semi-structural estimate of the price elasticity of imports.

Equation (3) has three elements. The first term involves an import-weighted average of σ_{kj}. The other two terms reflect the composition of sectoral trade. Note γ_j has a level effect on η_j^M, through the second and third summations in equation (3) . Both are likely smaller in magnitude than the first one. We later calibrate several values for γ_j, and show how the end value for η_j^M is affected.

The price elasticity of exports is more involved as it depends on the elasticities of substitution country j faces in *all* exporting destinations. By analogy, we use equation (1) to derive demand from country i addressed to producers in j, namely C_{kji}. Simple algebrae implies the sectoral elasticity of exports is given by

$$\eta^X_{jk} = \sum_{i\neq j} x_{kji}\left[(1-\sigma_{ki})\frac{\partial \ln P_{kji}/P_{kii}}{\partial \ln P_{kji}/P_{kii}} + (\sigma_{ki}-\gamma_i)\frac{\partial \ln P_{ki}/P_{kii}}{\partial \ln P_{kji}/P_{kii}} + (\gamma_i - 1)\frac{\partial \ln P_i/P_{kii}}{\partial \ln P_{kji}/P_{kii}}\right]$$

$$= \sum_{i\neq j} x_{kji}\left[(1-\sigma_{ki}) + (\sigma_{ki}-\gamma_i)w_{kji} + (\gamma_i - 1)\sum_k w_{ki}w_{kji}\right] \quad (4)$$

where

$$x_{kji} = \frac{P_{kji}C_{kji}}{\sum_{i\neq j} P_{kji}C_{kji}}$$

is the share of country j's exports of product k sold in country i and

$$w_{kji} = \frac{P_{kji}C_{kji}}{\sum_l P_{kli}C_{kli}}$$

is the share of products from j in country i's consumption of k.

The aggregate price elasticity of exports writes

$$\eta^X_j = \sum_k x_{kj} \sum_{i\neq j} x_{kji}\left[(1-\sigma_{ki}) + (\sigma_{ki}-\gamma_i)w_{kji} + (\gamma_i - 1)\sum_k w_{ki}w_{kji}\right] \quad (5)$$

The price elasticity of exports is a weighted average of elasticities of substitution in destination markets. The weighting scheme involves both the share of each sector in overall exports, and the share of importing country i in j's exports. Equations (5) has three components: an adequately weighted average of σ_{ki}, and two terms that purely reflects the specialization of trade. These involve γ_i, which we calibrate, and that has level effects on η^X_j, albeit probably small.

Equations (3) and (5) demonstrate both aggregate import and export elasticities are weighted averages of sector-specific elasticities of substitution, σ_{ki}. All that is needed for estimates of η^X_j and η^M_j are therefore sector and country-specific estimates of the elasticity of substitution, and calibrated values for γ_i, x_{kij}, x_{kj}, m_{kj}, w_{kij}, and w_{ki}. We now turn to our proposed structural estimation of the preference parameter σ_{ki}, across sectors k and countries i.

3 Estimation and Data

3.1 Estimation

Identification at sectoral level across countries is achieved thanks to the multilateral dimension of disaggregated trade data. Following Feenstra (1994), we identify σ_{ki} using the cross-section of traded quantities and prices across exporters selling goods to each considered destination. We now describe our implementation of Feenstra's methodology.[1]

Demand is given in equation (1), which rewrites

$$C_{kijt} = \left(\frac{P_{kijt}}{P_{kjt}}\right)^{1-\sigma_{kj}} \frac{\beta_{kijt}^{\sigma_{kj}-1} P_{kjt} C_{kjt}}{P_{kijt}}$$

where t is a time index, and we used the fact that $P_{kjt}C_{kjt} = \alpha_{kj}^{\gamma_j - 1} \left(\frac{P_{kjt}}{P_{jt}}\right)^{1-\gamma_j} P_{jt}C_{jt}$. Following Feenstra (1994), impose a simple supply structure

$$P_{kijt} = \exp(\upsilon_{kijt}) C_{kijt}^{\omega_{kj}}$$

where υ_{kijt} denotes a technological shock that can take different values across sectors and exporters and ω_{kj} is the inverse of the price elasticity of supply in sector k.[2]

Define $s_{kijt} = \frac{P_{kijt}C_{kijt}}{P_{kjt}C_{kjt}}$ and rewrite demand as

$$s_{kijt} = \left(\frac{P_{kijt}}{P_{kjt}}\right)^{1-\sigma_{kj}} \beta_{kijt}^{\sigma_{kj}-1}$$

This defines expenditure shares, which tend to alleviate measurement error, following Kemp (1962). We do not observe domestically produced consumption. In addition, prices are measured Free on Board. Let tilded variables denote the observed counterparts to theory-implied prices and quantities. We observe

[1] The framework borrows from Imbs and Méjean (2009).

[2] Crucially, all exporters selling goods in a given market share the same supply elasticity by assumption.

$\tilde{P}_{kijt} \equiv P_{kijt}/\tau_{kijt}$. The empirical market shares are therefore given by

$$\tilde{s}_{kit} \equiv \frac{\tilde{P}_{kijt} C_{kijt}}{\sum_{i \neq j} \tilde{P}_{kijt} C_{kijt}} = \frac{s_{kijt}}{\tau_{kijt}} \left(1 + \frac{P_{kjjt} C_{kjjt}}{\sum_{i \neq j} \tilde{P}_{kijt} C_{kijt}} \right) \equiv \frac{s_{kijt}}{\tau_{kijt}} \mu_{kjt}$$

Taking logarithms, demand rewrites in first differences as

$$\Delta \ln \tilde{s}_{kijt} = (1 - \sigma_{kj}) \Delta \ln \tilde{P}_{kijt} + \Phi_{kjt} + \varepsilon_{kijt} \tag{6}$$

with $\Phi_{kjt} \equiv (\sigma_{kj} - 1)\Delta \ln P_{kjt} + \Delta \ln \mu_{kjt}$, a time-varying intercept common across all varieties, and $\varepsilon_{kijt} \equiv (\sigma_{kj} - 1)\Delta \ln \beta_{kijt} - \sigma_{kj} \Delta \ln \tau_{kijt}$ an error term that captures random trade cost and taste shocks, via changes in τ_{kijt} and β_{kijt}.

After rearranging, substituting in log-linearized supply yields

$$\Delta \ln \tilde{P}_{kijt} = \Psi_{kjt} + \frac{\omega_{kj}}{1 + \omega_{kj}\sigma_{kj}} \varepsilon_{kijt} + \delta_{kijt} \tag{7}$$

with $\Psi_{kjt} \equiv \frac{\omega_{kj}}{1+\omega_{kj}\sigma_{kj}} \left[\Phi_{kjt} + \Delta \ln \sum_i (\tilde{P}_{kijt} C_{kijt}) \right]$ a time-varying factor common across varieties, which subsumes sector specific prices and quantities. $\delta_{kijt} \equiv \frac{1}{1+\omega_{kj}\sigma_{kj}} \Delta \upsilon_{kijt}$ is an error term. Note that Ψ_{kjt} can also capture sector-specific inflation in country j, which may be present in prices, computed on the basis of nominal unit values.

Under standard assumptions on taste shocks β_{kijt} and technology shocks υ_{kijt} , it is possible to identify the system formed by equations (6) and (7). Identification rests on the cross-section of exporters i to the considered economy, and is achieved in relative terms with respect to a reference country r. The following estimable regression summarizes the information contained in the system:

$$Y_{kijt} = \theta_{1kj} X_{1kijt} + \theta_{2kj} X_{2kijt} + u_{kijt} \tag{8}$$

where $Y_{kijt} = (\Delta \ln \tilde{P}_{kijt} - \Delta \ln \tilde{P}_{krjt})^2$, $X_{1kijt} = (\Delta \ln \tilde{s}_{kijt} - \Delta \ln \tilde{s}_{krjt})^2$, $X_{2kijt} = (\Delta \ln \tilde{s}_{kijt} - \Delta \ln \tilde{s}_{krjt})(\Delta \ln \tilde{P}_{kijt} - \Delta \ln \tilde{P}_{krjt})$ and $u_{kijt} = (\varepsilon_{kijt} - \varepsilon_{krjt})(\delta_{kijt} - \delta_{krjt}) \frac{(\sigma_{kj}-1)(1+\omega_{kj})}{1+\omega_{kj}\sigma_{kj}}$. Equation (8) still suffers from an endogeneity problem. We follow Feenstra (1994) and instrument the regressors with their country-sector specific averages. Identification is based on the cross-sectional dimension of equation (8). We also include an intercept specific to each HS6 category, to account for the measurement error arising from using unit values to approximate prices, and correct the estimation for heteroskedasticity across exporters i.

The system summarized by equation (8) can accommodate developments that are specific to each sector k. But in macroeconomic applications where the universe of economic activities that form Gross Domestic Product is considered, it is important to allow for more general, aggregate influences. For instance, aggregate inflation in country j is likely to create a component in Y_{kijt} that is common across all sectors k. If it were a shock in the exporting economy, that would correspond to a common component of υ_{kit} across all k. We allow for such correlated effects in as general and parsimonious a manner as possible. We implement a correction suggested by Pesaran (2006) to purge all "Common Correlated Effects" (CCE) from sector level data, and estimate

$$Y_{kijt} = \theta_0 + \theta_{1k}\hat{X}_{1kij} + \theta_{2k}\hat{X}_{2kij} + \theta_{3k}Y_{ijt} + \theta_{4k}X_{1ijt} + \theta_{5k}X_{2ijt} + u_{kijt} \quad (9)$$

where the intercept is allowed to vary by HS6 category, to account for measurement error. Equation (9) is therefore estimated on the within-sector dimension. Hatted variables are the instrumented versions of X_{1kijt} and X_{2kijt}, and Y_{ijt}, X_{1ijt} and X_{2ijt} control for the time-varying component of the variables of interest that is common across all sectors. In particular, following Pesaran (2006), Y_{ijt}, X_{1ijt} and X_{2ijt} are the cross-sector arithmetic averages of Y_{kijt}, X_{1kijt} and X_{2kijt}.[3]

The main difference between the estimator used here and Feenstra (1994) corresponds to the controls for CCE, captured by θ_{3k}, θ_{4k} and θ_{5k}. They are present in equation (9) because, unlike Feenstra (1994), this paper's purpose is macroeconomic in nature. It is important to purge estimates of θ from aggregate

[3] We select sequentially which CCE to include, starting with all three of them. We drop whichever terms are not significant, and iterate until all included CCE are significant. This procedure is implemented on the version of equation (9) where sectoral heterogeneity is assumed away. In practice, we end up including only Y_{ijt} for most countries j.

influences at the sector level, since not doing so would result in double-counting across the panel of sectors. In practice we let the data tell us which CCE are effectively significant for each country j, and modify specification (9) accordingly.

Estimates of equation (8) map directly with the parameters of interest, since

$$\theta_{1kj} = \frac{\omega_{kj}}{(\sigma_{kj}-1)(1+\omega_{kj})}, \quad \theta_{2kj} = \frac{\omega_{kj}\sigma_{kj} - 2\omega_{kj} - 1}{(\sigma_{kj}-1)(1+\omega_{kj})}$$

With consistent, country- and sector-specific estimates of θ_{1kj} and θ_{2kj}, it is straightforward to infer the parameters of interest. In particular, the model implies

$$\begin{aligned} \hat{\sigma}_{kj} &= 1 + \frac{\hat{\theta}_{2kj} + \Delta_{kj}}{2\hat{\theta}_{1kj}} \text{ if } \hat{\theta}_{1kj} > 0 \text{ and } \hat{\theta}_{1kj} + \hat{\theta}_{2kj} < 1 \\ \hat{\sigma}_{kj} &= 1 + \frac{\hat{\theta}_{2kj} - \Delta_{kj}}{2\hat{\theta}_{1kj}} \text{ if } \hat{\theta}_{1kj} < 0 \text{ and } \hat{\theta}_{1kj} + \hat{\theta}_{2kj} > 1 \end{aligned} \tag{10}$$

with $\Delta_{kj} = \sqrt{\hat{\theta}_{2kj}^2 + 4\hat{\theta}_{1kj}}$. Standard deviations are obtained using a first-order approximation around these point estimates.[4]

With consistent estimates of σ_{kj}, we can infer values for the aggregate import and export price elasticities using equations (3) and (5). We can feed the definitions of both trade elasticities with estimates of $\hat{\sigma}_{kj}$ that vary by sector and by country. Then equations (3) and (5) can be used to decompose the international differences in η_j^M and η_j^X. We can also estimate a version of equation (8) where the elasticities of substitution are constrained to sectoral homogeneity, i.e. $\sigma_{kj} = \sigma_j$. We now discuss the differences that makes.

3.2 Heterogeneity Bias

Estimates of η_j^X and η_j^M have different interpretations depending on whether they are computed on the basis of heterogeneous values of $\hat{\sigma}_{kj}$, or using estimates constrained to homogeneity across sectors. In

[4] As is apparent, there are combinations of estimates in equation (10) that do not correspond to any theoretically consistent estimates of $\hat{\sigma}_{kj}$. We follow Broda and Weinstein (2006) and use a search algorithm that minimizes the sum of squared residuals in equation (10) over the intervals of admissible values of the elasticities. The standard errors are bootstrapped in these instances.

Imbs and Méjean (2009), we argue constrained estimates suffer from a heterogeneity bias. The end value obtained for $\hat{\sigma}_j$ does not necessarily reflect a weighted average of sector-specific estimates that is grounded in theory. This can be a problem for the calibration choice of a representative agent in general equilibrium. A representative agent whose preferences are to reflect the heterogeneity in substitution elasticities across sectors should be endowed with an adequately weighted average of $\hat{\sigma}_{kj}$, not with the value obtained for $\hat{\sigma}_j$. We show in Imbs and Méjean (2009) that using aggregate data in fact constrains estimates of $\hat{\sigma}_{kj}$ to homogeneity across sectors. As a result, we argue estimates based on aggregate data should not be used to calibrate the elasticity of substitution of a representative agent in an international context.

In this paper, our purpose is different. We are after estimates of reduced form elasticities, inferred from the calibration of theory-implied weighting schemes, applied to structural estimates of σ, the elasticity of substitution. Since we are concerned with reduced form elasticities, the possibility of a heterogeneity bias in constrained estimates of $\hat{\sigma}_j$ only matters in as much as it alters the end interpretation of η_j^M and η_j^X.

In particular, if it is aggregate data that the estimates of η_j^M and η_j^X are to replicate, then constrained estimates of $\hat{\sigma}_j$ are of the essence. Identification issues notwithstanding, the elasticity of imports (or exports) to a shift in relative prices, that is estimated on aggregate data implicitly imposes that $\hat{\sigma}_{kj} = \hat{\sigma}_j$ for all k. The intuition is straightforward. Aggregating the data suppresses mechanically any sectoral dimension from the estimation, which may result in different results than aggregating sectoral *estimates*. If a discrepancy exists, there is a heterogeneity bias, and it is caused by the assumption that $\hat{\sigma}_{kj} = \hat{\sigma}_j$ in macro data. In that case, the behavior of macro data can only be mimicked by constrained estimates of $\hat{\sigma}_j$, fed into η_j^M and η_j^X.[5]

In contrast, the values of η_j^M and η_j^X based on heterogeneous estimates of $\hat{\sigma}_{kj}$ cannot hope to reproduce the dynamic response of aggregate trade to price shocks. But they are informative nonetheless. Constrained trade elasticities do not depend on the specialization of trade, simply because the weighting schemes in equations (3) and (5) become largely innocuous if $\hat{\sigma}_{kj} = \hat{\sigma}_j$. Unconstrained trade elasticities, on the other hand, will take different values depending on the specialization of production and trade across sectors. The

[5] To be precise, we show in Imbs and Méjean (2009) that aggregate data in fact assumes $\hat{\sigma}_{kj} = \hat{\sigma}_j = \gamma_j$, since with macro aggregates, different sectors become impossible to differentiate from each other. There is however nothing we can say about the empirical value of γ_j, which we calibrate throughout the paper.

unconstrained estimate of η_j^X will take larger values in countries specialized in substitutable exports, but not necessarily its constrained counterpart. Since the two alternatives are conceptually different, we report both variants in this report.

The sectoral dimension plays another important role. Disaggregated data lend power to our estimation, so that we are able to identify with precision country-specific estimates. In contrast, the conventional literature has often had to use cross-country data to identify one parameter, typically assumed to be the same for all considered countries. The assumption is almost certainly not innoucous. Anecdotal evidence is plentiful that the elasticity is actually heterogeneous across countries. There are observable differences in the trade performance of Euro-zone countries in response to a given Euro appreciation. Journalistic discussions are frequent, for instance of the differences between resilient German and French exports. And in general, global shocks to international relative prices do not appear to have identical consequences across countries. The entry of China in world markets, and the accompanying fall in the relative price of Chinese goods, has not affected trade balances identically everywhere. Yet, these international differences are absent from most modelling choices in academia or policy circles. We conjecture such is the case for lack of reliable estimates across countries.

3.3 Data

A structural estimate of $\hat{\sigma}_j$ requires that we observe the cross-section of imported quantities and unit values at the sector level, and for all countries j. We choose to use the multilateral trade database ComTrade, released by the United Nations. The data traces multilateral trade at the 6-digit level of the harmonized system (HS6), and cover around 5,000 products for a large cross-section of countries. We focus on the recent period, and use yearly data between 1995 and 2004. We start in 1995, as before then the number of reporting countries displayed substantial variation. In addition, the unit values reported in ComTrade after 2004 display large time variations that seem to correspond to a structural break. The database is freely available online, to facilitate replication of our results. We focus on the data based on export declarations, to maximize country coverage.

Thanks to the multilateral dimension of our data, we are able to estimate $\hat{\sigma}_j$ for a wide range of countries j. Identification requires the cross-section of countries exporting to j be wide enough, for all sectors. And since the precision of our estimates depends on the time-average of these trade data, we also need the cross-section of exporting countries (and goods) to be as stable over time as possible. We therefore only retain goods for which a minimum of 20 exporting countries are available throughout the period we consider. In addition, both unit values and market shares are notoriously plagued by measurement error. We limit the influence of outliers, and compute the median growth rate at the sector level for each variable, across all countries and years. We exclude the bilateral trade flows for all sectors whose growth rates exceed five time that median value in either unit values or in market share. On average, the resulting truncated sample covers about 85 percent of world trade. Table 1 presents some summary statistics for the 33 countries we have data for. The number of sectors ranges from 10 to 26. We also report the total number of exporters into each country j, equal to the product of the number of sectors in country j, times the number of exporters for each sectors. This suggests the average number of exporters ranges from 35 in Sri Lanka to more than 90 in the United States. For each sector, our data implies an average number of exporting countries of 53.

The main constraint on the range of countries for which we can obtain end estimates of η_j^X and η_j^M is not imposed by the availability of trade data. It is the calibration of sectoral shares that is limited by the availability of adequate data. Computing aggregate trade elasticities requires the calibration of six weights. We need values for m_{kj} and x_{kj}, which denote the value share of sector k in the aggregate imports and exports of country j, respectively. We need a value for x_{kji}, which is the share of country j's exports of product k sold in country i. There are also three consumption shares: w_{kj}, which denotes the share of sector k in country j's nominal consumption, w_{kjj}, the share of domestically produced goods in sector k consumption, and w_{kji}, the share of sector k consumption in country i that is imported from country j.

Consumption data require information on domestic production at sectoral level, across as large a cross-section of countries as possible. This is absent from conventional international trade databases. We resort instead to UNIDO data, which report nominal sectoral output at the 3-digit ISIC (revision 2) level. UNIDO data are incomplete for some countries. Since aggregation can become misleading for countries where too few sectors are reported, we impose a minimum of 10 sectors for all countries j. This tends to exclude small or developing economies, such as Hong Kong, Panama, Slovakia or Poland. The data are expressed in USD, and available at a yearly frequency. To limit the consequences of measurement error, we use five-year averages of the relevant UNIDO data. We experiment with weights computed between 1991 and 1995, or between 1996 and 2000. We merge multilateral trade data into the ISIC classification.

The UNIDO dataset is focused on manufacturing goods only, which truncates somewhat the original coverage in trade data. But the vast majority of traded goods are manufactures, so that the sampling remains minimal. We have experimented with weights implied by the OECD Structural Analysis database (STAN): for countries covered by both datasets, the end elasticities were in fact virtually identical - even though STAN provides information on all sectors of the economy. UNIDO has sectoral information on many more countries, not least non-OECD members like China. Such coverage is important in its own right, but it is also of the essence when it comes to computing export elasticities. The price elasticity of exports involves an average across destination markets for all countries considered. Focusing on just OECD economies would complicate the interpretation of our end estimates, as they would ignore non-OECD trade flows, which have increased in magnitude of late. The last column in Table 1 reports the percentage of total trade as implied by ComTrade, that we continue to cover once we restrict the sample to sectors for which we have UNIDO data. The coverage is below 20 percent for small open economies such as Hong Kong, Singapore, or the Philippines, and around three-quarters for large developed economies such as the US, France or Spain.

The values for m_{kj} and x_{kj} are calculated as the ratios of sectoral imports and exports, relative to their respective aggregate at the country level. The value for x_{kji} is computed as the ratio of good k's exports from country j that are sold in i. A choice must be made when it comes to the sectors used in computing x_{kj}. A first option is to consider the coverage in ComTrade data for both m_{kj} and x_{kj} , but without any guarantee they are the same. Such choice implies that we allow for the sectoral patterns of exports from country j to differ from what it imports. The assumption seems realistic and plausible, but it takes missing data from ComTrade at face value - even though data collection issues or measurement error may in reality be the reason why a sector is missing from ComTrade. In addition, the choice also implies that the trade weights x_{kj} are computed on a different sample than the consumption weights. This may create issues of comparability, as the consumption weights are constructed on the basis of a sectoral coverage put together by Di Giovanni and Levchenko (2009), with a view to accounting for outliers.

An alternative is to use the trade data compiled by Di Giovanni and Levchenko (2009), and effectively imposing no specialization, as countries import and export the same sectors. The assumption is clearly at odds with the evidence, but it is prudent. There are many sectors with zero entries in disaggregated trade data, and the distinction between effectively non traded goods, measurement error, or improvement in data collection over time is often difficult. With a no-specialization assumption, we maximize the comparability of our estimates of trade elasticities on the import and the export sides. This is done at the potential cost of realism.

For the three other weights, we require information on both production and trade at the sectoral level. In order to maximize concordance and comparability, we use a dataset built by Di Giovanni and Levchenko (2009) who merge information on production from UNIDO and on bilateral trade flows from the World Trade Database compiled by Feenstra et al (2005). Domestic consumption at the sectoral level is computed as production net of exports, and overall consumption is production net of exports but inclusive of imports. We have

$$w_{kjj} \equiv \frac{Y_{kj} - X_{kj}}{Y_{kj} - X_{kj} + M_{kj}}$$

where X_{kj} (M_{kj}) denotes country j's exports (imports) of good k.

$$w_{kji} = \frac{X_{kji}}{Y_{ki} - X_{ki} + M_{ki}} = \frac{x_{kji}}{\sum_j x_{kji}} (1 - w_{kii})$$

where X_{kji} are country j's exports of good k sold in country i. And

$$w_{kj} \equiv \frac{Y_{kj} - X_{kj} + M_{kj}}{\sum_k (Y_{kj} - X_{kj} + M_{kj})}$$

We experimented with alternative combinations of data sources. Rather than using the dataset merged by Di Giovanni and Levchenko (2009), we combined data from ComTrade for sectoral imports or exports, and from UNIDO for output. In all instances, however, we made use of output data corresponding to the UNIDO data treated for outliers by Di Giovanni and Levchenko (2009). This is important, for it ensures the compatibility of production and trade data, and the meaningfulness of the measures w of local consumption. We did however use production data from the OECD STAN as well, with substantially smaller country coverage. For those countries where data are available across all three alternatives, our end results are virtually unchanged.

4 Results

We present first cross-country estimates of import elasticities. We then turn to exports. In both instances, we report elasticity estimates obtained from both constrained and unconstrained estimates of the preference parameter σ_{kj}. We also present extensive robustness analysis. We consider different calibrated values for the cross-sector elasticity of substitution γ, and use different sub-periods to compute the weights w, m and x.

4.1 Import Price Elasticities

Table 2 reports the constrained estimates of η_j^M for 33 countries at various levels of development. The elasticities are computed imposing $\gamma = 1$. The left panel of the table uses weights computed on the basis of a 5-year average between 1991 and 1995. There is a lot of heterogeneity across countries, with values ranging from 0.5 in Hong Kong to 2.7 in Italy or Finland. Small, open countries tend to display low elasticities, with Hong Kong, Singapore or Austria all taking values around 1. Large, rich economies tend to have high elasticities, with France, the US, Spain, Japan, Australia or Italy all around 2.

Commodity exporters tend to have large elasticities, with elasticitites above 1.5 for Venezuela, Australia, or Indonesia. The exception is Kuwait, which takes one of the lowest values in our sample, at 0.6. Kuwait is extremely specialized in oil products, and is probably an outlier in that sense. There is no apparent mapping between income per capita and import elasticities, as there are developed countries at both ends of the distribution. The import elasticity in China is estimated at 1.57, with a tight standard error of 0.2.

In the US and Japan, the import elasticity is closer to 2, with standard errors below 0.2. The big European economies have estimates in the same region - with French, Spanish and German elasticities at or around 2. The value for Italy is higher, at 2.7, although the estimate is less precise with standard errors equal to 0.4. Small European economies tend to have low elasticities: Greece and Austria are below 1, Hungary, Norway, the UK and Portugal are close to 1.5. The main lesson from Table 2 is a large cross-country heterogeneity.

The estimated elasticities change little when weights used in aggregation are measured over the end of the 1990's, between 1996 and 2000. For the countries with observations in both sub-periods, the discrepancy between the values on the left and right panels of Table 1 is virtually never significant, compared to the standard deviations reported in the table. Exceptions are Canada, whose elasticity falls from 1.79 to 1.36, Australia, which goes from 2.04 to 1.54, and perhaps Israel and Germany, which go from 2.19 to 1.54 and 1.72 to 1.34, respectively. In general, the magnitude and ranking of import elasticities are largely robust to aggregation weights that are computed over different time periods. Of course, this can reflect the well known persistence in international trade patterns - even though the 1990's saw the rapid emergence of China in world trade.

Table 3 presents the unconstrained estimates of η_j^M for the same parametrization. Estimates are now substantially more heterogeneous, ranging from 0.7 to 7.14. Interestingly, the ranking of countries is altered somewhat. There is now a clear tendency for large, emerging markets to have high elasticity values. China is the most striking, that now tops the ranking of 33 countries with an elasticity estimate of 7.14. The same can be said of India, Korea, or Turkey. Rich developed countries now tend to be at the bottom of the distribution, with elasticity estimates below 3. This difference in ranking reflects the importance of specialization in driving the (unconstrained) elasticity of imports. Rich countries tend to import goods that are not substitutable, whereas the opposite can be said of large developing economies. Such importance of specialization was not apparent from Table 2. These relative rankings remain virtually unchanged when the weighting scheme is chosen using data from the second half of the 1990's.

Tables 4 and 5 present the constrained and unconstrained estimates of η_j^M for $\gamma = 2$. Goods across sectors are now assumed to be substitutes. From equation (3), we expect alternative values of γ to have a minimal impact on the level of aggregate trade elasticities. The two Tables confirm this to be the case. The magnitudes and distribution of elasticities in Table 4 resembles Table 2, and unconstrained estimates in Table 5 are virtually identical to Table 3. In particular, the same two results obtain. Constrained estimates are relatively less heterogeneous. And unconstrained estimates reflect differences in income per capita, and presumably of the specialization of production and trade. The same can be said for estimates implied by a calibration of $\gamma = 0.5$, which are reported in Tables 6 and 7.

4.2 Export Price Elasticities

Computing the price elasticities of exports entails averaging across two dimensions: across sectors, and across destination markets. The latter is performed with the weights x_{kji}, which capture the international allocation of exports from country j. We only estimate substitution elasticities for the destination countries in our sample, so that $\sum_{i \neq j} x_{kji}$ is in fact not unity in our data. In what follows, we experiment both with and without normalization , i.e. report estimates of η_j^X arising from the actual weights x_{kji} that we observe, and also from weights that are normalized so that $\sum_{i \neq j} x_{kji} = 1$. The normalization amounts to assuming the export markets we do not observe are in fact negligible in country j's trade.

Table 8 reports constrained estimates of the export elasticity for 33 countries, where we do not normalize $\sum_{i \neq j} x_{kji}$ to unity, and use the UNIDO coverage to compute the export weights x_{kj}, i.e. we assume specialization away. Estimates correspond to $\gamma = 1$. Table 9 reports the corresponding unconstrained elasticities. Constrained estimates display some heterogeneity, with values ranging from 0.9 to 2.25. Developed countries tend to take relatively low values, with Singapore, the US, Germany, the UK, Japan, France or Austria all around 1.5. Small open economies, such as Canada, Portugal, Hong Kong or Israel take larger values, around 2. Using weights from the late 90's does not alter these magnitudes, nor indeed the ranking.

As with import elasticities, heterogeneity is magnified for unconstrained elasticities. Estimates in Table 9 now range from 1.43 to 5.21. The ranking of countries remains however largely unchanged, with developed countries taking relatively low values. Interestingly, both India and China have export elasticities that are similar in magnitude with the US, France, the UK, or more generally rich OECD economies. This result does not depend on whether estimates are constrained or unconstrained. Unconstrained export elasticities are highest once again for small open economies, with Australia, Canada and Hong Kong taking the highest three values. Using weights from the late 1990's barely alters these results, with the possible exception of the German elasticity that increases significantly, whereas Korea's and Israel's fall in absolute value.

Tables 10 and 11 impose the normalization that $\sum_{i \neq j} x_{kji} = 1$, but continue to use the same sectoral coverage for x_{kj} as is done for import elasticities. As expected, the normalization has level effects on all estimates, since x_{kji} now take higher values for all countries. Constrained estimates now range from 1.9 to 3.0, whereas unconstrained estimates range from 3.5 to 6.0. Unconstrained estimates continue to display higher heterogeneity. In addition, the ranking of countries is altered. With the exception of China, which takes the lowest value in both Tables, the rank correlation between income per capita and η_j^X (in absolute value) is now firmly negative. Rich countries take relatively low values, as one would expect from countries that export branded, differentiated goods. Poor countries, and small open economies tend to take high values - in a way that is even more apparent for the unconstrained estimates in Table 11. The case of China is intriguing, and its ranking changes drastically compared with Tables 8 and 9. An explanation may lie in the fact that imposing $\sum_{i \neq j} x_{kji} = 1$ in fact assumes our data actually contains all export markets of country j. The difference between the estimates reported in Tables 8 and 10 will be largest, the more imperfect our actual coverage. Estimates of η_j^X for China in fact increase little between Tables 8 and 10, which suggests our coverage of China's trade may in fact be substantially better than what it is for other countries, where estimates of η_j^X increase more. In fact, the coverage percentages reported in Table 1 do suggest that, amongst developing economies, Chinese data have relatively high coverage, almost equal to 50% of all available trade flows in ComTrade. To limit the incidence of cross-country differences in coverage on our elasticity estimates, we do not impose $\sum_{i \neq j} x_{kji} = 1$ in what follows.

Tables 12 and 13 report elasticity estimates imposing sectoral specialization as implied by ComTrade data. Comparing Tables 12 and 8, it is patent that using ComTrade data has little consequence on the magnitude and ranking of elasticity estimates. η_j^X range from 0.91 to 2.25 in Table 8 (without specialization), and from 1.02 to 2.24 in Table 12. Differences in ranking are too small to detect, and estimates for most big economies are virtually identical. The comparison between Tables 9 and 13 implies similar comments, with only one noticeable change, namely Venezuela's elasticity doubling to 6.7 when specialization is imposed. This suggests the assumption that Venezuela imports and exports in the same sectors is especially implausible.

Tables 14, 15, 16 and 17 report estimates for $\gamma = 2$, or $\gamma = 0.5$, while maintaining the specialization hypothesis. The elasticity estimates are virtually unchanged, and country rankings continue to have the properties discussed previously. Different calibrates values for γ seem to have little effects on the end estimates of η_j^X.

5 Conclusion

We describe a CES demand system where the price elasticities of exports and imports are given by weighted averages of the international elasticity of substitution. We adapt the econometric methodology in Feenstra (1994) and Imbs and Méjean (2009) to estimate structurally the substitution elasticity for a broad range of countries. We collect from a variety of sources the weights that theory implies should be used to infer both imports and exports elasticities. We report trade elasticity estimates for more than 30 countries, including most developed and the major developing economies. We find susbtantial cross-country dispersion, which is robust to alternative measurement strategies. Our benchmark calibration imply constrained import elasticitites ranging from 0.5 to 2.7. While constrained import elasticities display little cross-country correlation with income per capita, unconstrained elasticities are clearly higher for open, developing, specialized economies - and closer to 0 in rich countries. Export elasticities display relatively less dispersion, and range (in absolute value) from 0.9 to 2.25. Rich open economies tend to take low absolute values, and developing countries have relatively high estimates. Alternative calibration choices for the cross-sector elasticity γ have little end consequences on trade elasticities. The same is true of the period used to calibrate the weights used in aggregating elasticities of substitution.

6 References

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Table 1: Summary Statistics

	Weights 1991-1995			Weights 1996-2000		
	# sect	# sect×exp	% Trade	# sect	# sect×exp	% Trade
Australia	16	763	47.7	18	851	50.1
Austria	24	1127	69.5	19	898	60.9
Belgium				13	857	24.7
Canada	24	1415	63.0	24	1415	63.0
Sri Lanka	10	355	24.3	10	355	24.3
China	18	989	49.3	17	924	35.3
Finland	25	1116	61.0	22	998	59.8
France	25	1950	74.0	26	2007	74.1
Germany	20	1506	50.4	21	1602	65.7
Greece	17	904	47.1			
Hong Kong	11	596	15.3			
Hungary	19	856	44.5	15	667	18.8
Indonesia	14	575	39.5			
Israel	20	868	38.7	14	620	32.6
Italy	25	1854	71.3	26	1918	72.6
Japan	26	1470	61.4	26	1470	61.4
Korea	25	1170	54.6	25	1170	54.6
Kuwait	15	640	24.6	13	559	21.7
Taiwan	20	972	40.3			
Norway	20	904	48.7	23	1029	51.4
Philippines	15	589	17.1			
Poland				23	1113	63.6
Portugal	22	1029	64.4	26	1194	70.6
India	17	884	38.1	17	884	38.1
Singapore	10	496	12.3			
Slovakia	10	369	23.4			
Spain	26	1725	75.1	26	1725	75.1
Sweden	25	1305	72.0	21	1067	54.0
Turkey	23	1182	53.1	20	1020	48.9
Egypt	19	925	37.7			
United Kingdom	26	2015	77.4	24	1838	73.0
United States	26	2347	77.1	25	2258	76.7
Venezuela	20	758	37.1			

Table 2: ComTrade Data (export declarations), $\gamma = 1$, constrained estimates

	Weights 1991-1996			Weights 1996-2000		
	# sect.	η^M	SD	# sect.	η^M	SD
Hong Kong	11	-0.538	0.086			
Kuwait	15	-0.571	0.104	13	-0.594	0.111
Greece	17	-0.875	0.127			
Austria	24	-0.879	0.118	19	-0.711	0.089
Sweden	25	-0.965	0.085	21	-1.331	0.212
Egypt	19	-0.971	0.151			
Singapore	10	-1.078	0.189			
Hungary	19	-1.329	0.242	15	-1.964	0.635
Norway	20	-1.398	0.158	23	-1.399	0.155
Venezuela	20	-1.428	0.307			
Turkey	23	-1.434	0.221	20	-1.212	0.195
Taiwan	20	-1.437	0.119			
United Kingdom	26	-1.455	0.128	24	-1.209	0.105
Indonesia	14	-1.535	0.365			
Portugal	22	-1.559	0.272	26	-1.375	0.216
China	18	-1.571	0.233	17	-1.508	0.201
Sri Lanka	10	-1.664	0.575	10	-1.605	0.554
Germany	20	-1.7195	0.10158	21	-1.339	0.144
Korea	25	-1.765	0.306	25	-1.677	0.291
Canada	24	-1.787	0.307	24	-1.357	0.233
France	25	-1.844	0.214	26	-1.744	0.202
United States	26	-1.937	0.140	25	-2.091	0.223
Spain	26	-1.948	0.238	26	-1.721	0.211
Japan	26	-1.999	0.199	26	-1.897	0.189
Slovakia	10	-2.015	0.412			
Australia	16	-2.040	0.417	18	-1.541	0.251
Israel	20	-2.189	0.443	14	-1.540	0.283
India	17	-2.238	0.293	17	-2.276	0.298
Philippines	15	-2.393	0.674			
Italy	25	-2.681	0.397	26	-2.796	0.417
Finland	25	-2.715	0.790	22	-2.410	0.643
Belgium				13	-1.282	0.171
Poland				23	-1.225	0.166

Table 3: ComTrade Data (export declarations), $\gamma = 1$, Unconstrained estimates

	Weights 1991-1996			Weights 1996-2000		
	# sect.	η^M	SD	# sect.	η^M	SD
Hong Kong	11	-0.713	0.045			
Kuwait	15	-0.807	0.071	12	-0.857	0.020
Singapore	10	-1.697	0.223			
Sweden	25	-1.943	0.111	21	-2.052	0.168
Austria	24	-1.996	0.108	19	-1.778	0.102
Norway	20	-2.158	0.105	23	-2.443	0.117
Indonesia	14	-2.163	0.155			
Greece	17	-2.283	0.110			
Venezuela	20	-2.391	0.219			
Germany	20	-2.433	0.057	21	-2.386	0.064
Egypt	19	-2.613	0.148			
United Kingdom	26	-2.635	0.077	24	-2.277	0.077
Portugal	22	-2.637	0.196	26	-2.370	0.158
France	25	-2.867	0.093	26	-2.717	0.094
Canada	24	-2.878	0.301	24	-2.163	0.206
Spain	26	-2.905	0.125	26	-2.548	0.111
Italy	25	-2.920	0.087	26	-2.951	0.086
Hungary	19	-3.063	0.200	15	-2.085	0.110
Slovakia	10	-3.124	0.238			
Finland	25	-3.148	0.230	22	-3.041	0.191
Israel	20	-3.193	0.173	14	-3.587	0.270
Philippines	15	-3.337	0.228			
Taiwan	20	-3.461	0.219			
Australia	16	-3.647	0.342	18	-2.927	0.292
United States	26	-4.196	0.189	25	-4.112	0.216
Sri Lanka	10	-4.476	0.052	10	-4.135	0.054
Korea	25	-4.723	0.344	25	-4.485	0.363
Japan	26	-5.150	0.344	26	-4.869	0.319
India	17	-6.170	0.384	17	-6.052	0.397
Turkey	23	-6.403	0.293	20	-5.179	0.239
China	18	-7.139	0.484	17	-6.697	0.675
Belgium				13	-1.963	0.116
Poland				23	-3.069	0.207

Table 4: ComTrade Data (export declarations), $\gamma = 2$, Constrained estimates

	Weights 1991-1996			Weights 1996-2000		
	# sect.	η^M	SD	# sect.	η^M	SD
Hong Kong	11	-0.576	0.086			
Kuwait	15	-0.619	0.104	13	-0.655	0.111
Greece	17	-0.973	0.127			
Austria	24	-1.024	0.118	19	-0.824	0.089
Sweden	25	-1.102	0.085	21	-1.458	0.212
Singapore	10	-1.127	0.189			
Egypt	19	-1.135	0.151			
Hungary	19	-1.454	0.242	15	-2.135	0.635
Taiwan	20	-1.493	0.119			
Turkey	23	-1.549	0.221	20	-1.338	0.195
United Kingdom	26	-1.559	0.128	24	-1.340	0.105
Norway	20	-1.566	0.158	23	-1.550	0.155
Venezuela	20	-1.586	0.307			
China	18	-1.653	0.233	17	-1.576	0.201
Indonesia	14	-1.717	0.365			
Portugal	22	-1.727	0.272	26	-1.542	0.216
Sri Lanka	10	-1.760	0.575	10	-1.718	0.554
Germany	20	-1.817	0.102	21	-1.417	0.144
Korea	25	-1.872	0.306	25	-1.782	0.291
Canada	24	-1.911	0.307	24	-1.479	0.233
France	25	-1.931	0.214	26	-1.840	0.202
United States	26	-2.011	0.140	25	-2.166	0.223
Japan	26	-2.048	0.199	26	-1.960	0.189
Spain	26	-2.074	0.238	26	-1.864	0.210
Slovakia	10	-2.079	0.412			
Australia	16	-2.207	0.417	18	-1.725	0.251
India	17	-2.309	0.293	17	-2.337	0.298
Israel	20	-2.339	0.443	14	-1.702	0.283
Philippines	15	-2.516	0.674			
Italy	25	-2.758	0.397	26	-2.897	0.417
Finland	25	-2.874	0.790	22	-2.549	0.643
Belgium				13	-1.371	0.171
Poland				23		

Table 5: ComTrade Data (export declarations), $\gamma = 2$, Unconstrained estimates

	Weights 1991-1996			Weights 1996-2000		
	# sect.	η^M	SD	# sect.	η^M	SD
Hong Kong	11	-0.751	0.045			
Kuwait	15	-0.855	0.071	12	-0.922	0.020
Singapore	10	-1.746	0.223			
Sweden	25	-2.080	0.111	21	-2.179	0.168
Austria	24	-2.141	0.108	19	-1.891	0.102
Norway	20	-2.327	0.105	23	-2.593	0.117
Indonesia	14	-2.345	0.155			
Greece	17	-2.381	0.110			
Germany	20	-2.531	0.057	21	-2.464	0.064
Venezuela	20	-2.548	0.219			
United Kingdom	26	-2.739	0.077	24	-2.408	0.077
Egypt	19	-2.777	0.148			
Portugal	22	-2.805	0.196	26	-2.538	0.158
France	25	-2.954	0.093	26	-2.813	0.094
Italy	25	-2.998	0.087	26	-3.051	0.086
Canada	24	-3.002	0.301	24	-2.285	0.206
Spain	26	-3.031	0.125	26	-2.692	0.111
Slovakia	10	-3.188	0.238			
Hungary	19	-3.188	0.200	15	-2.256	0.110
Finland	25	-3.307	0.230	22	-3.180	0.191
Israel	20	-3.342	0.173	14	-3.748	0.270
Philippines	15	-3.459	0.228			
Taiwan	20	-3.518	0.219			
Australia	16	-3.814	0.342	18	-3.110	0.292
United States	26	-4.270	0.189	25	-4.187	0.216
Sri Lanka	10	-4.572	0.052	10	-4.248	0.054
Korea	25	-4.830	0.344	25	-4.591	0.363
Japan	26	-5.199	0.344	26	-4.933	0.319
India	17	-6.240	0.384	17	-6.113	0.397
Turkey	23	-6.518	0.293	20	-5.305	0.239
China	18	-7.221	0.484	17	-6.765	0.675
Belgium				13	-2.052	0.116
Poland				23	-3.201	0.207

Table 6: ComTrade Data (export declarations), $\gamma = 0.5$, Constrained estimates

	Weights 1991-1996			Weights 1996-2000		
	# sect.	η^M	SD	# sect.	η^M	SD
Hong Kong	11	-0.520	0.086			
Kuwait	15	-0.547	0.104	13	-0.564	0.111
Austria	24	-0.806	0.118	19	-0.654	0.089
Greece	17	-0.826	0.127			
Egypt	19	-0.888	0.151			
Sweden	25	-0.897	0.085	21	-1.267	0.212
Singapore	10	-1.053	0.189			
Hungary	19	-1.267	0.242	15	-1.878	0.635
Norway	20	-1.313	0.158	23	-1.324	0.155
Venezuela	20	-1.349	0.307			
Turkey	23	-1.377	0.221	20	-1.149	0.195
United Kingdom	26	-1.403	0.128	24	-1.143	0.105
Taiwan	20	-1.409	0.119			
Indonesia	14	-1.443	0.365			
Portugal	22	-1.476	0.272	26	-1.290	0.216
China	18	-1.530	0.233	17	-1.474	0.201
Sri Lanka	10	-1.616	0.575	10	-1.548	0.554
Germany	20	-1.671	0.102	21	-1.300	0.144
Korea	25	-1.711	0.306	25	-1.625	0.291
Canada	24	-1.726	0.307	24	-1.295	0.233
France	25	-1.801	0.214	26	-1.696	0.202
Spain	26	-1.885	0.238	26	-1.649	0.210
United States	26	-1.899	0.140	25	-2.054	0.223
Australia	16	-1.956	0.417	18	-1.450	0.251
Japan	26	-1.975	0.199	26	-1.865	0.189
Slovakia	10	-1.983	0.412			
Israel	20	-2.115	0.443	14	-1.460	0.283
India	17	-2.203	0.293	17	-2.246	0.298
Philippines	15	-2.332	0.674			
Finland	25	-2.636	0.790	22	-2.341	0.643
Italy	25	-2.642	0.397	26	-2.746	0.417
Belgium				13	-1.238	0.171
Poland				23	-1.159	0.166

Table 7: ComTrade Data (export declarations), $\gamma = 0.5$, Unconstrained estimates

	Weights 1991-1996			Weights 1996-2000		
	# sect.	η^M	SD	# sect.	η^M	SD
Hong Kong	11	-0.695	0.045			
Kuwait	15	-0.783	0.071			
Singapore	10	-1.672	0.223			
Sweden	25	-1.875	0.111	21	-1.988	0.168
Austria	24	-1.923	0.108	19	-1.722	0.102
Indonesia	14	-2.071	0.155			
Norway	20	-2.074	0.105	23	-2.367	0.117
Greece	17	-2.234	0.110			
Venezuela	20	-2.312	0.219	12	-0.824	0.020
Germany	20	-2.384	0.057	21	-2.347	0.064
Egypt	19	-2.530	0.148			
Portugal	22	-2.553	0.196	26	-2.286	0.158
United Kingdom	26	-2.582	0.077	24	-2.212	0.077
Canada	24	-2.816	0.301	24	-2.102	0.206
France	25	-2.824	0.093	26	-2.669	0.094
Spain	26	-2.841	0.125	26	-2.477	0.111
Italy	25	-2.881	0.087	26	-2.901	0.086
Hungary	19	-3.001	0.200	15	-1.999	0.110
Finland	25	-3.069	0.230	22	-2.972	0.191
Slovakia	10	-3.092	0.238			
Israel	20	-3.118	0.173	14	-3.506	0.270
Philippines	15	-3.276	0.228			
Taiwan	20	-3.433	0.219			
Australia	16	-3.563	0.342	18	-2.835	0.292
United States	26	-4.159	0.189	25	-4.075	0.216
Sri Lanka	10	-4.428	0.052	10	-4.078	0.054
Korea	25	-4.669	0.344	25	-4.433	0.363
Japan	26	-5.126	0.344	26	-4.837	0.319
India	17	-6.134	0.384	17	-6.021	0.397
Turkey	23	-6.345	0.293	20	-5.115	0.239
China	18	-7.098	0.484	17	-6.663	0.675
Belgium				13	-1.919	0.116
Poland				23	-3.004	0.207

Table 8: Export Elasticities - No Specialization in x_{kj}, Raw x_{kji}, $\gamma = 1$, Constrained estimates

	Weights 1991-1996		Weights 1996-2000	
	η^X	SD	η^X	SD
Slovakia	-0.907	0.039		
Venezuela	-1.189	0.054		
Turkey	-1.400	0.048	-1.261	0.054
Singapore	-1.412	0.059		
India	-1.413	0.048	-1.218	0.055
Egypt	-1.449	0.067		
US	-1.451	0.086	-1.157	0.075
Germany	-1.473	0.063	-1.673	0.070
UK	-1.505	0.053	-1.544	0.064
Japan	-1.596	0.061	-1.462	0.090
Hungary	-1.604	0.070	-1.643	0.088
China	-1.613	0.070	-1.492	0.076
Finland	-1.624	0.047	-1.619	0.059
Indonesia	-1.634	0.058		
Australia	-1.636	0.063	-1.222	0.057
France	-1.655	0.067	-1.666	0.073
Austria	-1.703	0.070	-1.772	0.095
Korea	-1.703	0.061	-1.345	0.068
Italy	-1.716	0.058	-1.598	0.064
Norway	-1.775	0.061	-1.884	0.082
Sweden	-1.780	0.080	-1.814	0.080
Spain	-1.874	0.091	-1.926	0.097
Israel	-1.896	0.068	-1.672	0.086
Sri Lanka	-1.936	0.061	-2.019	0.100
Hong Kong	-1.962	0.128		
Greece	-2.009	0.104		
Portugal	-2.093	0.088	-2.115	0.097
Canada	-2.249	0.138	-2.525	0.245
Belgium			-1.811	0.092
Kuwait			-0.361	0.021
Taiwan				
Philippines				
Poland			-1.499	0.075

Table 9: Export Elasticities - No Specialization in x_{kj}, Raw x_{kji}, $\gamma = 1$, Unconstrained estimates

	Weights 1991-1996		Weights 1996-2000	
	η^X	SD	η^X	SD
Slovakia	-1.430	0.033		
Turkey	-2.368	0.035	-2.057	0.029
Germany	-2.469	0.034	-3.117	0.081
Egypt	-2.496	0.052		
Finland	-2.555	0.038	-2.596	0.050
Austria	-2.607	0.036	-2.936	0.050
India	-2.610	0.038	-2.248	0.037
UK	-2.654	0.040	-2.802	0.053
Hungary	-2.683	0.054	-2.686	0.057
Venezuela	-2.763	0.153		
China	-2.836	0.056	-2.680	0.064
France	-2.894	0.047	-2.976	0.050
Sweden	-2.918	0.056	-3.181	0.079
US	-2.944	0.106	-2.390	0.089
Indonesia	-3.010	0.045		
Italy	-3.062	0.035	-2.881	0.035
Singapore	-3.105	0.079		
Spain	-3.216	0.066	-3.378	0.068
Greece	-3.276	0.077		
Portugal	-3.321	0.058	-3.680	0.057
Korea	-3.350	0.056	-2.673	0.065
Japan	-3.355	0.116	-2.979	0.134
Norway	-3.406	0.094	-3.519	0.089
Israel	-3.491	0.070	-2.793	0.069
Sri Lanka	-3.580	0.092	-3.674	0.095
Australia	-3.588	0.106	-2.651	0.077
Canada	-4.937	0.360	-5.046	0.432
Hong Kong	-5.206	0.270		
Belgium			-2.862	0.046
Kuwait			-0.831	0.022
Taiwan				
Philippines				
Poland			-2.295	0.031

Table 10: Export Elasticities - No Specialization in x_{kj}, Normalized x_{kji}, $\gamma = 1$, Constrained estimates

	Weights 1991-1996		Weights 1996-2000	
	η^X	SD	η^X	SD
China	-1.973	0.084	-2.585	0.130
Japan	-2.044	0.078	-2.401	0.146
US	-2.098	0.124	-1.948	0.123
Korea	-2.265	0.080	-2.370	0.116
Hong Kong	-2.268	0.148		
Indonesia	-2.319	0.083		
Italy	-2.340	0.079	-2.309	0.093
Canada	-2.383	0.145	-2.653	0.257
Austria	-2.386	0.097	-2.482	0.132
Slovakia	-2.389	0.101		
France	-2.440	0.099	-2.372	0.104
UK	-2.459	0.088	-2.562	0.107
Finland	-2.459	0.071	-2.475	0.091
Germany	-2.471	0.105	-2.390	0.100
Australia	-2.486	0.095	-2.556	0.118
Norway	-2.510	0.086	-2.662	0.117
Hungary	-2.580	0.111	-2.567	0.133
Israel	-2.592	0.093	-2.794	0.147
India	-2.596	0.090	-2.543	0.114
Venezuela	-2.603	0.126		
Sweden	-2.614	0.117	-2.572	0.113
Spain	-2.615	0.126	-2.654	0.133
Singapore	-2.619	0.112		
Turkey	-2.657	0.092	-2.664	0.112
Sri Lanka	-2.763	0.087	-2.616	0.123
Portugal	-2.776	0.118	-2.617	0.121
Egypt	-2.875	0.143		
Greece	-3.017	0.151		
Belgium			-2.619	0.134
Kuwait			-2.424	0.150
Taiwan				
Philippines				
Poland			-2.500	0.124

Table 11: Export Elasticities - No Specialization in x_{kj}, Normalized x_{kji}, $\gamma = 1$, Unconstrained estimates

	Weights 1991-1996		Weights 1996-2000	
	η^X	SD	η^X	SD
China	-3.498	0.067	-4.699	0.110
Austria	-3.630	0.049	-4.086	0.067
Slovakia	-3.795	0.084		
Finland	-3.948	0.056	-4.022	0.077
Germany	-4.152	0.056	-4.378	0.105
Italy	-4.183	0.048	-4.175	0.050
US	-4.227	0.145	-3.981	0.138
France	-4.262	0.067	-4.278	0.074
Indonesia	-4.317	0.067		
Japan	-4.346	0.157	-4.860	0.207
Hungary	-4.349	0.076	-4.177	0.081
Sweden	-4.365	0.090	-4.520	0.114
UK	-4.398	0.071	-4.664	0.089
Portugal	-4.408	0.073	-4.578	0.068
Turkey	-4.475	0.060	-4.335	0.057
Spain	-4.498	0.085	-4.683	0.088
Korea	-4.513	0.077	-4.820	0.119
India	-4.765	0.064	-4.676	0.076
Greece	-4.881	0.105		
Israel	-4.893	0.119	-4.905	0.150
Egypt	-4.921	0.119		
Norway	-5.045	0.133	-5.056	0.118
Sri Lanka	-5.100	0.124	-4.818	0.120
Canada	-5.189	0.367	-5.263	0.441
Australia	-5.459	0.160	-5.487	0.152
Singapore	-5.738	0.146		
Venezuela	-5.761	0.259		
Hong Kong	-6.022	0.318		
Belgium			-4.114	0.063
Kuwait			-5.286	0.332
Taiwan				
Philippines				
Poland			-3.835	0.047

Table 12: Export Elasticities - Specialization in x_{kj}, Raw x_{kji}, $\gamma = 1$, Constrained estimates

	Weights 1991-1996		Weights 1996-2000	
	η^X	SD	η^X	SD
Slovakia	-1.020	0.045		
Venezuela	-1.216	0.068		
Singapore	-1.264	0.044		
Australia	-1.380	0.056	-1.017	0.048
Egypt	-1.446	0.060		
India	-1.467	0.047	-1.315	0.065
United States	-1.477	0.088	-1.178	0.077
Turkey	-1.512	0.051	-1.352	0.058
Germany	-1.518	0.064	-1.629	0.068
United Kingdom	-1.542	0.055	-1.541	0.064
China	-1.554	0.065	-1.403	0.072
Finland	-1.573	0.045	-1.647	0.062
Japan	-1.602	0.061	-1.472	0.091
Hungary	-1.638	0.068	-1.782	0.094
France	-1.663	0.068	-1.679	0.074
Korea	-1.672	0.059	-1.291	0.063
Austria	-1.674	0.067	-1.752	0.094
Italy	-1.688	0.057	-1.613	0.065
Indonesia	-1.692	0.060		
Hong Kong	-1.748	0.083		
Israel	-1.782	0.069	-1.828	0.112
Sweden	-1.788	0.081	-1.799	0.083
Norway	-1.793	0.064	-1.791	0.075
Spain	-1.878	0.092	-1.918	0.096
Greece	-1.888	0.086		
Sri Lanka	-1.984	0.075	-2.256	0.152
Portugal	-2.090	0.084	-2.073	0.092
Canada	-2.238	0.136	-2.518	0.243
Belgium			-1.790	0.083
Kuwait			-0.239	0.014
Taiwan				
Philippines				
Poland			-1.572	0.082

Table 13: Export Elasticities - Specialization in x_{kj}, Raw x_{kji}, $\gamma = 1$, Unconstrained estimates

	Weights 1991-1996		Weights 1996-2000	
	η^X	SD	η^X	SD
Slovakia	-1.658	0.025		
Singapore	-2.414	0.040		
Egypt	-2.454	0.047		
Finland	-2.592	0.039	-2.766	0.055
Austria	-2.593	0.036	-2.915	0.048
India	-2.634	0.035	-2.309	0.034
United Kingdom	-2.671	0.038	-2.756	0.049
Turkey	-2.702	0.044	-2.410	0.035
Germany	-2.768	0.054	-3.002	0.070
France	-2.879	0.044	-3.000	0.051
China	-2.893	0.053	-2.608	0.044
Israel	-2.909	0.045	-2.734	0.042
Hungary	-2.940	0.052	-3.351	0.070
Sweden	-2.990	0.062	-3.151	0.065
United States	-3.025	0.111	-2.452	0.094
Italy	-3.053	0.036	-2.953	0.038
Australia	-3.116	0.113	-2.344	0.083
Greece	-3.236	0.052		
Spain	-3.264	0.069	-3.415	0.070
Korea	-3.374	0.059	-2.650	0.066
Japan	-3.385	0.121	-3.036	0.145
Indonesia	-3.435	0.125		
Norway	-3.484	0.072	-3.445	0.070
Portugal	-3.680	0.066	-3.718	0.067
Sri Lanka	-3.774	0.086	-4.241	0.159
Hong Kong	-4.360	0.183		
Canada	-5.291	0.334	-5.485	0.411
Venezuela	-6.671	0.070		
Belgium			-3.056	0.043
Kuwait			-0.538	0.024
Taiwan				
Philippines				
Poland			-2.574	0.038

Table 14: Export Elasticities - Specialization in x_{kj}, Raw x_{kji}, $\gamma = 2$, Constrained estimates

	Weights 1991-1996		Weights 1996-2000	
	η^X	SD	η^X	SD
Slovakia	-1.022	0.045		
Venezuela	-1.223	0.068		
Singapore	-1.269	0.044		
Australia	-1.389	0.056	-1.023	0.048
Egypt	-1.447	0.060		
India	-1.471	0.047	-1.320	0.065
United States	-1.510	0.088	-1.210	0.077
Turkey	-1.522	0.051	-1.359	0.058
Germany	-1.536	0.064	-1.646	0.068
United Kingdom	-1.547	0.055	-1.546	0.064
China	-1.577	0.065	-1.425	0.072
Finland	-1.582	0.045	-1.657	0.062
Japan	-1.629	0.061	-1.489	0.091
Hungary	-1.640	0.068	-1.786	0.094
France	-1.671	0.068	-1.688	0.074
Austria	-1.678	0.067	-1.755	0.094
Korea	-1.680	0.059	-1.295	0.063
Italy	-1.700	0.057	-1.622	0.065
Indonesia	-1.706	0.060		
Hong Kong	-1.784	0.083		
Israel	-1.788	0.069	-1.835	0.112
Sweden	-1.797	0.081	-1.807	0.083
Norway	-1.797	0.064	-1.795	0.075
Spain	-1.887	0.092	-1.927	0.096
Greece	-1.892	0.086		
Sri Lanka	-1.988	0.075	-2.266	0.152
Portugal	-2.099	0.084	-2.080	0.092
Canada	-2.260	0.136	-2.538	0.243
Belgium			-1.798	0.083
Kuwait			-0.239	0.014
Taiwan				
Philippines				
Poland			-1.576	0.082

Table 15: Export Elasticities - Specialization in x_{kj}, Raw x_{kji}, $\gamma = 2$, Unconstrained estimates

	Weights 1991-1996		Weights 1996-2000	
	η^X	SD	η^X	SD
Slovakia	-1.660	0.025		
Singapore	-2.419	0.040		
Egypt	-2.455	0.047		
Austria	-2.597	0.036	-2.918	0.048
Finland	-2.601	0.039	-2.776	0.055
India	-2.639	0.035	-2.314	0.034
United Kingdom	-2.675	0.038	-2.761	0.049
Turkey	-2.712	0.044	-2.416	0.035
Germany	-2.785	0.054	-3.019	0.070
France	-2.887	0.044	-3.010	0.051
Israel	-2.915	0.045	-2.741	0.042
China	-2.917	0.053	-2.630	0.044
Hungary	-2.942	0.052	-3.354	0.070
Sweden	-2.998	0.062	-3.160	0.065
United States	-3.058	0.111	-2.484	0.094
Italy	-3.065	0.036	-2.962	0.038
Australia	-3.125	0.113	-2.351	0.083
Greece	-3.240	0.052		
Spain	-3.272	0.069	-3.424	0.070
Korea	-3.383	0.059	-2.654	0.066
Japan	-3.412	0.121	-3.053	0.145
Indonesia	-3.449	0.125		
Norway	-3.487	0.072	-3.449	0.070
Portugal	-3.689	0.066	-3.725	0.067
Sri Lanka	-3.778	0.086	-4.252	0.159
Hong Kong	-4.397	0.183		
Canada	-5.313	0.334	-5.506	0.411
Venezuela	-6.677	0.070		
Belgium			-3.064	0.043
Kuwait			-0.538	0.024
Taiwan				
Philippines				
Poland			-2.577	0.038

Table 16: Export Elasticities - Specialization in x_{kj}, Raw x_{kji}, $\gamma = 0.5$, Constrained estimates

	Weights 1991-1996		Weights 1996-2000	
	η^X	SD	η^X	SD
Slovakia	-1.019	0.045		
Venezuela	-1.213	0.068		
Singapore	-1.261	0.044		
Australia	-1.375	0.056	-1.013	0.048
Egypt	-1.445	0.060		
United States	-1.460	0.088	-1.162	0.077
India	-1.464	0.047	-1.313	0.065
Turkey	-1.508	0.051	-1.349	0.058
Germany	-1.509	0.064	-1.620	0.068
United Kingdom	-1.540	0.055	-1.538	0.064
China	-1.542	0.065	-1.392	0.072
Finland	-1.568	0.045	-1.642	0.062
Japan	-1.588	0.061	-1.464	0.091
Hungary	-1.637	0.068	-1.781	0.094
France	-1.660	0.068	-1.674	0.074
Korea	-1.667	0.059	-1.289	0.063
Austria	-1.672	0.067	-1.751	0.094
Italy	-1.683	0.057	-1.609	0.065
Indonesia	-1.685	0.060		
Hong Kong	-1.729	0.083		
Israel	-1.779	0.069	-1.825	0.112
Sweden	-1.784	0.081	-1.794	0.083
Norway	-1.792	0.064	-1.789	0.075
Spain	-1.874	0.092	-1.914	0.096
Greece	-1.886	0.086		
Sri Lanka	-1.981	0.075	-2.251	0.152
Portugal	-2.085	0.084	-2.070	0.092
Canada	-2.227	0.136	-2.508	0.243
Belgium			-1.786	0.083
Kuwait			-0.238	0.014
Taiwan				
Philippines				
Poland			-1.571	0.082

Table 17: Export Elasticities - Specialization in x_{kj}, Raw x_{kji}, $\gamma = 0.5$, Unconstrained estimates

	Weights 1991-1996		Weights 1996-2000	
	η^X	SD	η^X	SD
Slovakia	-1.657	0.025		
Singapore	-2.411	0.040		
Egypt	-2.453	0.047		
Finland	-2.587	0.039	-2.761	0.055
Austria	-2.591	0.036	-2.913	0.048
India	-2.632	0.035	-2.306	0.034
United Kingdom	-2.668	0.038	-2.754	0.049
Turkey	-2.698	0.044	-2.406	0.035
Germany	-2.759	0.054	-2.994	0.070
France	-2.875	0.044	-2.996	0.051
China	-2.882	0.053	-2.596	0.044
Israel	-2.905	0.045	-2.730	0.042
Hungary	-2.939	0.052	-3.349	0.070
Sweden	-2.986	0.062	-3.147	0.065
United States	-3.008	0.111	-2.436	0.094
Italy	-3.047	0.036	-2.948	0.038
Australia	-3.111	0.113	-2.341	0.083
Greece	-3.234	0.052		
Spain	-3.260	0.069	-3.411	0.070
Korea	-3.370	0.059	-2.648	0.066
Japan	-3.371	0.121	-3.028	0.145
Indonesia	-3.428	0.125		
Norway	-3.482	0.072	-3.443	0.070
Portugal	-3.675	0.066	-3.715	0.067
Sri Lanka	-3.772	0.086	-4.236	0.159
Hong Kong	-4.342	0.183		
Canada	-5.280	0.334	-5.475	0.411
Venezuela	-6.667	0.070		
Belgium			-3.052	0.043
Kuwait			-0.538	0.024
Taiwan				
Philippines				
Poland			-2.573	0.038