Trade Elasticities*

Jean Imbs† Isabelle Méjean‡

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Abstract

We estimate the aggregate export and import price elasticities implied by a Constant Elasticity of Substitution (CES) demand system, for more than 30 countries at various stages of development. Trade elasticities are given by weighted averages of sector-specific elasticities of substitution, that we estimate structurally. Both weights and substitution elasticities can be chosen to compute the response of trade to specific shocks to relative prices, bilateral or global. We document considerable, significant cross-country heterogeneity in multi-lateral trade elasticities, which is virtually absent from estimates constrained to mimic aggregate data. The international dispersion in import price elasticities depends mostly on preference parameters, whereas export price elasticites vary with the composition of trade. We simulate the demand-based response of trade to specific exogenous shifts in international prices. We consider shocks to EMU-wide, US or China's relative prices, as well as country-specific shocks within the EMU zone. The trade responses to an external EMU-shock are considerably heterogeneous across member countries; in contrast, a within-EMU (Greek, Portuguese, German) shock to relative prices has largely homogeneous consequences on Eurozone trade patterns.
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PRELIMINARY AND INCOMPLETE

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†Paris School of Economics, Swiss Finance Institute and CEPR. Corresponding author: Paris School of Economics, 106-112 boulevard de l'Hopital, Paris, France 75013. jeanimbs@gmail.com, www.hec.unil.ch/jimbs

‡Ecole Polytechnique, CREST and CEPR, isabelle.mejean@polytechnique.edu, http://www.isabellemejean.com

1 Introduction

The response of traded quantities to exogenous shifts in relative prices is often used to gauge a country's external performance. Export elasticities illustrate the resilience of exporters in the face of a sudden deterioration in their position. The price elasticity of imports summarizes the competition between domestic and foreign producers in the face of an adjustment in demand. A trade elasticity is a reduced form estimate, but one that is relevant to policy - and ultimately to calibration choices.

Recent work has shown trade elasticities can reflect supply decisions on the part of individual producers. International prices differ, for instance because of tariffs or transport costs, and firms decide accordingly to enter or exit export markets. The aggregate response of trade to such relative price shocks is a trade elasticity, one that derives only from parameters on the supply side of the economy.[1] The result is striking because it takes the counterpoint to a venerable literature, that views trade elasticities as determined by end consumers' preferences.

In this paper, trade elasticities are governed by the demand side of the economy. The response of aggregate imports and exports to changes in relative prices depends on consumers' willingness to substitute domestic and foreign goods, just as it does in a venerable literature. We use a sectoral version of a conventional Constant Elasticity of Substitution (CES) demand system to motivate a parsimonious and quasi-structural estimation of trade elasticities. The price elasticity of imports is a trade-weighted average of the sectoral elasticities of substitution of the domestic consumer; the price elasticity of exports is similar, but the average is now taken both across sectors and destination markets.

We implement a now standard structural method due to Feenstra (1994) to estimate elasticities of substitution at a sector level, using disaggregated data on traded quantities and prices. We also collect the trade weights implied by a CES demand system, that are required to average sectoral substitution elasticities into aggregate trade elasticities. Both trade weights and

[1] See for instance Eaton and Kortum (2002), Chaney (2008), Dekle, Eaton and Kortum (2008), or Arkolakis, Demidova, Klenow and Rodriguez-Clare (2008).

elasticities of substitution can be chosen to reflect the nature of a specific relative price shock. A domestic shock to production costs in country A, for instance, affects the price of domestic goods relative to every competing varieties produced abroad. The price elasticities of trade are then multi-lateral, and computed using trade weights and elasticities of substitution across all destinations.

We consider specific alternatives, focusing each time on adequately chosen sub-sets of trade weights and elasticities of substitution. We simulate the trade consequences in EMU member countries of a price shock that is external to the single currency area. In this case, the trade weights used in aggregation are computed on the basis of trade between member countries and the rest of the non-EMU world. We then consider internal price shocks, like a Greek, Portuguese or German wage shock, and investigate their consequences on intra-EMU trade. In this case, the weighthing scheme reflects trade within the single currency area only. We also explore the global consequences of a shock to production prices in China (e.g. its entry into the World Trade Organization (WTO)), or in the United States. Each time we choose the appropriate parametrization reflecting the nature of the shock we consider. This flexibility - and the possibility to consider specific, non multi-lateral price shocks - is unique to our methodology.

Given parsimonious data requirements, we are able to obtain theory-implied estimates of multi-lateral import and export price elasticities for 28 countries, at various levels of development. We uncover large differences across countries. Import elasticities range from 0.7 in Hong Kong to more than 7 in China. Export elasticities range from 1.7 in Slovakia to almost 5 in Canada. Such dispersion is absent from conventional estimates of trade elasticities, even though they are implied by a CES demand system as ours are. The difference arises because our approach builds on sectoral data; the trade elasticities we compute depend directly on the specialization of trade, across both sectors and trade partners. Conventional estimates are typically obtained from aggregate data, which tend to mitigate the importance of sectoral specialization. In addition, the sectoral dimension adds econometric precision to our country-specific estimates.

Anecdotal evidence is in fact plentiful that trade elasticities are 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 resilience of German exports, focused on differentiated consumption goods. 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, does not seem to have affected trade balances identically everywhere. The measures of trade elasticities that we introduce are well suited to capture such cross-country heterogeneity; by construction they embed directly the sectoral and geographic specialization of trade.

We decompose the dispersion in trade elasticities into cross-country differences in sectoral elasticities of substitution and cross-country differences in the specialization of trade. We find differences in preferences explain the lion's share of the dispersion in import elasticities, while differences in the international and sectoral patterns of exports explain most of the cross-country variation in export elasticities. Conditional on our model - and its identifying assumptions - this suggests imports price elasticities differ across countries because of preferences, whereas export price elasticities are determined by patterns of trade. The former are therefore likely stable over time, whereas the latter will change in response to shifts in the international specialization of trade, like China's entry in world trade or the formation of regional trade areas.

Armed with cross-country sectoral estimates of substitution elasticities and trade weights, we simulate the response of trade to specific, bilateral, shocks to relative prices. We consider four experiments. An EMU-wide price shock, that does not affect internal relative prices within the zone. A shock to relative prices within the EMU zone, e.g. to Greek, Portuguese or German prices. A shock to prices in China, reflecting for instance the country's entry into the WTO. And a shock to US prices. In each case, we compute the trade responses implied by adequately chosen substitution elasticities and trade weights. For instance, to compute the response of EMU imports to a shock to external EMU prices, we consider member countries elasticities

of substitution, aggregated up using the sectoral allocation of trade with respect to non-EMU partners. To compute the response of US exports to a shock in Chinese prices, we aggregate up Chinese elasticities of substitution using the sectoral allocation of US exports towards China. In all cases, we maintain the assumption that preference parameters and trade weights are left unchanged by the considered shock in relative prices, so that the trade responses we compute ought to be interpreted as comparative statics obtained in partial equilibrium.

The trade responses to an EMU-wide price shock display considerable heterogeneity across EMU member countries. A one percent increase in relative domestic prices lowers Finnish imports by two percent, but Austrian imports by 0.5 percent only. Most country-level estimates are significantly different from the EMU-wide response, equal to 1.4 percent. We show these large differences arise because of the heterogeneous elasticities of substitution we estimate for each member country. We simulate import elasticities using the same methodology but holding substitution elasticities constant across countries. Virtually no cross-country heterogeneity subsists. Imports in EMU member countries respond to an external price shock in a heterogeneous manner. In our model, they do so mostly because preferences are heterogeneous.

The same is true of export responses, that range from 1.2 to 2.3 percent. Most country-level responses are significantly away from the EMU-wide average, equal to 1.7. For exports, however, differences across countries do not arise because of preferences. Rather, export elasticities towards non-EMU destinations differ across EMU member countries because the pattern of specialization of external EMU trade is highly country-specific.

Such results are not innocuous within a single currency area. An external shock to relative prices has vastly heterogeneous consequences on EMU member countries' exports and imports. Such heterogeneity exists only in our proposed measure of trade elasticity. It is absent from conventional estimates arising from aggregate data. In contrast, changes in relative prices within EMU have little consequences on the patterns of trade within the zone. TO BE COMPLETED

The rest of the paper is structured as follows. Section 2 reviews the vast literature on trade elasticities. Section 3 develops the model that relates trade and substitution elasticities,

discusses our structural estimation of substitution elasticities across countries, along with the data we use. Section 4 presents the estimates we obtain for multi-lateral import and export elasticities, and analyses their cross-country dispersion. Section 5 discusses the simulated trade consequences of specific bilateral price shocks. Section 6 concludes.

2 Literature

Estimating trade elasticities is an old business in economics. A venerable empirical literature goes back to at least Orcutt (1950), or Houtakker and Magee (1969). The basic specification writes:

$$\ln M_{it} = \alpha_{0i} + \alpha_{1i} \ln \frac{PM_{it}}{WPI_{it}} + u_{it}$$

where M_{it} is country i total imports, PM_{it} is an index of import prices, and WPI_{it} is the wholesale price index in i. The price elasticity of imports is given by α_{1i}. For exports, the estimations writes:

$$\ln X_{it} = \beta_{0i} + \beta_{1i} \ln \frac{PX_{it}}{PXW_{it}} + v_{it}$$

where X_{it} denotes country i total exports, PX_{it} is an index of export prices, and PXW_{it} is a world index of export prices. The parameter of interest is β_{1i}. Houtakker and Magee (1969) also include controls for domestic or world GDP, to estimate the income elasticity of imports (or exports, respectively). As is well known, the specifications arise directly from a linearized version of a one-good CES demand system. Our focus here is on the price elasticities of trade flows, and our approach remains silent on income elasticities.

These early specifications have undergone decades of econometric sophistication, surveyed in Marquez (2002). They include allowances for dynamics, differences between short and long run elasticities, the importance of heterogeneity, the stability of trade relationships, and of course endogeneity issues. Attempts to alleviate endogeneity pervade this vast empirical literature, and range from the estimation of simultaneous equations, co-integration analysis to the

instrumentation of relative price changes. See for instance Marquez (1990), Gagnon (2003), or Hooper, Johnson and Marquez (1998).

In practice, there is little cross-country evidence on trade elasticities. In their survey, Goldstein and Kahn (1985) report estimates for 15 countries, all OECD members. Marquez (2002) or Kwack et al (2007) report some estimates for 8 Asian economies, including Hong Kong, the Philippines, whereas Cheung et al (2009) estimate Chinese trade elasticities. Figure 1 displays the estimates reported by Houtakker and Magee (1969). With 15 developed importing countries and 26 exporting economies, these results may well be amongst the broadest cross section in the literature up to now. Most estimates are not significantly different from zero, often because of wide standard error bands. They are not significantly different from each other either. We in fact do not know much about the international cross section of trade elasticities.

One explanation may be econometric. Identification of trade elasticities is complicated by the potential endogeneity of traded goods prices to their quantities. In spite of a vast literature, little is available to address the issue *systematically*, in a large cross section of countries. In addition, with macroeconomic data, identification is obtained on the time dimension. Econometric power is limited accordingly, sometimes drastically. For instance, China did not release an import price index until 2005.

Estimates are therefore often imprecise, so much so that international differences are rarely significant, in spite of continuing econometric refinements. Goldstein and Kahn (1985) report values for export elasticities that, depending on the source paper, range from -2.27 to -0.34 for France, from -3.00 to -0.50 for Japan, or from -2.32 to -0.32 for the U.S. These are in fact point estimates, corresponding to different estimators. Accounting for uncertainty, it is not clear whether any of these elasticities are effectively significantly different from zero, nor indeed from each other. Marquez (2002) surveys values for the US price elasticity of imports oscillating between -4.8 and -0.3, between -0.2 and -2.8 for Canada and between 0.15 and -3.4 for Japan.

Such imprecision makes it difficult to use trade elasticities for any cross-country purposes. Trade resilience, or trade performance are effectively estimated to be the same across countries. And the calibration choices implied by trade elasticities are also made accordingly. In many a multi-country model, the international elasticity of substitution is effectively calibrated to be the same across all countries. The choice is made for lack of reliable cross-country estimates. We know little about its empirical validity.

3 Theory, Estimation, and Data

We first review the CES demand system used to derive expressions for the price elasticities of aggregate imports and exports. We then briefly describe the estimation of sectoral elasticities of substitution, introduced by Feenstra (1994). We close with a review of the data needed for this structural estimation, and for the weights used in averaging up sectoral elasticities.

3.1 Theory

We build on a Constant Elasticity of Substitution (CES) demand system, with nested 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 $i \in I_{kj}$ 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

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

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 producing 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 $\tau_{kij} > 1$ for $i \neq j$, and $\tau_{kjj} = 1$. Utility maximization implies that demand for variety i in each sector k is given by

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

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 how aggregate quantities respond to changes in aggregate international relative prices. We compute the response of trade to a shock affecting all relative prices in country j, across all sectors k and all partners i. We later consider the response of trade to specific, bilateral, price shocks. Let η_{kj}^M (η_{kj}^X) denote the response of country j's sectoral imports (exports),

and η_j^M (η_j^X) their aggregate counterparts. By definition:

$$\begin{aligned}\eta_j^M &\equiv \frac{\partial \ln \sum_k \sum_{i\neq j} P_{kij}C_{kij}}{\partial \ln\{P_{kij}/P_{kjj}\}_{\forall k,i\neq j}} \equiv \sum_k m_{kj}\eta_{kj}^M \\ \eta_j^X &\equiv \frac{\partial \ln \sum_k \sum_{i\neq j} P_{kji}C_{kji}}{\partial \ln\{P_{kji}/P_{kii}\}_{\forall k,i\neq j}} \equiv \sum_k x_{kj}\eta_{kj}^X\end{aligned}$$

where $\eta_{kj}^M = \frac{\partial \ln \sum_{i\neq j} P_{kij}C_{kij}}{\partial \ln\{P_{kij}/P_{kjj}\}_{\forall k,i\neq j}}$ and $\eta_{kj}^X = \frac{\partial \ln \sum_{i\neq j} P_{kji}C_{kji}}{\partial \ln\{P_{kji}/P_{kii}\}_{\forall k,i\neq j}}$, and weights are 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.

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

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

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:

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

The response of aggregate imports 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, which has three elements. The first term, largest in magnitude, involves an import-weighted average of σ_{kj}. The other two reflect the composition of sectoral trade; both are smaller in magnitude than the first summation. The parameter γ_j has a level effect on η_j^M, through the second and third summations in equation (3).

The price elasticity of exports depends on the elasticities of substitution country j faces in all exporting destinations. 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

$$\begin{aligned}
\eta_{jk}^X &= \sum_{i\neq j} x_{kji}\left[(1-\sigma_{ki})\frac{\partial \ln P_{kji}/P_{kii}}{\partial \ln P_{kji}/P_{kii}} + (\sigma_{ki}-\gamma_i)\frac{\partial \ln P_{ki}/P_{kii}}{\partial \ln P_{kji}/P_{kii}} + (\gamma_i - 1)\frac{\partial \ln P_i/P_{kii}}{\partial \ln P_{kji}/P_{kii}}\right] \\
&= \sum_{i\neq j} x_{kji}\left[(1-\sigma_{ki}) + (\sigma_{ki}-\gamma_i)w_{kji} + (\gamma_i - 1)\sum_k w_{ki}w_{kji}\right] \qquad (4)
\end{aligned}$$

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

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

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 x_{kj}, and the share of importing country i in j's exports, x_{kji}. Equations (5) has three components: an adequately weighted average of σ_{ki}, and two terms, smaller in magnitude, that reflect the specialization of trade. These involve γ_i, which we calibrate.

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_j^X and η_j^M are 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 the structural estimation of the preference parameter σ_{ki}, across sectors k and countries i.

3.2 Estimation

Following Feenstra (1994), we identify σ_{ki} using the cross-section of traded quantities and prices across exporters selling goods to each considered destination. This is possible thanks to the multilateral dimension of disaggregated trade data.[2] Demand is given in equation (1), which

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

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. Following Feenstra (1994), we 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.[3]

Kemp (1962) argued using expenditure share $s_{kijt} = \frac{P_{kijt}C_{kijt}}{P_{kjt}C_{kjt}}$ tends to alleviate measurement error. Rewrite demand as

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

We do not observe domestically produced consumption, and prices are measured Free on Board. Let tilded variables denote the observed counterparts to theory-implied prices and quantities. We observe $\tilde{P}_{kijt} \equiv P_{kijt}/\tau_{kijt}$. 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 and first differences, demand becomes

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

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

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

taste shocks. 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.

Under an orthogonality assumption between 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 $Y_{kit} = (\Delta \ln \tilde{P}_{kijt} - \Delta \ln \tilde{P}_{krjt})^2$, $X_{1kijt} = (\Delta \ln \tilde{s}_{kijt} - \Delta \ln \tilde{s}_{krjt})^2$, $X_{2kijt} = (\Delta \ln \tilde{s}_{kijt} - \Delta \ln \tilde{s}_{krjt})(\Delta \ln \tilde{P}_{kijt} - \Delta \ln \tilde{P}_{krjt})$ and $u_{kijt} = (\varepsilon_{kijt} - \varepsilon_{krjt})(\delta_{kijt} - \delta_{krjt}) \frac{(\sigma_{kj}-1)(1+\omega_{kj})}{1+\omega_{kj}\sigma_{kj}}$. Feenstra (1994) showed that, in a CES demand system, X_{1kijt} and X_{2kijt} can be instrumented by their time averages, which averages away demand shocks. Identification is therefore based on the cross-sectional dimension of equation (8). We also include an observation-specific intercept to account for measurement error, correct the estimation for heteroskedasticity across exporters i, and include Common Correlated Effects to avoid double counting of macroeconomic influence at the sector level.[4]

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

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

[4]See Imbs and Méjean (2009) for details.

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.[5]

3.3 Data

A structural estimate of $\hat{\sigma}_{kj}$ requires that we observe the cross-section of imported quantities and unit values at the sector level, and for all countries j. We use the trade database BACI, released by CEPII, that harmonizes UN-ComTrade export and import declarations. The data trace 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.

Thanks to the multilateral dimension of our data, we are able to estimate $\hat{\sigma}_{kj}$ 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

[5]The appendix details the computation of standard deviations. As is apparent, there are combinations of estimates in equation (8) 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 (8) over the intervals of admissible values of the elasticities.

summary statistics for the 28 countries we have data for. The number of sectors ranges from 10 to 27. 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 20 in Guatemala to more than 50 in Homg Kong. For each sector, our data implies an average number of exporting countries of 28.

The main data constraint 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.

To compute the latter three weights, we require information on domestic consumption at sectoral level, across as large a cross-section of countries as possible. This is absent from conventional international trade databases, and raises issues of concordance since we require information on both production and trade at the sectoral level. In order to maximize 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. But then we continued to use 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.

In general, UNIDO data report nominal sectoral output at the 3-digit ISIC (revision 2) level. 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 Panama 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. 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 recently increased in magnitude. 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.

4 Trade Elasticities

We first present cross-country estimates of import elasticities. We then turn to exports. In both instances, we discuss the determinants of the cross-country dispersion in estimates. We close with a comparison with what is implied by parameters constrained to homogeneity across sectors.

4.1 Imports Price Elasticities

Figure 2 reports import elasticities for the 28 countries estimates are available, ordered by increasing absolute value. The values are computed imposing $\gamma = 1$, and using weights computed over 1991-1995. The estimates range from -0.7 in Hong Kong to more than -7 for China. All elasticities are significantly different from zero, and most are also significantly different from each other. Most countries have estimates between -3 and -5.

China, India, or Turkey all have import elasticities below -5, whereas rich developed countries tend to have estimates larger than -4. Large, emerging economies tend to have high import elasticities, whereas OECD member economies are closer to zero. Exceptions are Japan and the US, both with elasticities below -4.

Such cross country dispersion may reflect the specialization of imports across sectors, and/or the preferences of the importing representative consumer. From equation (3), we can begin to investigate the theoretical importance of one versus the other in determining the value of η_j^M.

Define $\Theta_j = \sum_k m_{kj}\left[(1-\sigma_{kj}) + (1-w_{kjj})(\sigma_{kj}-1)\right]$. Simple algebra implies

$$\frac{\partial \ln \eta_j^M}{\partial \ln m_{kj}} = \frac{\partial \ln \eta_j^M}{\partial \ln w_{kjj}} = -\frac{1}{\Theta_j} m_{kj}(\sigma_{kj}-1)w_{kjj}$$

and

$$\frac{\partial \ln \eta_j^M}{\partial \ln \sigma_{kj}} = -\frac{1}{\Theta_j} m_{kj}\sigma_{kj}w_{kjj}$$

By definition, cross-country differences in σ_{kj} are likely to have first-order effect on the dispersion of η_j^M. The weights m_{kj} and w_{kjj} have theoretically smaller (and identical) effects.

This conclusion holds for given cross-country dispersion in m_{kj}, σ_{kj}, and w_{kjj}. We now decompose the variance of all three determinants of η_j^M into a cross-country and a cross-sector component. For $c_{kj} \in \{\sigma_{kj}, m_{kj}, w_{kjj}\}$, we compute

$$\underbrace{E(c_{jk}-\bar{c})^2}_{Overall} = \underbrace{E(c_{jk}-\bar{c}_k)^2}_{Cross-country} + \underbrace{E(\bar{c}_k-\bar{c})^2}_{Cross-sector}$$

where $\bar{c} = \frac{1}{N}\sum_j \sum_k c_{jk}$ and $\bar{c}_k = \frac{1}{N_k}\sum_j c_{jk}$. The decomposition is performed in Table 3. More than 80 percent of the variance in σ_{kj} originates in cross-country dispersion; in contrast, the cross-country dimension explains 74 percent of the variance in m_{kj} and 39 of that in w_{kjj}. Given that σ_{kj} has first order theoretical effect on import elasticities, our data are strongly suggestive that most of the cross-country dispersion in η_j^M stems from international differences in σ_{kj}. The result is important, for it suggests a contrario that import elasticities are resilient to shifts in the international specialization of trade.

Table 2 presents the cross-country estimates of η_j^M that underpin Figure 2, along with the number of sectors used in each country. For robustness, the Table also reports the elasticity estimates obatined if m_{kj} and w_{kjj} are computed over 1996-2000 instead of the first half of the 1990's. This is an important robustness check, since the assumption of constant weights in the face of shocks to relative prices is implausible in theory. Reassuringly, changing the period over which trade weights are computed has little impact on the end estimates of η_j^M. The ranking of

countries and the range of estimates both remain largely unchanged. Such robustness reflects the well known persistence in trade patterns.

4.2 Export Price Elasticities

Figure 3 reports our estimates of export elasticities for the 28 countries data are available, ordered by increasing absolute value. The results correspond to $\gamma = 1$, and export weights computed over 1991-1995. Estimates of η_j^X range from -1.7 in Slovakia to more than -4.5 in Canada and Hong Kong. All are estimated precisely, significantly different from zero, and significantly different from each other in most instances.

Germany, Finland, Austria or the UK all have export elasticities relatively close to zero, with values around -3; Taiwan, Canada or Hong Kong lay at the top of the distribution we estimate, with values between -4.5 and -5. Most countries have estimates between -3 and -4. Large economies, like the US, China or Japan have average estimates, around -4. Such dispersion can correspond to the international pattern of exports, as η_j^X depends not only on the sectoral allocation of trade, but also its specialization across destination markets.

From equation (5), we can once again investigate the relative importance of the various determinants of η_j^X. In particular, define $\Xi_j = \sum_k x_{kj} \sum_{i \neq j} x_{kji} \left[(1 - \sigma_{ki}) + w_{kji}(\sigma_{ki} - 1)\right]$. Simple algebra implies:

$$\begin{aligned}
\frac{\partial \ln \eta_j^X}{\partial \ln \sigma_{ki}} &= -\frac{1}{\Xi_j} x_{kj} x_{kji} \sigma_{ki} (1 - w_{kji}) \\
\frac{\partial \ln \eta_j^X}{\partial \ln x_{kj}} &= -\frac{1}{\Xi_j} x_{kj} \sum_{i \neq j} x_{kji} \left(1 - w_{kji}\right) (\sigma_{ki} - 1) \\
\frac{\partial \ln \eta_j^X}{\partial \ln x_{kji}} &= -\frac{1}{\Xi_j} x_{kj} x_{kji} \left(1 - w_{kji}\right) (\sigma_{ki} - 1) \\
\frac{\partial \ln \eta_j^X}{\partial \ln w_{kji}} &= \frac{1}{\Xi_j} x_{kj} x_{kji} w_{kji} (\sigma_{ki} - 1)
\end{aligned}$$

Since it is proportional to x_{kj}, it is $\frac{\partial \ln \eta_j^X}{\partial \ln x_{kj}}$ that takes the largest absolute value amongst these

four partial derivatives. The three other terms are indeed proportional to $x_{kj} \cdot x_{kji}$, which is by definition smaller. This suggests it is the sectoral composition of exports from country j that determines most of its export elasticity. One can actually show that, for small enough values of w_{kji}, it is $\frac{\partial \ln \eta_j^X}{\partial \ln \sigma_{ki}}$ that comes second.[6] In other words, the characteristics of demand in destination countries, and the geographical allocation of exports both have only second-order consequences on the resilience of exports to price shocks.

Such ranking takes as given the cross-country dispersion in all four determinants of η_j^X. We perform a variance decomposition of σ_{ki}, x_{kj}, x_{kji}, and w_{kji} following an approach analogous to the previous section. For $c_{kj} \in \{\sigma_{ki}, x_{kj}, x_{kji}, w_{kji}\}$, we compute

$$\underbrace{E(c_{kji} - \bar{c})^2}_{Overall} = \underbrace{E(c_{kji} - \bar{c}_{kj})^2}_{Cross-destination} + \underbrace{E(\bar{c}_{kj} - \bar{c}_k)^2}_{Cross-origin} + \underbrace{E(\bar{c}_k - \bar{c})^2}_{Cross-sector}$$

where $\bar{c} = \frac{1}{N} \sum_k \sum_j \sum_i c_{kji}$, $\bar{c}_{kj} = \frac{1}{N_{kj}} \sum_i c_{kji}$, and $\bar{c}_k = \frac{1}{N_k} \sum_j \sum_i c_{kji}$. The decomposition is performed in Table 5. Interestingly, 70 percent of the variance in x_{kj} is across countries. Since this weight has first order effects on η_j^X, the bulk of the cross-country dispersion in export elasticities does in fact correspond to the cross-sector allocation of trade. The Table also confirms the finding that most of the dispersion in σ_{ki} is across destination countries, with 84 percent of total variance. But that has only second order consequences on the aggregate elasticity, because exporting countries face identical dispersion in σ_{ki} across destination markets. The fact export elasticities vary with the sectoral allocation of trade is important. It suggests that, contrary to imports, the resilience of exports to price shocks does depend on the sectoral specialization of exported goods, and not only on their substitutability.

Table 4 reports the cross-country estimates of η_j^X that underpin Figure 3, along with the values implied by trade weights computed over the late 1990's. Once again, the end estimates are virtually unchanged for those countries where they are available for both periods. Chile becomes the country with the lowest export elasticity.

[6] For $w_{kji} < \frac{1}{2}$, we have $\left|\frac{\partial \ln \eta_j^X}{\partial \ln \sigma_{ki}}\right| > \left|\frac{\partial \ln \eta_j^X}{\partial \ln x_{kji}}\right| > \left|\frac{\partial \ln \eta_j^X}{\partial \ln w_{kji}}\right|$

4.3 Homogeneity and Conventional Trade Elasticities

If estimates of η_j^M and η_j^X are to replicate aggregate data, then estimates of $\hat{\sigma}_j$ constrained to homogeneity across sectors are of the essence. Identification issues notwithstanding, the elasticity of imports (or exports) 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.[7]

But constraining $\hat{\sigma}_{kj} = \hat{\sigma}_j$ acts to reduce the information content in the corresponding aggregate trade elasticities. We showed import elasticities differ across countries mostly because $\hat{\sigma}_{kj}$ do so. In fact, the bulk of the cross-country dispersion in $\hat{\sigma}_{kj}$ is at the sector level: a given sectoral elasticity tends to take different values across countries.[8] Import elasticities computed on the basis of constrained estimates of $\hat{\sigma}_j$ will tend to display less cross-country dispersion. In addition, trade weights mechanically affect less estimates of η_j^M based on $\hat{\sigma}_j$. Consider equation (3), rewritten with the assumption that $\hat{\sigma}_{kj} = \hat{\sigma}_j$ for all k:

$$\bar{\eta}_j^M \equiv (1 - \sigma_j) + (\sigma_j - \gamma_j) \sum_k m_{kj}(1 - w_{kjj}) + (\gamma_j - 1) \sum_k w_{kj}(1 - w_{kjj})$$

The sectoral specialization of imports enters only in the second and third terms, which tend to be small in magnitude.

The cross-country dispersion in estimates of $\bar{\eta}_j^M$ is mechanically smaller than what is implied by the unconstrained estimates described in the previous section. Aggregate data tend to dampen the differences in import elasticities across countries. This is apparent from Figure

[7] To be precise, we show in Imbs and Méjean (2009) that aggregate data in fact imply $\hat{\sigma}_{kj} = \hat{\sigma}_j = \gamma_j$, 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.

[8] We regressed $\hat{\sigma}_{kj}$ on country and sector fixed effects. The two taken together explain barely 30 percent of the variance in $\hat{\sigma}_{kj}$. Therefore, the vast majority of the variance in $\hat{\sigma}_{kj}$ corresponds to sector-specific differences across countries.

4, where we report cross-country estimates of $\bar{\eta}_j^M$. The dispersion of estimates reduces considerably, with values now ranging between -0.3 and -3.3, and most estimates smaller than -2 in absolute value. While all estimates continue to be significantly different from zero, they are not always distinguishable from each other. Standard error bands are wider for constrained estimates.[9] The results in Figure 4 are consistent with the fact that cross-country aggregate data tend to yield values for import elasticities that are close to zero. But our estimates of $\bar{\eta}_j^M$ continue to be precise enough that they are significantly different from zero. In that sense, our replication of estimates based on aggregate data, albeit imperfect, dominates what would be obtained from the conventional time series estimation described in Section 2.

Constraining elasticities of substitution to homogeneity is likely to have consequences on export elasticities as well, though perhaps somewhat dampened. Substituting $\hat{\sigma}_{kj} = \hat{\sigma}_j$ for all k in equation (5) implies

$$\bar{\eta}_j^X \equiv \sum_k x_{kj} \sum_{i \neq j} x_{kji} \left[(1 - \sigma_i) + (\sigma_i - \gamma_i) w_{kji} + (\gamma_i - 1) \sum_k w_{ki} w_{kji} \right]$$

The theoretical difference between $\bar{\eta}_j^X$ and η_j^X is difficult to quantify in theory. Figure 5 presents cross country estimates of $\bar{\eta}_j^X$. The dispersion is reduced, as well. All estimates lie between -0.8 and -1.9, and in fact most values are close to -1.5. They are all significantly different from zero - but in most instances not significantly different from each other. Once again, our estimates of export elasticities that are meant to replicate what aggregate data imply, do fall in the same ballpark as what the literature has uncovered over the decades. But precision is improved relative to conventional time-series estimates.

5 Comparative Statics

We have considered shocks affecting all relative prices, across all sectors but also all trade partners. For each country in our sample, we have considered a shift in domestic costs fully

[9] See the Appendix for an intuition.

passed through into prices, changing international prices with respect to all other countries. We now consider alternative experiments, involving specific sub-groups of countries. The nature of the shock conditions the specific elasticities of substitution and weights that we choose to compute η_j^M and η_j^X. We consider four exercises. First, we simulate the trade consequences on member countries of an EMU-wide shock to relative prices, that leaves within-EMU relative prices unchanged. For this purpose, we average up elasticities of substitution using weights reflective of the trade patterns between EMU member countries and non-EMU partners. Second, we simulate shocks to within-EMU relative prices. We consider shocks to Greek, Portuguese and German prices. We evaluate the consequences of each shock on trade within the area, once again using only the pertinent within-EMU trade weights. Third, we consider a shock to Chinese relative prices, and its consequence on world imports and exports. Fourth and finally, we consider a shock to US relative prices.

The approach is akin to comparative statics. The patterns of trade are assumed invariant to the price shock. Such is not the case in general equilibrium, where new developments on relative costs, for instance, have consequences on the entry or exit decision of firms, and ultimately on the pattern of trade. Mechanisms like this are central to the literature modeling the supply and exporting decisions of heterogeneous firms, pioneered by Melitz (2003) or Eaton and Kortum (2002). They are assumed away here, just as they were in a venerable literature where trade patterns are determined by demand. We propose to improve on this literature, introducing a sectoral dimension to (demand-based) estimates of trade elasticities. This has two interesting consequences. First, we introduce measures of trade elasticities that vary across countries because of the specialization of trade. That is largely absent from the (demand based) conventional estimates obtained from macroeconomic data. Second, we are able to replicate the (demand based) trade elasticity estimates implied by aggregate data. But we achieve econometric precision thanks to the sectoral dimension.

5.1 An EMU-wide price shock

We first consider the case of a permanent shock to euro-zone costs, that is fully passed-through into prices, but leaves relative prices within the EMU unchanged. Overall EMU exports and imports respond as predicted by a conventional demand system. But the intensity of the response varies across countries according to the heterogeneous estimates we uncover. In particular, the country level response differs as the specialization of trade varies across EMU member countries, sometimes drastically. A country that trade mostly within EMU will not be affected by the shock we consider. As we will show, conventional estimates of trade elasticities, based on aggregate data, will largely miss such heterogeneous response.

We adapt equations (3) and (5) to the shock considered. We introduce import weights $m_{kj}^{NEMU} \equiv \frac{\sum_{i \notin EMU} P_{kij} C_{kij}}{\sum_k \sum_i P_{kij} C_{kij}}$, the share of extra-EMU imports of good k in country j's aggregate imports, and w_{kj}^{NEMU} the share of products from non-EMU countries in country j's consumption of good k. From equation (3), the response of country j's aggregate imports to the price shock is

$$\eta_j^M = \sum_k m_{kj}^{NEMU}(\sigma_{kj} - 1) - \sum_k m_{kj}^{NEMU}(1 - w_{kjj})(\sigma_{kj} - \gamma_j) - (\gamma_j - 1)\sum_k w_{kj} w_{kj}^{NEMU} \quad (9)$$

By analogy, the price elasticity of non-EMU exports is given by a slightly amended version of equation (5):

$$\eta_j^X = \sum_k x_{kj} \left[\sum_{l \notin EMU} x_{kjl}(1 - \sigma_{kl}) + \sum_{l \notin EMU} x_{kjl} w_{kjl}(\sigma_{kl} - \gamma_l) + \sum_{l \notin EMU} x_{kjl}(\gamma_l - 1) \sum_k w_{kl} w_{kjl} \right] \quad (10)$$

where exporting destinations are limited to non-EMU member countries.

Figure 6 reports the import and export elasticities implied by equations (9) and (10) for the EMU member countries we observe. The Figure also plots EMU-wide elasticities, which are aggregated across member countries:

$$\eta^M_{EMU} = \sum_{j \in EMU} m^{EMU}_j \eta^M_j$$
$$\eta^X_{EMU} = \sum_{j \in EMU} x^{EMU}_j \eta^X_j$$

where m^{EMU}_j and x^{EMU}_j measure the share of country j in overall EMU imports and exports.

The left panel of Figure 6 reports elasticity estimates computed with sector-specific values for σ_{kj}. There is considerable heterogeneity across the import responses of EMU countries. The extreme values of η^M_j suggest a 1 percent shock to external EMU prices has an effect on Finnish imports that is four times its value for Austria. German or Spanish imports, in turn, are three times more responsive. Such different responses are striking within a single currency area. Most countries in the left panel of Figure 6 display import responses that differ significantly from the EMU-wide adjustment.

What drives such heterogeneity? In the right panel of Figure 6, we have re-computed import elasticities according to equations (9) and (10), but now we have used for all countries the estimates of σ_{kj} we obtained for Finland.[10] Strikingly, the dispersion in import elasticities is virtually unchanged even once all countries are endowed with Finnish substitution elasticities. The result suggests the main reason why EMU countries have heterogeneous import responses to an EMU wide price shock is the specialization of trade. Even though elasticities of substitution are different (and drive most of the cross-country dispersion in multi-lateral η^M_j), within EMU the key driver of heterogeneous import elasticities is the pattern of trade.

The cross-country dispersion is smaller for η^X_j. Export elasticities also take highest value for Finland, -2.2 percent, and lowest in Austria, -1.3 percent. Exports in Portugal and France are relatively responsive to external EMU-wide shocks, whereas Spanish or German exports are more resilient. Six of the eight estimates of export elasticities are between -1.5 and -2. The

[10]There is one sector that is not traded by Finland, but is by other EMU member countries. We experimented with simply dropping this sector, or keeping the country estimates. The results presented in the paper follow the former approach. They are virtually identical using the latter.

lower right panel of Figure 6 shows the cross-section of η_j^X that are implied by equation (10), but endowing all countries with the values for x_{kj} we compute for Finland.[11] The dispersion is virtually unchanged, and continue to be substantially lower than for import elasticities.

Figure 7 reports the estimates of trade elasticities built from a single value for σ_j, constrained to homogeneity across sectors. The left panel of the Figure illustrates how cross-country dispersion is reduced, with values for $\bar{\eta}_j^M$ranging from -0.7 to -0.2, and from -1 to -0.5 for $\bar{\eta}_j^X$. Relative to unconstrained values, the country ranking is virtually identical. In other words, the use of aggregate data misses entirely on the heterogeneous response of trade across EMU member countries to an EMU wide price shock.

5.2 Within-EMU price shock

We now consider shocks to relative prices within EMU. We simulate the trade responses within the single currency area of three specific shocks: a Greek, a Portuguese and a German shift in relative prices. The first two concern small economies whose external balances are of topical interest; the last one concerns the biggest economy in the Union. Let $I = \{Greece, Portugal, Germany\}$ index the EMU economy whose prices are shocked. The import elasticity in EMU country $j \neq I$ is given by

$$\eta_{Ij}^M \equiv \sum_k m_{kj}^I (1 - \sigma_{kj}) + \sum_k m_{kj}^I w_{kIj} (\sigma_{kj} - \gamma_j) + (\gamma_j - 1) \sum_k w_{kj} w_{kIj} \tag{11}$$

where $m_{kj}^I \equiv (P_{kIj} C_{kIj})/(\sum_k P_{kIj} C_{kIj})$ is the share of sector k in country j's imports from country I, and $w_{kIj} \equiv (P_{kIj} C_{kIj})/(\sum_i P_{kij} C_{kij})$ is the share of country I in country j's consumption of good k. This is now a bilateral elasticity of j's imports from the shocked country I. An increase in country I's relative prices is going to reduce other countries' aggregate imports, the effect being stronger as the share of country I in that country's imports is larger.

[11] We continue to omit the one sector exported by some EMU countries, but not by Finland. It makes virtually no difference if the values of x_{kj} in that sector are kept instead for all concerned countries.

The same increase in country I's prices affect exports in all other countries, as their relative competitiveness increases. Bilateral exports to I increase, and the elasticity of exports to I in country j is given by

$$\eta_{Ij}^{X} \equiv \sum_{k} x_{kj}^{I} \left[(\sigma_{kI} - 1) - (\sigma_{kI} - \gamma_{I}) w_{kjI} - (\gamma_{I} - 1) \sum_{k} w_{kI} w_{kjI} \right] \tag{12}$$

where x_{kj}^{I} is the share of goods k in country j's exports to country I, and w_{kjI} is the share of good k produced in j in country I's consumption. This is a bilateral elasticity of country j's exports to I.

5.3 China and the US

In this section, we consider bilateral elasticities in response to two specific, country-level, price shocks. In particular, we compute the worldwide consequences of permanent shocks to costs - fully passed through in prices - in the two largest economies, the US and China. Theoretically, the implied elasticities are given by equations (11) and (12), defined in the previous section, with $I = \{US, China\}$. h

6 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 compute trade elasticity estimates for 28 countries, including most developed and the major developing economies. We find susbtantial cross-country dispersion, which is robust to alternative measurement strategies. Such dispersion is absent from conventional estimates of trade elasticities, obtained from aggregate time series.

Most of the cross-country dispersion in import elasticities can be explained by heterogeneous sectoral elasticities of substitution. This is a preference parameter in our model and in our identification, which therefore imply import elasticities are an empirical object likely to remain stable in the face of shocks to the pattern of international trade. Most of the cross-country dispersion in export elasticities can be explained by the sectoral specialization of exports. The pattern of exports does respond to large changes in the structure of international trade, and thus so do export elasticities.

We perform four exercises in comparative statics. We first consider a shock to EMU-wide prices. On the basis of the demand system we estimate, we compute the trade response in individual member countries. We find considerable heterogeneity in the response of imports, which corresponds to differences in the sectoral specialization of trade. The response of exports, in contrast, displays less heterogeneity within EMU. TO BE COMPLETED.

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8 Appendix: Variances

$$\hat{\eta}_j^M = \eta_j^M + \sum_k \frac{\partial \eta_j^M}{\partial \sigma_{kj}} (\hat{\sigma}_{kj} - \sigma_{kj}) \tag{13}$$

$$\Rightarrow \quad Var\ \hat{\eta}_j^M = \sum_k m_{kj}^2 w_{kjj}^2 Var\ \hat{\sigma}_{kj} \tag{14}$$

$$\hat{\bar{\eta}}_j^M = \bar{\eta}_j^M + \frac{\partial \bar{\eta}_j^M}{\partial \sigma_j} (\hat{\sigma}_j - \sigma_j) \tag{15}$$

$$\Rightarrow \quad Var\ \hat{\bar{\eta}}_j^M = \left[\sum_k m_{kj} w_{kjj} \right]^2 Var\ \hat{\sigma}_j \tag{16}$$

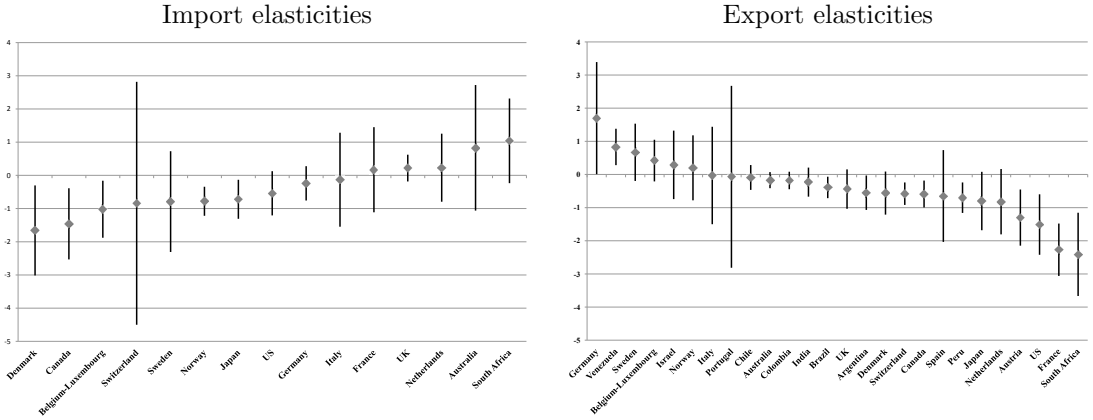
$$\hat{\eta}_j^X = \eta_j^X + \sum_k \sum_{i \neq j} \frac{\partial \eta_j^X}{\partial \sigma_{ki}} (\hat{\sigma}_{ki} - \sigma_{ki}) \tag{17}$$

$$\Rightarrow \quad Var\ \hat{\eta}_j^X = \sum_k \sum_{i \neq j} x_{kj}^2 x_{kji}^2 (1 - w_{kji})^2 Var\ \hat{\sigma}_{ki} \tag{18}$$

$$\hat{\bar{\eta}}_j^X = \bar{\eta}_j^X + \sum_k \sum_{i \neq j} \frac{\partial \eta_j^X}{\partial \sigma_i} (\hat{\sigma}_i - \sigma_i) \tag{19}$$

$$\Rightarrow \quad Var\ \hat{\bar{\eta}}_j^X = \sum_{i \neq j} \left[\sum_k x_{kj} x_{kji} (1 - w_{kji}) \right]^2 Var\ \hat{\sigma}_i \tag{20}$$

Figure 1: Houtakker and Magee (1969) elasticity estimates



Note: The grey circles are the point estimates found in Houtakker and Magee (1969). Lines around the cirecles correspond to the confidence interval, at the 5% level.

Figure 2: Distribution of unconstrained import elasticities ($\gamma = 1$)

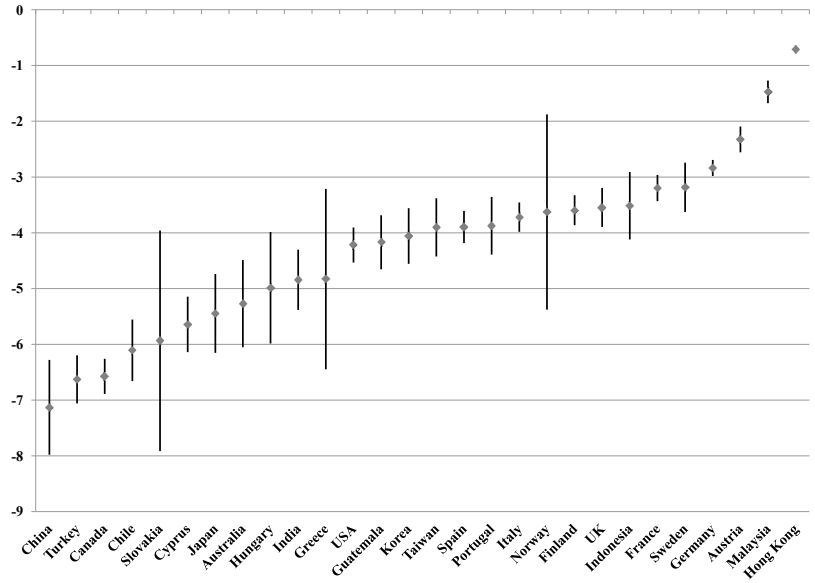


Table 1: Summary Statistics

	# sect	# sect×exp	% Trade
Australia	17	569	47.1
Austria	24	562	71.7
Canada	24	562	64.5
Chile	17	569	33.5
China	20	566	51.2
Cyprus	18	568	24.4
Finland	26	560	65.4
France	26	560	78.4
Germany	21	565	50.8
Greece	17	569	42.8
Guatemala	18	568	36.9
Hong Kong	11	575	16.9
Hungary	19	567	47.1
Indonesia	15	571	42.5
Italy	25	561	72.6
Japan	26	560	61.1
Korea	26	560	58.6
Malaysia	18	568	50.4
Taiwan	20	566	40.1
Norway	20	566	49.8
Portugal	22	564	62.3
India	18	568	33.7
Slovakia	10	576	27.9
Spain	26	560	73.3
Sweden	25	561	72.9
Turkey	24	562	57.5
United Kingdom	26	560	81.1
United States	27	559	74.3

Table 2: Unconstrained import elasticities, $\gamma = 1$

	Weights 1991-1996			Weights 1996-2000		
	# sect.	η^M	SD	# sect.	η^M	SD
Australia	17	-5.270	0.399	19	-4.087	0.334
Austria	24	-2.325	0.118	19	-1.915	0.085
Belgium				13	-1.979	0.080
Canada	24	-6.572	0.161	24	-4.602	0.116
Chile	17	-6.108	0.281	17	-5.116	0.277
China	20	-7.129	0.434	18	-7.416	0.839
Cyprus	18	-5.642	0.254	18	-4.569	0.243
Finland	26	-3.595	0.137	22	-3.475	0.199
France	26	-3.198	0.120	26	-3.133	0.145
Germany	21	-2.836	0.073	22	-3.525	0.286
Greece	17	-4.830	0.825			
Guatemala	18	-4.169	0.247			
Hong Kong	11	-0.714	0.036			
Hungary	19	-4.985	0.510	15	-2.418	0.248
India	18	-4.844	0.275	18	-4.874	0.267
Indonesia	15	-3.515	0.308			
Italy	25	-3.720	0.134	26	-3.889	0.151
Japan	26	-5.446	0.361	26	-5.191	0.341
Korea	26	-4.058	0.254	26	-3.764	0.213
Malaysia	18	-1.473	0.103	15	-2.862	0.226
Norway	20	-3.628	0.893	23	-4.196	1.289
Poland				24	-3.638	0.249
Portugal	22	-3.875	0.264	26	-3.552	0.255
Slovakia	10	-5.938	1.009	11	-3.246	0.506
Spain	26	-3.896	0.147	26	-3.485	0.131
Sweden	25	-3.186	0.225	21	-3.244	0.226
Taiwan	20	-3.903	0.266			
Turkey	24	-6.628	0.220	21	-5.333	0.173
United Kingdom	26	-3.545	0.178	24	-3.175	0.18
USA	27	-4.218	0.160	26	-4.020	0.152

Table 3: Decomposition of the Determinants of η^M

	Cross-country	Cross-sector
σ_{kj}	.842	.158
m_{kj}	.258	.742
w_{kjj}	.612	.388

Table 4: Unconstrained Export Elasticities - $\gamma = 1$

	Weights 1991-1996		Weights 1996-2000	
	η^X	SD	η^X	SD
Australia	-3.832	0.087	-3.013	0.065
Austria	-3.167	0.042	-3.906	0.175
Belgium				
Canada	-4.825	0.254	-4.984	0.251
Chile	-2.665	0.053	-2.172	0.056
China	-3.533	0.073	-3.373	0.081
Cyprus	-4.441	0.092	-3.107	0.045
Finland	-3.229	0.066	-3.393	0.092
France	-3.493	0.059	-3.821	0.088
Germany	-3.239	0.061	-3.803	0.092
Greece	-4.153	0.076		
Guatemala	-2.745	0.111		
Hong Kong	-4.523	0.169		
Hungary	-3.072	0.050	-3.175	0.065
India	-3.065	0.058	-2.771	0.054
Indonesia	-3.497	0.057		
Italy	-3.818	0.053	-3.753	0.059
Japan	-4.000	0.100	-3.602	0.107
Korea	-4.211	0.074	-3.401	0.080
Malaysia	-3.467	0.126	-2.318	0.046
Norway	-4.299	0.190	-4.378	0.162
Poland				
Portugal	-4.257	0.096	-4.872	0.145
Slovakia	-1.743	0.026	-2.738	0.082
Spain	-3.901	0.118	-4.387	0.151
Sweden	-3.624	0.080	-3.998	0.125
Taiwan	-4.460	0.103		
Turkey	-2.653	0.046	-2.577	0.053
United Kingdom	-3.225	0.049	-3.561	0.077
United States	-4.285	0.090	-3.747	0.086

Table 5: Decomposition of the Determinans of η^X

	Cross-destination	Cross-origin	Cross-sector
σ_{kji}	.840	.001	.158
x_{kj}		.698	.302
x_{kji}	.970	.021	.009
w_{kji}	.635	.339	.026

Figure 3: Distribution of unconstrained export elasticities ($\gamma = 1$)

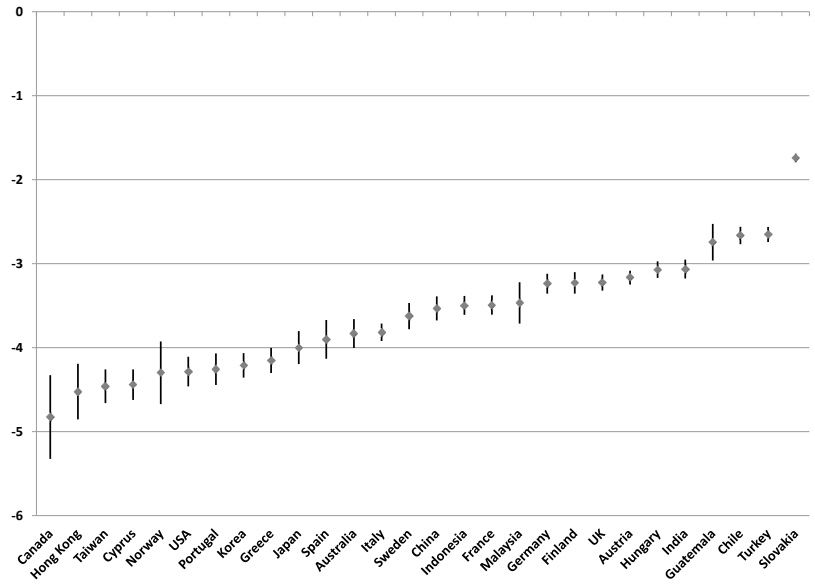


Figure 4: Distribution of constrained import elasticities ($\gamma = 1$, $\sigma_{kj} = \sigma_j, \ \forall k$)

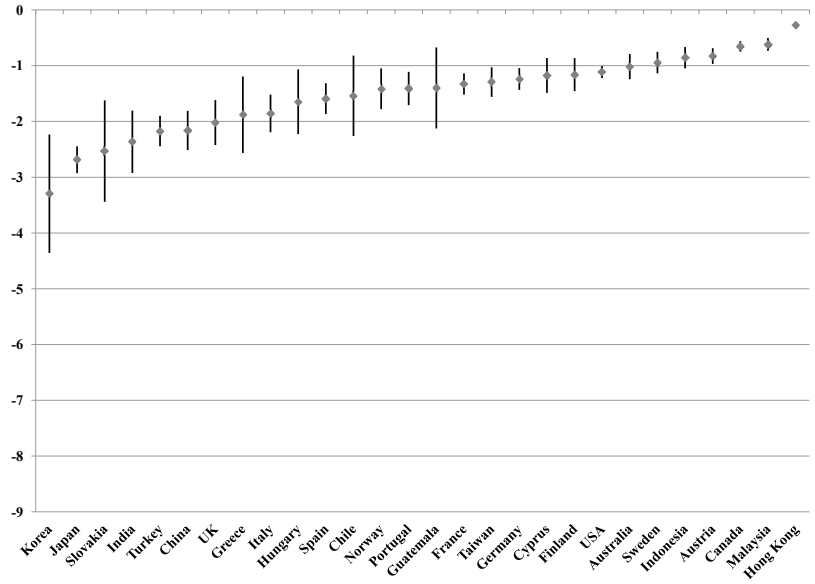


Figure 5: Distribution of constrained export elasticities ($\gamma = 1$, $\sigma_{kj} = \sigma_j$, $\forall k$)

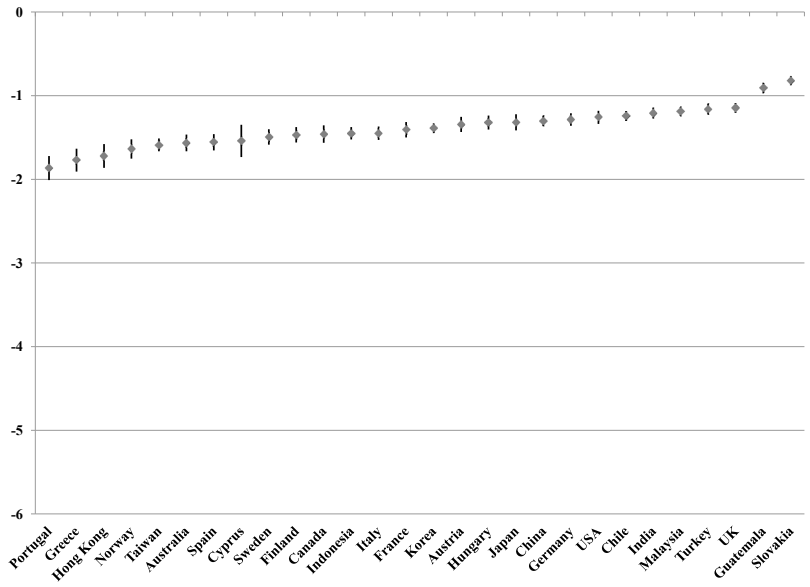


Figure 6: Unconstrained trade elasticities for EMU members

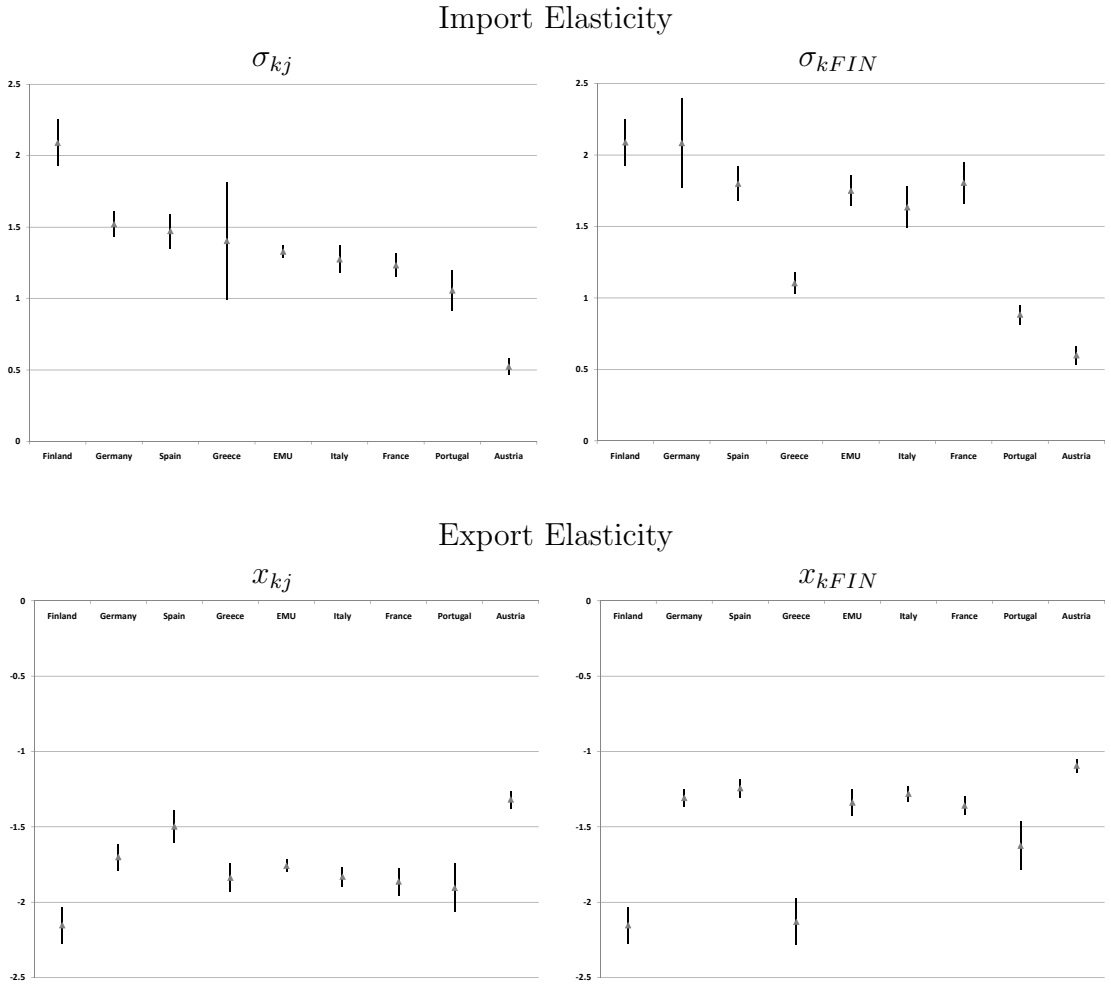


Figure 7: Constrained trade elasticities for EMU members