

Trade Elasticities*

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PRELIMINARY. DO NOT CIRCULATE

Abstract

Arkolakis et al (2011) show the welfare gains from trade can be summarized by import shares and the price elasticity of imports. The claim holds in a multi-sector model, which nests a one-sector version. Both versions must have identical welfare predictions, provided they are calibrated adequately. For 29 countries, we compute multi-sector welfare on the basis of sectoral import shares and estimated sector-level trade elasticities. From the one-sector version, we ask what value of the trade elasticity at country-level can replicate such welfare measures. This is the trade elasticity that must be used in the one-sector trade models considered in Arkolakis et al (2011). The elasticities are always significantly different from conventional macroeconomic estimates. Across 29 countries, they range from -2.1 to -6.6 , with lowest values in China and India. They take values around -4 in developed countries. We use the theory to decompose these international differences. We find the distribution of sectoral trade elasticities is key.

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1 Introduction

Trade elasticities measure the response of traded quantities to price shocks. They are often used to gauge a country’s “external performance”, but their fundamental economic interpretation is in fact model dependent. In an endowment economy with Armington preferences like Krugman (1980), the price elasticity of imports maps directly with the elasticity of substitution between international varieties. In a Ricardian model with perfect competition among heterogeneous firms like Eaton and Kortum (2002), or with monopolistic competition like Mélitz (2003), it relates with the distribution of firm productivity.

In spite of such differences, import price elasticities are sufficient to summarize the fundamental features of all three classes of models, as recently demonstrated by Arkolakis et al (2011) [ACRC henceforth]. In all three models, the welfare consequences of moving from observed, current trade to autarky can be evaluated with estimates of import price elasticities ε , and the observed shares of domestic expenditures λ_{jj} . The result generalizes to a multi-sector setting.

In this paper, we follow Imbs and Méjean (2011) and compute the aggregate welfare gains W_{MS} implied by sector-level trade elasticities ε^s , and sector-level expenditure shares λ_{jj}^s . We do so using the multi-sector model of ACRC, which we show nests the one-sector version. Since they come from the same model, the two versions must imply identical welfare, provided they are calibrated adequately. We ask what value of ε must be chosen for the one-sector model to imply the value of W_{MS} obtained from sectoral data. Such value represents the calibration of import price elasticities that belongs in the one-sector trade models in ACRC.

The exercise is conducted for 29 countries at various levels of development. The one-sector estimates of ε that mimics W_{MS} , denoted by ε_{OS} , range from -2.1 to -6.6 , with lowest values in China and India. The cross-country average equals -4.33 . Developed economies have estimates around -4 , -4.5 in the US, and slightly lower values for Canada, Japan, Korea or Singapore, close to -5 . The highest elasticities are found for specialized small open economies, Kuwait and Sri Lanka. The values of ε_{OS} are clearly significantly different from classic trade

elasticity estimates obtained from aggregate data. Consequently, the welfare measure calibrated to aggregate data, W_A , is systematically different from W_{MS} , even though in theory both arise from *two nested versions of the same model*. If the calibration is adequate, two different values cannot both be true simultaneously.

While the shares of domestic expenditures λ_{jj} and λ_{jj}^s are readily available from accounting data, trade elasticities must be estimated. Estimates obtained from aggregate data are not necessarily the same as those implied by sectoral information. In particular, the trade elasticity implied by aggregate data may in fact differ from an (adequately) weighted average of sector-specific trade elasticities. If it does, a heterogeneity bias prevails, and the price elasticity of imports obtained from aggregate data is effectively disconnected from sectoral information.¹ In that case, W_A differs from W_{MS} because of an econometric issue associated with an estimation performed on aggregates. Put differently, in the presence of heterogeneity bias, aggregate data should not be used when evaluating the gains from trade.

The result matters beyond the evaluation of the welfare gains from trade. To understand why, it is useful to re-visit the intuition behind ACRC's result. Changes in country j 's trade are triggered by shifts in production costs w_i in the foreign country i , and/or by changes in trade costs τ_{ij} . Neither affect domestic income. Therefore, trade-induced changes in real income, i.e. in welfare, can only stem from the price index. How does the domestic price index in country j respond to shifts in w_i and/or τ_{ij} ? That depends on the model. In the Armington model, the price response is proportional to the changes in the quantities traded, as substitution occurs between imports and domestic goods, scaled by the (inverse of the) import price elasticity. Thus, two parameters only, λ_{jj} and ε , summarize the welfare gains from trade in country j . The former because it reflects traded quantities in the current state of economy j , and the latter because it maps changes in quantities into changes in prices, which are what matter for

¹Levchenko, Lewis and Tesar (2010) estimate trade responses at sectoral level. Their elasticities are largest in sectors used as intermediate goods. Bussiere et al (2011) obtain trade elasticities for each component of aggregate demand. Their estimates of overall, aggregate trade elasticity are much larger than conventional ones.

welfare.²

In other words, W_A differs from W_{MS} because the bias in the estimated trade elasticity implies a bias in the response of prices and real income to terms of trade shocks. Within the confines of the (static) models in ACRC, aggregate data should not be used to calibrate the response of aggregate income to a terms of trade shock. Doing so simply assumes away the sectoral dimension of the data.

The claim generalizes to the models with endogenous supply considered by ACRC. In Ricardian trade models, the domestic price index aggregates the (minimum) prices of goods across production locations. But the price response to terms of trade shocks is not necessarily proportional to the changes in imports vs. domestic expenditures anylonger. Entry decisions (triggered by changes in production or trade costs) can differ across locations, with heterogeneous consequences on trade flows. ACRC show how Eaton and Kortum (2002) assume away this heterogeneity so that trade flows are still proportional to price shocks. The key assumption is the Frechet distribution of firm productivity, which in fact also pins down the ratio of traded quantities to relative prices. So the price elasticity of imports continues to determine the response of real income to terms of trade shocks, i.e. welfare.

With monopolistic competition and firm entry, the domestic price index incorporates not only production costs across locations, but also the number and mass of (imperfectly substitutable) varieties produced in each location. Either one can respond to terms of trade shocks, which in general breaks down the proportionality between prices and changes in imports vs. domestic expenditures. ACRC show that in Méltz (2003), the assumption that firm-specific productivity is distributed following a Pareto distribution restores the proportionality between prices and traded flows. The proportion is now given by the Pareto coefficient. Once again, the price elasticity of imports is key to determining the response of real income to terms of trade shocks, and welfare.

²To be precise, λ_{jj} is relevant when measuring the welfare losses associated with going from current trade to autarky.

We find that ε_{OS} , the one-sector trade elasticity implied by W_{MS} , is significantly different from an estimate obtained from aggregate data, ε_A . But what is it similar to? Using theory, we compare ε_{OS} with an adequately weighted average ε_W of sectoral estimates ε^s . The weights represent the size of each sector in overall expenditures, and their degree of openness to international trade. These weighted averages are calculated for 29 countries. They are virtually identical to ε_{OS} . We conclude that the calibration of one-sector models should build from averaged sectoral estimates, if the model is to mimick the response of real income to terms of trade shocks that is implied by a multi-sector model. A calibration using aggregate data does not correspond to the models described in ACRC.

We perform our exercise for 29 countries at various levels of development. There are large international differences in the estimates of ε_{OS} , which are typically absent in estimates from macroeconomic data. Using the definition of ε_W as a weighted average of ε^s , international differences can come either from the weights or the estimates of ε^s . We decompose international differences in ε_{OS} into three components: (i) the dispersion in estimates of ε^s across countries, (ii) differences in sectoral openness to trade, and (iii) differences in the sectoral allocation of expenditures.

Trade in most economies in western Europe is slightly less elastic (-4) than in the US (-4.5). The difference reflects the fact that sectoral trade is slightly less elastic than in the US, especially in large and open sectors. An exception is Germany, whose estimated elasticity is low relative to the rest of Europe, equal to -3.54 . At the sector level, German trade is in fact more elastic than in the US, especially in open and large sectors. But on the other hand, the large sectors in Germany are relatively more open, so that foreign shocks have sizeable effects on the domestic price index - i.e the trade elasticity is low.

The Chinese trade elasticity is -6.60 , the lowest in the sample. This happens because in China, large and open sectors are relatively *more* elastic than in the US. The same is true, albeit to a smaller extent, of Japan, Korea, and Canada. India, whose elasticity is -6.13 , has such low estimate because consumption falls on closed sectors, so that foreign shocks have

relatively little effect on the domestic price index - i.e. the trade elasticity is high.

The international differences uncovered in this paper point to the importance of sectoral specialization in explaining the elasticity of trade - measured by ε_{OS} - and ultimately the response of real income to terms-of-trade shocks, i.e. welfare. Of particular relevance in our sample is the dispersion in sector-specific elasticities, and its covariance with sectoral openness and expenditure share. In contrast, estimates of ε_A are virtually identical across countries. There are no international differences to explain in the first instance. And since they built from aggregate data, estimates of ε_A are essentially independent on the sectoral composition of output and trade.

The rest of the paper is structured as follows. Section 2 shows the one-sector version of ACRC is a special case of the multi-sector model, with identical welfare implications. Section 3 describes our estimation of sector-level elasticities, and data sources. Section 4 computes the one-sector trade elasticities ε_{OS} implied by the multi-sector model. They are compared with macroeconomic estimates, then with adequately weighted averages of sector level estimates. The section closes with a decomposition of international differences in trade elasticities. Section 5 verifies the robustness of the results. Section 6 concludes.

2 Welfare in One vs. Multi-Sector Trade Models

This section establishes that the one-sector version of ACRC is nested into the multi-sector model they develop in section 5.1. We follow Appendix A in ACRC, and consider first perfect, then monopolistic competition. By definition, changes in aggregate welfare W_{MS} associated with moving to autarky in the multi-sector world are given by

$$\partial \ln W_{MS} = \partial \ln Y - \partial \ln P \tag{1}$$

where Y is aggregate income and P is the price index.

2.1 Perfect competition

Labor markets clear. Assuming balanced trade, $\partial \ln Y = \partial \ln w = 0$, where the second equality comes from the choice of labor as the numeraire. The change in welfare corresponding to a change in trade costs, e.g. a move to autarky, is entirely driven by a change in prices. By definition, the price index is given by

$$P = \prod_s (P_s)^{\eta^s} \quad (2)$$

where η^s denotes the (Cobb-Douglas) expenditure share in sector s , and P_s is the sector-specific price index. The welfare loss associated with a move to autarky is given by

$$\partial \ln W_{MS} = - \sum_s \eta^s \partial \ln P_s \quad (3)$$

Under perfect competition, the sectoral price index is a weighted average of the marginal costs of production across all exporting countries, with weights given by import shares. Appendix A in ACRC shows that for small changes in the price index, the definition simplifies into

$$\partial \ln P_s = - \frac{1}{\varepsilon^s} \partial \ln \lambda_{jj}^s \quad (4)$$

where λ_{jj}^s is the share of domestic expenditures in sector s . Substituting into equation (3),

$$\partial \ln W_{MS} = \sum_s \frac{\eta^s}{\varepsilon^s} \partial \ln \lambda_{jj}^s \quad (5)$$

Now consider $\lambda_{jj} = \frac{\sum_s X_{jj}^s}{\sum_s Y^s}$, the aggregate share of domestic expenditures, where X_{jj}^s and Y^s denote domestic and total expenditures in sector s , respectively. By definition

$$\partial \ln \lambda_{jj} = \frac{\sum_s \eta^s \partial \lambda_{jj}^s}{\lambda_{jj}} = \sum_s \frac{\lambda_{jj}^s}{\lambda_{jj}} \eta^s \partial \ln \lambda_{jj}^s \quad (6)$$

A one-sector version of the model imposes unique parameters, i.e. $\varepsilon^s = \varepsilon$ and $\lambda_{jj}^s = \lambda_{jj}$. This of course says nothing about the calibration of either ε or λ_{jj} . The latter does have a definition, spelled out in equation (6), and is directly observable. The former, in contrast, must be estimated - and there is nothing in the model suggestive that this should be based on aggregate data. Let \bar{W}_{MS} denote the welfare implied by the multi-sector model, with sectoral heterogeneity assumed away. With such simplifications,

$$\partial \ln \bar{W}_{MS} = \frac{1}{\varepsilon} \sum_s \eta^s \partial \ln \lambda_{jj}^s = \frac{1}{\varepsilon} \sum_s \partial \ln \lambda_{jj} = \partial \ln W_{OS} \quad (7)$$

where W_{OS} is the welfare response in the one-sector model. Under perfect competition, the one sector model is a special case of the multi-sector version.

2.2 Monopolistic Competition

Under monopolistic competition with free entry, a zero profit condition holds that equates gross profit to fixed entry costs. Once again income $Y = L$, so that $\partial \ln Y = 0$. With restricted entry, ACRC impose that profits be proportional to income to obtain the same result. In either case, therefore, we continue to have equation (3).

Now however P_s is an ideal price index composed of all the imperfectly substitutable varieties produced across exporting countries. Appendix A in ACRC shows that for small changes in prices, we have

$$\partial \ln P_s = -\frac{1}{\varepsilon^s} \partial \ln \lambda_{jj}^s + \frac{1}{\varepsilon^s} \partial \ln N^s \quad (8)$$

where N^s denotes the measure of goods that can be produced in sector s . Substituting into welfare gives

$$\partial \ln W_{MS} = \sum_s \frac{\eta^s}{\varepsilon^s} \partial \ln \lambda_{jj}^s - \sum_s \frac{\eta^s}{\varepsilon^s} \partial \ln N^s \quad (9)$$

Consider again a one-sector version of the model, imposing the additional constraints that $\varepsilon^s = \varepsilon$ and $\lambda_{jj}^s = \lambda_{jj}$. The thus constrained measure of multi-sector welfare \bar{W}_{MS} changes

according to

$$\partial \ln \bar{W}_{MS} = \frac{1}{\varepsilon} \partial \ln \lambda_{jj} - \frac{1}{\varepsilon} \sum_s \eta^s \partial \ln N^s \quad (10)$$

If entry is restricted, then by assumption $N^s = \bar{N}^s$ and we have once again $\partial \ln \bar{W}_{MS} = \partial \ln W_{OS}$. Under free entry, gross profits are just sufficient to cover the total fixed costs of new varieties, $\Pi^s = N^s F^s$. The sector specific budget constraint rewrites

$$Y_s = wL^s + \Pi^s - N^s F^s = wL^s = L^s \quad (11)$$

Sectoral labor allocation varies with sectoral income. Assuming with ACRC that sectoral profits are proportional to sectoral income, we have

$$\partial \ln N^s = \partial \ln \Pi^s = \partial \ln L^s \quad (12)$$

So constrained welfare rewrites

$$\partial \ln \bar{W}_{MS} = \frac{1}{\varepsilon} \partial \ln \lambda_{jj} - \frac{1}{\varepsilon} \sum_s \eta^s \partial \ln L^s \quad (13)$$

But we have: $0 = \partial \ln Y = \partial \ln \sum_s Y_s = \frac{\sum_s \partial Y_s}{\sum_s Y_s} = \sum_s \eta^s \partial \ln Y_s = \sum_s \eta^s \partial \ln L_s$. So once again, $\partial \ln \bar{W}_{MS} = \partial \ln W_{OS}$.

Inasmuch as they stem from the same theory, the two versions must have the same welfare implications provided ε and λ_{jj} are calibrated adequately, i.e.

$$\partial \ln W_{MS} = \partial \ln W_{OS} \quad (14)$$

The natural analogy for one-sector variables is to obtain them from aggregate data. But a fundamental difference exists between ε and λ_{jj} . By definition, $\lambda_{jj} = \sum_s \eta_s \lambda_{jj}^s$: the aggregate domestic share is a weighted average of sectoral shares. The share of domestic expenditures is directly observable, from sectoral or indeed aggregate data. In contrast, ε must be estimated.

There is potentially a difference between a value for ε that is obtained from aggregate data, and a weighted average of ε^s , obtained from sectoral data. This difference can be understood as a heterogeneity bias.

The following experiment is therefore possible: (i) use sectoral data to estimate ε^s , and to calibrate λ_{jj}^s , (ii) compute the welfare loss $\partial \ln W_{MS}$, (iii) calibrate λ_{jj} from aggregate data, and (iv) use equation (14) to back the value of ε , denoted by ε_{OS} , that must be used to obtain identical welfare gains across the two versions.³ In particular, equation (14) implies

$$\varepsilon_{OS} = \frac{\partial \ln \lambda_{jj}}{\partial \ln W_{MS}} \quad (15)$$

The question is whether ε_{OS} can be estimated with aggregate data.

3 Estimation and Data

We first review the Armington demand system used to estimate the price elasticities of imports at sector level. The approach is similar to Feenstra (1994), adapted to sector-level data by Imbs and Méjean (2011). Then we review the data needed for the estimation and the welfare computations.

3.1 Estimation

Demand in sector s of country j is defined as an Armington aggregator of varieties indexed by i . Sectoral demand at time t is given by

$$C_{jt}^s = \left[\sum_{i \in I} (\beta_{ijt}^s C_{ijt}^s)^{\frac{\sigma^s - 1}{\sigma^s}} \right]^{\frac{\sigma^s}{\sigma^s - 1}} \quad (16)$$

³The actual formulae for W_{MS} depends on market structure. It is given by equation (5) under perfect competition, and equation (13) under monopolistic competition. For lack of reliable employment sectoral data across the countries in our sample, we focus on the perfect competition measure.

where $i \in I$ indexes both countries and varieties, including j , the variety produced domestically. The parameter β_{ijt}^s denotes preference shocks. Aggregate consumption in country j combines demand in each sector $s = 1, \dots, S$ according to

$$C_{jt} = \prod_{s \in S} \frac{(C_{jt}^s)^{\eta_j^s}}{(\eta_j^s)^{\eta_j^s}} \quad (17)$$

where η_j^s denotes the share of expenditures on sector s in country j . The representative agent chooses her consumption allocation on the basis of prices inclusive of transport costs τ_{ijt}^s . Utility maximization implies that nominal demand for variety i in each sector s is given by

$$P_{ijt}^s C_{ijt}^s = \left(\frac{P_{ijt}^s}{P_{jt}^s} \right)^{1-\sigma^s} (\beta_{ijt}^s)^{\sigma^s-1} P_{jt}^s C_{jt}^s \quad (18)$$

where P_{ijt}^s is the local currency price of variety i of good s , and $P_{jt}^s = \left[\sum_{i \in I} \left(\frac{P_{ijt}^s}{\beta_{ijt}^s} \right)^{1-\sigma^s} \right]^{\frac{1}{1-\sigma^s}}$.

The price elasticity of sectoral imports $\varepsilon^s = 1 - \sigma^s$ can be estimated from equation (18). It scales the (logarithm) response of sectoral imports to an exogenous shock in (logarithm) relative prices. As is well known, endogeneity issues notwithstanding, ε^s can be obtained from a gravity estimation derived from equation (18). To see this, assume shifts in relative prices are driven by shocks to trade costs. Equation (18) can be rewritten in logarithms:

$$\ln P_{ijt}^s C_{ijt}^s = A_{jt}^s + \varepsilon^s \tau_{ijt}^s + \nu_{ijt}^s \quad (19)$$

where $A_{jt}^s = \ln P_{jt}^s C_{jt}^s$ and $\nu_{ijt}^s = \ln (\beta_{ijt}^s)^{\sigma^s-1}$. Equation (19) represents a classical gravity regression, that falls in the category described in section 6 of ACRC. Estimates arising from equation (19) can therefore be interpreted within the three classes of models discussed there.

Of course, equation (19) is not necessarily well identified. For one thing, trade costs are not the only cause for price shocks. For another, the residuals ν_{ijt}^s are not necessarily orthogonal to τ_{ijt}^s . To achieve identification, Feenstra (1994) or Imbs and Méjean (2011) impose a simple

supply structure, given by

$$P_{ijt}^s = \tau_{ijt}^s \exp(v_{ijt}^s) (C_{ijt}^s)^{\omega^s} \quad (20)$$

where v_{ijt}^s denotes a technological shock, and ω^s is the inverse elasticity of supply in sector s . Shocks to relative prices can now arise either because of trade costs, or technological developments. Feenstra (1994) famously showed that combining equations (18) and (20) can achieve identification under further assumptions on the fundamental shocks v_{ijt}^s and β_{ijt}^s .

We implement his procedure, adapted to sector level data, as described in Imbs and Méjean (2011). For each importing country j and each sector s , identification is obtained in the cross-section of exporters i . The data needed are therefore multi-lateral information on traded quantities and prices. To alleviate measurement error, imported values are replaced with market shares $m_{ijt}^s = \frac{P_{ijt}^s C_{ijt}^s}{P_{jt}^s C_{jt}^s} \mu_{jt}^s$ where $\mu_{jt}^s = 1 + \frac{P_{jt}^s C_{jt}^s}{\sum_{i \neq j} P_{ijt}^s C_{ijt}^s}$ corrects for domestic consumption in sector s . Prices are measured as unit values. The estimation is performed for each sector, and for each country. Common correlated effects are accounted for across the sectors of a single country following Imbs and Méjean (2011).

Many alternatives approaches exist to estimate trade elasticities. Most rest on an identification of the exogenous component of trade costs in equation (19), typically constrained to be time-invariant. Eaton and Kortum (2002) invoke an arbitrage argument to reason that bilateral trade costs must be approximated by the maximum difference in price levels observed across goods for a given country pair. Simonovska and Waugh (2011) show that this approach creates biased estimates of τ_{ij} , because of the limited number of goods for which such information is available. They propose instead to implement a simulated method of moments on Eaton and Kortum (2002). Bernard, Eaton, Jensen and Kortum (2003), or Eaton, Kortum and Kramarz (forthcoming) fit a multi-country Ricardian model on detailed firm-level data, which delivers - among others - trade elasticity estimates.

It should be clear this paper does not rely on the very methodology used to estimate trade elasticities. Any alternative econometric approach is acceptable, provided it falls under the

umbrella of the gravity equation (19). That will guarantee the argument developed in ACRC applies to the resulting estimates. The necessary data must also be relatively parsimonious, so that elasticities can be estimated at sectoral level, for a range of countries. The only instance of such estimates that we are aware of is due to Broda, Greenfield and Weinstein (2006), who also make use of the Feenstra (1994) approach. Of course, an open question is how different the results would be with alternative sectoral trade elasticity estimates. Although a precise assessment of this concern is impossible for lack of alternative international, sectoral estimates, we provide a discussion in Section 4.

3.2 Data

Sectoral information is needed on the cross-section of imported quantities and unit values for a cross-section of countries. In the main text, elasticities are estimated on the basis of the United Nations ComTrade database, using export declarations for maximum coverage. Alternatively, a set of estimates was obtained using the BACI dataset, compiled by CEPII, which is meant to harmonize ComTrade and scan for measurement errors or duplicates. The alternative results are reported in Section 5.

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. The universe of products is partitioned into sectors according to the 3-digit ISIC (revision 2) level, which makes for a maximum of 26 sectors. The data are yearly between 1995 and 2004. Before 1995, the number of reporting countries is unstable. And the unit values reported in ComTrade experience a structural break in 2004. Values of ε^s are estimated on the complete time series.

Identification requires that the cross-section of countries exporting to j be wide enough for all sectors, and remain so over time. We retain goods for which a minimum of 20 exporting countries are available throughout the period. Both unit values and market shares are notoriously plagued by measurement error. We compute the median growth rate at the sector level

for each variable, across all countries and years. The bilateral trade flows are excluded for all sectors with growth rates in excess of five times that median value, either in unit values or in market share. The resulting 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. The Table also reports the total number of exporters into each country j . It is given by the number of sectors in country j , multiplied by the number of exporting countries for each sector. The average number of exporters ranges from 35 in Sri Lanka to more than 90 in the United States. For each sector, the data imply an average number of exporting countries equal to 53.

The main data constraint is not imposed by trade data. The binding limitation concerns the weights that enter W_{MS} , and are used to decompose the source of international differences in trade elasticities. Both η^s and λ_{jj}^s require information on domestic consumption at sectoral level, that must be compatible with the trade data in ComTrade. The constraint raises issues of concordance since information is needed on both production and trade at the sectoral level. This is what reduces the coverage from 33 to 29 countries. We use a dataset built by Di Giovanni and Levchenko (2009) who merge information on production at the 3-digit ISIC (revision 2) level 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 posit

$$\eta_s \equiv \frac{Y_j^s - X_j^s + M_j^s}{\sum_k (Y_j^s - X_j^s + M_j^s)} \quad (21)$$

where X_j^s (M_j^s) denotes country j 's exports (imports) in sector s . And

$$\lambda_{jj}^s \equiv \frac{Y_j^s - X_j^s}{Y_j^s - X_j^s + M_j^s} \quad (22)$$

To focus on meaningful computations, a minimum of 10 sectors is imposed for all countries. The constraint tends to exclude small or developing economies, such as Panama or Poland.

The UNIDO data are in USD, and available at a yearly frequency. The values of η^s and λ_{jj}^s are computed over five-year averages in order to limit the consequences of cyclical fluctuations in trade. Two sets of estimations are reported. In the main text, we use average weights between 1991 and 1995. For robustness, we also consider averages between 1996 and 2000.

The UNIDO dataset is focused on manufacturing goods only, which can bring into question the validity of trade elasticity estimates. But the vast majority of traded goods are manufactures, so that the truncation remains minimal. We have experimented with the values for η^s and λ_{jj}^s implied by the OECD Structural Analysis database (STAN), which provides information on all sectors of the economy. For countries covered by both datasets, i.e. OECD members, the end elasticities were in fact virtually identical. At least for OECD members, this suggests the sampling caused by the UNIDO dataset is kept to a minimum. The last column in Table 1 reports the fraction of total trade afforded by UNIDO data. The coverage is below 40 percent for small open economies such as Hong Kong, Cyprus, or Chile, but above 70 percent for large developed economies such as the US, France or Spain. Coverage is clearly limited for small open, developing economies. But, contrary to OECD data, it leaves the door open to some analysis for the developing world, not least China where coverage is above 50%.

4 Trade Elasticities in the One-Sector Model

We report the estimates of ε^s implied by sectoral data for the 29 countries with the required data, and discuss the corresponding values of ε_{OS} . One-sector elasticities are first compared with conventional macroeconomic estimates, and then with the weighted averages of sectoral elasticities ε^s implied by theory. International differences in these weighted averages are then ascribed to different weights across countries, vs. different estimates of ε^s .

4.1 Sector-Level Estimates, Multi-Sector Welfare Loss and ε_{OS}

Table 2 presents some summary statistics of the estimates of ε^s implied by ComTrade data between 1995 and 2004. There is considerable heterogeneity in mean sectoral elasticities across countries. Developed countries display average values around -4 , Germany at -4.06 , France at -4.25 , or Spain at -4.30 . Specialized oil exporters also display relatively low average sectoral elasticities. Both Kuwait and Venezuela have estimates below 4 in absolute value. In contrast, developing exporting economies present estimates at least twice larger. China has the largest mean sectoral elasticity, equal to -9.22 , closely followed by India, Korea and Sri Lanka.

Sectoral heterogeneity is sizeable within countries as well. The distribution of estimates tends to be most disparate and skewed in developing economies. For instance, estimates of ε^s range between -2.56 and -29.01 in China, with a median of -5.85 , substantially below the mean of -9.22 . In India, estimates range from -1.41 to -29.01 . Ranges tend to be narrower for European developed countries, such as France, Germany, Italy or Spain. The distributions there tend to be more symmetric, with mean and median elasticities closer together.

Country and sector effects each explain approximately 10 percent of the cross-country dispersion in estimates of ε^s . Close to 80 percent of the variance in ε^s must therefore correspond to international differences in the trade elasticity for each sector s . The result is apparent from Table 2, where some sectoral estimates are drastically different from one country to the next. For instance, the price elasticity of Textiles imports is -6.00 in Spain, but -29.01 in Sri Lanka. Imports of Rubber Products are inelastic in Kuwait ($\varepsilon^s = -1.56$), but elastic in the Philippines ($\varepsilon^s = -24.06$). Such disparities may correspond to differences in the very nature of the goods imported. For instance, textile imports into Spain are likely to be of higher quality than the (probably intermediate) goods imported by Sri Lanka.

Such heterogeneity in sectoral estimates is relevant to this paper inasmuch as it affects the welfare gains from trade implied by the multi-sector version of ACRC. We now compute the welfare loss $\partial \ln W_{MS}$ associated with a move from current trade to autarky, across all the

countries with relevant data. The welfare loss is given by equation (5):

$$\partial \ln W_{MS} = \sum_s \frac{\eta^s}{\varepsilon^s} \partial \ln \lambda_{jj}^s$$

whose estimation requires calibrated values for η^s and λ_{jj}^s . Welfare losses get close to zero for low values of $\partial \ln \lambda_{jj}^s$, i.e. in closed economies where λ_{jj}^s is close to 1. They decrease in trade elasticities, because low ε^s means large price responses to shifts in the quantities traded, i.e. large welfare loss. And they increase in η^s , i.e. in the share of sector s in expenditures.

Armed with the welfare gains implied by the multi-sector version, and observed values for λ_{jj} , the one-sector value of ε_{OS} is given by

$$\varepsilon_{OS} = \frac{\partial \ln \lambda_{jj}}{\partial \ln W_{MS}}$$

This is the trade elasticity that equates welfare in the one-sector model of ACRC to what is implied by its multi-sector version. As ACRC, we consider the welfare loss associated with a move to autarky, i.e. $\partial \ln \lambda_{jj} = \ln \lambda_{jj}$.

The first three columns in Table 3 report estimates of $\partial \ln W_{MS}$, the calibrated values of λ_{jj} , and the corresponding estimates of ε_{OS} for the 29 countries with data. Standard errors are obtained using the Delta method, as detailed in the Appendix. The welfare losses from autarky are highest in small open economies, like Hong Kong (40.3% of real income), Singapore (22.5%) or Kuwait (45.7%). They are lowest in large, closed economies, such as Japan (1.34%), India (1.97%), China (3.08%) or the US (3.97%). Finally the losses are estimated around 10% of real income for developed, West European economies.

The ranking correlates with measures of overall openness, as reflected in the aggregate share of domestic expenditures $\lambda_{jj} \equiv \frac{Y_j - X_j}{Y_j - X_j + M_j}$. It takes lowest values in small open economies, like Hong Kong or Singapore, and highest in large or closed countries, such as Japan, the US or India. But openness is not the sole determinant of welfare. $\partial \ln W_{MS}$ also decreases with trade

elasticities, and depends on their distribution across sectors. The case of Kuwait is illustrative, where the welfare loss is 45.7%, the highest in our sample. Kuwait is very open to trade, with $\lambda_{jj} = 0.38$. But it is not nearly as open as Hong Kong, where $\partial \ln W_{MS}$ is nonetheless also close to 40%. The large welfare loss in Kuwait comes from relatively low average sectoral elasticities, as illustrated in Table 1. It also come from the cross-sector correlation between λ_{jj}^s and ε_s . For given average openness and average trade elasticity, the welfare loss $\partial \ln W_{MS}$ takes higher (absolute) value if open sectors tend to display low elasticities. This tends to happen in Kuwait, an open economy on average, whose imports are specialized in sectors with low trade elasticities. The specialization of trade matters for welfare.

The third column in Table 3 reports the values of ε_{OS} implied by observed λ_{jj} and estimated W_{MS} . There are once again considerable cross-country differences. Estimates of the one-sector trade elasticity range from around -2 in Indonesia, Kuwait and Sri Lanka, down to -6 in China or India. Intermediate values between -4 and -5 are found for developed economies, with -4.48 for the US or -3.95 for the UK. No obvious correlate of ε_{OS} is apparent from Table 3, as developing economies can be found at either extreme of the range of estimates.

Are such estimates different from the literature? The average of ε_{OS} across countries equals -4.33 . This is close to half of the value found in Eaton and Kortum (2002), but not any different from the corrected estimates presented in Simonovska and Waugh (2011). It is within the range of estimates in Bernard, Jensen, Eaton and Kortum (2003) or in Eaton, Kortum and Kramarz (2011). These papers do not seek to estimate trade elasticities per se, and so no sectoral values are reported. Nor are trade elasticities obtained for more than one country in each paper.⁴ It is however reassuring that the average of ε_{OS} is no different from results in the literature that pertain to developed countries. In our sample, trade elasticities in western Europe and the US are close to -4 , as they are in the literature. This suggests the estimates of ε_{OS} in Table 3 do not crucially depend on the estimation approach adopted here. Neither do the comparisons and the decomposition of international differences presented in the next

⁴Broda and Weinstein (2006) do present international, sectoral estimates. They are close to those in this paper - unsurprisingly given the similarity in the estimation methods.

sections.

4.2 Comparisons

Table 3 does suggest an important result: the estimates of trade elasticities are unusual for one-sector models. We now compare our estimates of ε_{OS} with alternative candidates. Trade elasticity estimates that arise from aggregate data are first considered. Aggregate data are the most natural source when it comes to estimating parameters that enter one-sector models. It is self-evident from Table 3 that our estimates of ε_{OS} are significantly different from the conventional values for import price elasticities obtained in macroeconomics.⁵ For instance, Figure 1 reproduces the estimates obtained in Houtakker and Magee (1969) for 15 developed economies. No point estimates are below -2 , some are positive, and 10 out of 15 are not significantly different from zero. In fact, virtually no two estimates are significantly different from each other. For instance, the US price elasticity of imports is -0.5 , Japan's is -0.78 , and Canada's is -1.5 .

A similar exercise was conducted using ComTrade data, which we aggregated to country level in order to estimate a gravity equation. Using the notation from section 3, we estimate

$$\Delta \ln P_{ijt} C_{ijt} = A_{ij} + (\varepsilon_j^A + 1) \Delta \ln P_{ijt} + \tilde{\nu}_{ijt} \quad (23)$$

where $\Delta X_t = X_t - X_{t-1}$, $P_{ijt} C_{ijt} = \sum_s P_{ijt}^s C_{ijt}^s$ and $\Delta \ln P_{ijt} = \frac{1}{2} \sum_s \left(\frac{P_{ijt}^s C_{ijt}^s}{P_{ijt} C_{ijt}} + \frac{P_{ijt-1}^s C_{ijt-1}^s}{P_{ijt-1} C_{ijt-1}} \right) \Delta \ln P_{ijt}^s$ is a Tornqvist price index. Identification is obtained through time variation. Column 4 in Table 3 reports the estimates of ε_j^A . In stark contrast with ε_{OS} , the estimates of ε_j^A are all between 0 and -1 . Of course, equation (23) is acutely problematic econometrically, as changes in prices are endogenous. It is however extremely unlikely a correction for such endogeneity (which remains largely elusive in aggregate data) would imply estimates of ε_j^A in line with the implication

⁵See for instance the estimates reported in Francis et al (1976). The book contains summaries of available, pertinent, aggregate elasticities from the literature. These were used to calibrate the Deardorff-Stern Michigan CGE Model of World Production and Trade.

of ACRC, i.e. ε_{OS} . At the very least, no existing estimates using aggregate data come even close.

Estimates of ε_{OS} can therefore not be reproduced from aggregate data. But can they be obtained from sectoral data? We make use of the theory developed by ACRC to derive an analytical expression of ε_{OS} , as a weighted average of sectoral estimates.⁶ Consider welfare in the multi-sector model, given by equation (5). After some rearranging,

$$\partial \ln W_{MS} = \partial \ln \lambda_{jj} \sum_s \frac{\eta^s \Delta \lambda^s}{\varepsilon^s}$$

where $\Delta \lambda_s = \frac{\partial \ln \lambda_{jj}^s}{\partial \ln \lambda_{jj}} > 0$ denotes the openness of sector s relative to the country average. Note that $\Delta \lambda_s$ *increases* in relative sectoral openness as both λ_{jj}^s and λ_{jj} are smaller than 1. Welfare losses increase still with the overall openness of the economy, i.e. for low values of λ_{jj} . They also increase in relative sectoral openness $\Delta \lambda^s$, holding country average constant. By analogy with the expression for W_{OS} , we have

$$\varepsilon_{OS} = \left(\sum_s \frac{\eta^s \Delta \lambda^s}{\varepsilon^s} \right)^{-1} \equiv \varepsilon_W \quad (24)$$

The one-sector elasticity is a weighted average of sector level estimates ε_s , with weights given by $\eta^s \Delta \lambda^s$.

The weights reflect relative openness to trade, $\Delta \lambda_s$, and each sector's importance in overall consumption, η^s . The specialization of production and consumption matters therefore in two ways for one-sector elasticities. First, sectors that compose a large fraction of total expenditures receive a small weight. For a given shock and a given sectoral elasticity, a large value of η^s implies a large response of the overall price index, i.e. low aggregate trade elasticity. For the same reason, relatively open sectors enter with a small weight. For a given shock to traded quantities, a large value of $\Delta \lambda^s$ means a large response of the sectoral price index. The response of the aggregate price index is accordingly large, which means low aggregate trade elasticity.

⁶For simplicity and tractability, we focus on the perfect competition case.

The last column in Table 3 reports estimates of ε_W computed using equation (24). Perhaps unsurprisingly, the values obtained for ε_W are identical to ε_{OS} , albeit sometimes estimated with less precision.

4.3 International Dispersion

International differences in trade elasticities are absent in estimates obtained from aggregate data. Thus in macroeconomics, trade elasticities are customarily assumed to be identical across countries. By definition, they are invariant to differences in the specialization of trade across countries. This is an undesirable property in light of anecdotal and journalistic arguments that the specialization of production or trade has direct implications on countries' external performance.

The trade elasticity this paper suggests ought to be used in one-sector models does not share this property. Cross-country estimates of ε_{OS} displays considerable heterogeneity. Theory can be used to identify the sources of such heterogeneity, and ascribe them to the specialization of trade. Using the definition of ε_W , it is easy to show how the one-sector elasticity in country j will respond to different determinants. In particular, a Taylor expansion of equation (24) around a reference country implies

$$\Gamma (\varepsilon_W - \varepsilon_W^*) = \sum_s \frac{\Delta \lambda^s}{-\varepsilon^s} (\eta^s - \eta^{s*}) + \sum_s \frac{\eta^s}{-\varepsilon^s} (\Delta \lambda^s - \Delta \lambda^{s*}) + \sum_s \frac{\eta^s \Delta \lambda^s}{(\varepsilon^s)^2} (\varepsilon^s - \varepsilon^{s*}) \quad (25)$$

where $\Gamma = \left(\sum_s \frac{\eta^s \Delta \lambda^s}{\varepsilon^s} \right)^2$, and starred variables denote the reference country's parameters. Equation (25) implies that the international dispersion in trade elasticities depends on the sectoral composition of expenditures, openness, and trade elasticities, rather than on their average levels.

International differences are determined by three covariance terms. The first reflects international differences in the sectoral composition of expenditures. The second one reflects differences in sectoral openness, and the third quantifies the importance of different sectoral

trade elasticities. In absolute value, the elasticity in country j is relatively high if (i) consumers spend less (relative to the reference country) in open and inelastic sectors, (ii) large and inelastic sectors are also closed (relative to the reference), and (iii) sectors that are elastic (relative to the reference) also tend to be large and open.

Performing the decomposition described in equation (25) is straightforward, given the data requirements involved in computing W_{MS} . For reference, Figure 2 reproduces the cross-country estimates of ε_W reported in Table 3. The elasticities are represented relative to the US, which is the reference country used in implementing the decomposition in equation (25). Figure 3 reports the corresponding three elements for 29 countries, with $H = \sum_s \frac{\Delta\lambda^s}{-\varepsilon^s} (\eta^s - \eta^{s*})$, $\Lambda = \sum_s \frac{\eta^s}{-\varepsilon^s} (\Delta\lambda^s - \Delta\lambda^{s*})$, and $E = \sum_s \frac{\eta^s \Delta\lambda^s}{(\varepsilon^s)^2} (\varepsilon^s - \varepsilon^{s*})$. Since all three can take either sign, international differences in each component of equation (25) tend to be magnified.

It is interesting to note that high average estimates of ε^s , which tend to happen in the developing world as shown in Table 2, do not translate into large negative values for E . For instance, Sri Lanka and Venezuela have large positive values of E , whereas they are negative in China. In Indonesia, Egypt or the Philippines, E is actually close to zero. As is obvious from equation (25), there is no correlation between sectoral averages of ε^s and the value of E . International differences in ε_W arise because the sectoral *distributions* of ε^s , η^s and $\Delta\lambda^s$ change from one country to the next. Figure 3 suggests these international differences are smallest as regards $\Delta\lambda^s$, as Λ tends to be the least important element of $\varepsilon_W - \varepsilon_W^*$ across the 29 countries. In other words, the distribution of sectoral openness displays relatively small cross-country dispersion, as compared with consumption expenditures η^s or sectoral elasticities ε^s .

Several results are of particular interest. Both China and India display substantially larger elasticities of trade than the US, with values around -6.25 . Figure 3 reveals this happens for different reasons. Chinese trade tends to be elastic because $E < 0$, whereas $H < 0$ in India. Chinese trade would be as elastic as the US if it had similar sectoral elasticities. But they are in fact larger in absolute value: Chinese imports are elastic and specialized in large and open sectors. The same is true of Korea, Japan and Canada, albeit with smaller magnitudes. In

contrast, India's trade is elastic because India spends less than the US on open sectors. Indian trade would actually become less elastic than the US if it shared the same η^s .

Most European elasticities are slightly closer to zero than the US. This typically happens because E takes small, positive values. In the United Kingdom for instance, $E = 0.50$ explains most of the difference with the US. The same is true in France, Italy, Spain or Sweden, where positive values of E are the main reason for the (small) discrepancy with the US. These are countries where relatively inelastic sectors compose a large fraction of expenditures, in open sectors.

An major exception in Europe is Germany, where ε_W is estimated to be -3.54 . That is still close to the US or the rest of Europe. But Figure 3 illustrates the apparent similarity masks in fact large discrepancies. First, $H = 1.73$: German consumers spend more in open, inelastic sectors than their US counterpart. If this were the sole difference with the US, the German trade elasticity would jump up to -2.75 . But in addition, $E < 0$ in Germany, around -1.5 . This means open, large sectors in German expenditures tend to be more elastic than in the US. In the end, the German trade elasticity is close to the US because the sectoral allocation of consumption and the distribution of sectoral trade elasticities work in opposite directions.

The two countries with lowest estimates of ε_W are Kuwait and Sri Lanka. Figure 3 suggests both countries have large, positive values for E and Λ . The two countries present values of $\Delta\lambda^s$ that are systematically larger than in the US, across all sectors. The actual values of λ_{jj}^s cover broad ranges in both countries, with only a few closed sectors, so that the distributions in both countries are heavily skewed. This translates in high - a few very high - values of $\Delta\lambda^s$ as most sectors are more open than the average. Thus Λ takes positive values. The same is at play in a third country, Egypt. In addition, in both countries, the sectors with highest values for $\Delta\lambda^s$ happen to be relatively less elastic than the US, which implies $E > 0$. What makes both countries special is the concordance of both coincidences: a skewed distribution of $\Delta\lambda^s$, with high values precisely in relatively inelastic sectors.

The decomposition of ε_W is relevant to understanding the international dispersion in trade

elasticities. It is of course also important for welfare. The welfare gains from trade decrease in the trade elasticity, so that large estimates of ε_{OS} mean lower welfare than what is implied by aggregate data. For instance, Figure 3 suggests the welfare gains from trade in China would be substantially higher if the distribution of sectoral elasticities were closer to the US. They would be higher in India if India spent more on open sectors. They would be lower in Kuwait if openness were more evenly distributed across sectors. To our knowledge, there is no alternative methodology that implies such a close mapping between the sectoral specialization of consumption and production, the elasticity of trade, and ultimately the welfare gains from trade.

5 Robustness

5.1 Trade data from BACI

TBC

5.2 Calibration of η^s and λ_{jj}^s based on 1996-2000 data

TBC

5.3 Alternative Estimation of Sectoral Elasticities

TBC

6 Conclusion

The welfare gains from trade are computed for 29 countries, on the basis of the multi-sector model developed by Arkolakis et al (2011) [ACRC]. Welfare is given by the static response of real income to a terms-of-trade shock, and it is summarized by import shares and trade elasticities. The multi-sector welfare measure is computed from observed sectoral import shares

and estimated sectoral trade elasticities, obtained using the methodology developed by Feenstra (1994) and adapted in Imbs and Méjean (2011). The one-sector version is a special case, and should have identical welfare predictions to the multi-sector model, provided it is calibrated adequately. On the basis of observed import shares at country level, we estimate the trade elasticity implied by the one-sector version of ACRC, constrained to predict multi-sector welfare. This is the trade elasticity that belongs in the one-sector models considered by ACRC.

The estimates are significantly different from conventional, macroeconomic trade elasticities. They are larger in absolute value, and heterogeneous across countries, with values ranging between -2.1 and -6.6 . China and India have low estimates, below -6 . Western Europe and the US display estimates around -4 , whereas Canada, Japan, Korea, and Singapore are closer to -5 . The lowest values are found for specialized economies, like Kuwait or Sri Lanka. Using the theory, a decomposition of this international dispersion is introduced. Trade elasticities can differ because of the specialization of consumption, of production, or because of international differences in sector-level trade elasticities.

European elasticities are close to zero because, relative to the US, inelastic sectors are open, and compose a large fraction of expenditures. Germany is a European exception: spending falls on relatively elastic sectors. But spending also falls relatively more on open sectors, which means low elasticity since the price level is then responsive to foreign shocks. The latter effect dominates, so that German trade is slightly less elastic than the European average. China, and to a smaller extent Canada, Japan, and Korea have high elasticities because expenditures fall on sectors that are elastic relative to the US. India is an exception amongst exporting countries. Its trade is elastic because relatively less spending falls on open sectors, so that its price level is less responsive to foreign shocks. Inasmuch as welfare decreases in the trade elasticity, these decomposition carry through to welfare.

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Appendix: Variances

The variance of $\partial \ln W_{MS}$

Consider a Taylor expansion of $\partial \ln W_{MS} = \sum_s \frac{\eta^s}{\varepsilon^s} \ln \lambda_{jj}^s$ around its estimated value $\partial \ln \hat{W}_{MS}$.

We have:

$$\begin{aligned} \partial \ln W^{MS} &= \partial \ln \hat{W}_{MS} + \sum_s \left. \frac{\partial \ln W^{MS}}{\partial \varepsilon^s} \right|_{\varepsilon^s = \hat{\varepsilon}^s} (\varepsilon^s - \hat{\varepsilon}^s) \\ &= \partial \ln(\hat{W}^{MS}) + \sum_s \frac{-\eta^s}{(\hat{\varepsilon}^s)^2} (\varepsilon^s - \hat{\varepsilon}^s) \ln \lambda_{jj}^s \end{aligned}$$

The variance is therefore given by

$$Var(\partial \ln W^{MS}) = \sum_s \left(\frac{\eta^s}{(\hat{\varepsilon}^s)^2} \ln \lambda_{jj}^s \right)^2 Var(\hat{\varepsilon}^s)$$

The variance of ε_{OS}

A Taylor expansion of $\varepsilon_{OS} = \frac{\ln \lambda_{jj}}{\partial \ln W_{MS}}$ around its estimated value $\hat{\varepsilon}_{OS}$ implies

$$\varepsilon_{OS} = \hat{\varepsilon}_{OS} - \frac{\ln \lambda_{jj}}{(\partial \ln W_{MS})^2} (\partial \ln W_{MS} - \partial \ln \hat{W}_{MS})$$

The variance is therefore given by

$$Var(\varepsilon_{OS}) = \left(\frac{\ln \lambda_{jj}}{(\partial \ln W_{MS})^2} \right)^2 Var(\partial \ln W^{MS})$$

The variance of ε_W

A Taylor expansion of $\varepsilon_W = \left(\sum_s \frac{\eta_s \Delta \lambda_s}{\varepsilon_s} \right)^{-1}$ around its estimated value $\hat{\varepsilon}_W$ implies

$$\varepsilon_W = \hat{\varepsilon}_W + \frac{\sum_s \frac{\eta_s \Delta \lambda_s}{(\hat{\varepsilon}_s)^2}}{\left(\sum_s \frac{\eta_s \Delta \lambda_s}{\hat{\varepsilon}_s} \right)^2} (\varepsilon^s - \hat{\varepsilon}^s)$$

So that the variance is given by

$$Var(\varepsilon_W) = \left(\sum_s \frac{\eta_s \Delta \lambda_s}{\hat{\varepsilon}_s} \right)^{-4} \left(\sum_s \frac{\eta_s \Delta \lambda_s}{(\hat{\varepsilon}_s)^2} \right)^2 Var(\hat{\varepsilon}^s)$$

Table 1: Summary Statistics

	# sect	# sect \times 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: Cross country Estimates of ε^s
Summary Statistics

Australia			
	Nb Sect	16	
	Mean	-8.20	
	Median	-6.30	
	Minimum	-3.51	Other non-metallic mineral products
	Maximum	-23.26	Pottery, China and Earthenware
Austria			
	Nb Sect	24	
	Mean	-6.02	
	Median	-5.32	
	Minimum	-2.67	Other non-metallic mineral products
	Maximum	-22.96	Pottery, China and Earthenware
Canada			
	Nb Sect	24	
	Mean	-5.68	
	Median	-5.34	
	Minimum	-2.73	Footwear, except vulcanized or moulded rubber or plastic footwear
	Maximum	-10.81	Machinery, except electrical
China			
	Nb Sect	18	
	Mean	-9.22	
	Median	-5.85	
	Minimum	-2.56	Printing, publishing and allied industries
	Maximum	-29.01	Transport equipment
Egypt			
	Nb Sect	19	
	Mean	-5.70	
	Median	-4.76	
	Minimum	-2.61	Other manufacturing industries
	Maximum	-15.51	Transport equipment
Finland			
	Nb Sect	25	
	Mean	-6.56	
	Median	-4.95	
	Minimum	-2.61	Glass and glass products
	Maximum	-21.56	Wearing apparel, except footwear
France			
	Nb Sect	25	
	Mean	-4.25	
	Median	-3.89	
	Minimum	-2.48	Wood and cork products, except furniture
	Maximum	-6.63	Wearing apparel, except footwear

Germany			
	Nb Sect	20	
	Mean	-4.06	
	Median	-3.93	
	Minimum	-2.41	Glass and glass products
	Maximum	-7.35	Non-ferrous metal basic industries
Greece			
	Nb Sect	26	
	Mean	-6.06	
	Median	-4.41	
	Minimum	-2.61	Glass and glass products
	Maximum	-29.01	Leather and products of leather, leather substitutes and fur, except footwear and wearing apparel
Hong Kong			
	Nb Sect	11	
	Mean	-5.14	
	Median	-5.11	
	Minimum	-2.26	Printing, publishing and allied industries
	Maximum	-9.20	Food manufacturing
Hungary			
	Nb Sect	19	
	Mean	-7.18	
	Median	-5.37	
	Minimum	-2.85	Glass and glass products
	Maximum	-19.86	Furniture and fixtures, except primarily of metal
India			
	Nb Sect	17	
	Mean	-8.30	
	Median	-5.52	
	Minimum	-1.41	Printing, publishing and allied industries
	Maximum	-29.01	Plastic products not elsewhere classified
Indonesia			
	Nb Sect	14	
	Mean	-5.03	
	Median	-4.47	
	Minimum	-3.03	Glass and glass products
	Maximum	-11.36	Textiles
Israel			
	Nb Sect	20	
	Mean	-7.54	
	Median	-5.13	
	Minimum	-3.17	Leather and products of leather, leather substitutes and fur, except footwear and wearing apparel
	Maximum	-20.66	Non-ferrous metal basic industries

Italy			
	Nb Sect	25	
	Mean	-4.61	
	Median	-4.35	
	Minimum	-2.79	Other non-metallic mineral products
	Maximum	-10.61	Furniture and fixtures, except primarily of metal
Japan			
	Nb Sect	26	
	Mean	-5.30	
	Median	-5.05	
	Minimum	-2.18	Petroleum refineries
	Maximum	-11.09	Transport equipment
Korea			
	Nb Sect	25	
	Mean	-8.63	
	Median	-6.78	
	Minimum	-2.26	Printing, publishing and allied industries
	Maximum	-29.01	Other industries
Kuwait			
	Nb Sect	15	
	Mean	-3.72	
	Median	-3.21	
	Minimum	-1.56	Rubber products
	Maximum	-10.36	Electrical machinery apparatus, appliances and supplies
Norway			
	Nb Sect	20	
	Mean	-4.62	
	Median	-4.49	
	Minimum	-2.95	Paper and paper products
	Maximum	-7.25	Wearing apparel, except footwear
Philippines			
	Nb Sect	15	
	Mean	-6.31	
	Median	-5.05	
	Minimum	-2.66	Beverage industries
	Maximum	-24.06	Rubber products
Portugal			
	Nb Sect	22	
	Mean	-5.87	
	Median	-4.48	
	Minimum	-2.00	Pottery, china and earthenware
	Maximum	-25.96	Non-ferous metal basic industries

Singapore			
	Nb Sect	10	
	Mean	-7.07	
	Median	-5.57	
	Minimum	-3.29	Plastic products not elsewhere classified
	Maximum	-20.46	Furniture and fixtures, except primarily of metal
Slovakia			
	Nb Sect	10	
	Mean	-4.50	
	Median	-4.42	
	Minimum	-1.86	Professional and scientific, and measuring and controlling equipment not elsewhere classified, and photographic and optical goods
	Maximum	-6.90	Fabricated metal products, exc. mach. and equip.
Spain			
	Nb Sect	26	
	Mean	-4.30	
	Median	-4.29	
	Minimum	-2.38	Furniture and fixtures, except primarily of metal
	Maximum	-6.00	Textiles
Sri Lanka			
	Nb Sect	10	
	Mean	-8.44	
	Median	-4.43	
	Minimum	-1.78	Transport equipment
	Maximum	-29.01	Textiles
Sweden			
	Nb Sect	25	
	Mean	-4.86	
	Median	-4.09	
	Minimum	-2.11	Furniture and fixtures, except primarily of metal
	Maximum	-10.28	Textiles
USA			
	Nb Sect	26	
	Mean	-5.35	
	Median	-4.39	
	Minimum	-2.57	Other non-metallic mineral products
	Maximum	-21.01	Petroleum refineries
UK			
	Nb Sect	26	
	Mean	-4.62	
	Median	-4.20	
	Minimum	-2.26	Petroleum refineries
	Maximum	-8.94	Footwear, except vulcanized or molded rubber or plastic footwear

Venezuela			
	Nb Sect	20	
	Mean	-3.92	
	Median	-3.26	
	Minimum	-1.16	Iron and steel basic industries
	Maximum	-8.07	Electrical machinery apparatus, appliances and supplies

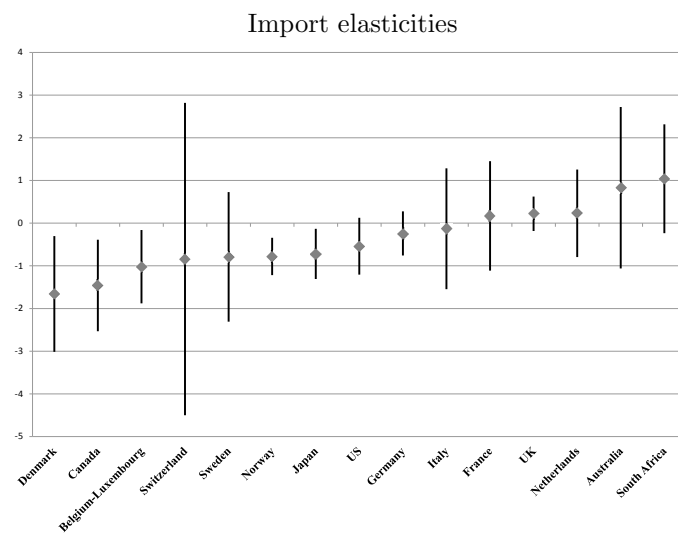
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Table 3: One-Sector Elasticity Estimates and Comparisons					
Country	$\partial \ln W_{MS}$	λ_{jj}	ϵ_{OS}	ϵ_A	ϵ_W
Australia	-0.0829(2.25E-05)	0.6926	-4.43(0.064)	-0.43(0.006)	-4.43(0.065)
Austria	-0.1601(5.21E-05)	0.5311	-3.95(0.032)	-0.81(0.006)	-3.95(0.032)
Canada	-0.0991(10.07E-04)	0.5935	-5.26(0.284)	-0.87(0.003)	-5.27(0.284)
China	-0.0308(5.58E-06)	0.8160	-6.60(0.256)	-0.62(0.007)	-6.60(0.054)
Egypt	-0.1112(4.55E-05)	0.6355	-4.08(0.061)	-0.61(0.009)	-4.08(0.256)
Finland	-0.0918(1.83E-05)	0.6547	-4.61(0.047)	-0.63(0.008)	-4.61(0.047)
France	-0.0853(4.83E-06)	0.7008	-4.17(0.012)	-0.65(0.003)	-4.17(0.012)
Germany	-0.1030(4.43E-06)	0.6943	-3.54(0.005)	-0.53(0.003)	-3.54(0.005)
Greece	-0.1735(1.05E-04)	0.5259	-3.70(0.048)	-0.49(0.007)	-3.70(0.048)
Hong Kong	-0.4033(7.52E-04)	0.1725	-4.36(0.088)	-0.59(0.009)	-4.36(0.088)
Hungary	-0.0999(3.34E-05)	0.6182	-4.82(0.078)	-0.43(0.007)	-4.82(0.078)
India	-0.0197(1.44E-06)	0.8862	-6.13(0.139)	-0.50(0.009)	-6.13(0.072)
Indonesia	-0.1618(2.37E-04)	0.6337	-2.82(0.072)	-0.70(0.018)	-2.82(0.093)
Israel	-0.0963(4.45E-05)	0.6547	-4.40(0.093)	-0.41(0.013)	-4.40(0.011)
Italy	-0.0878(5.17E-06)	0.7053	-3.98(0.011)	-0.71(0.004)	-3.98(0.034)
Japan	-0.0134(2.37E-07)	0.9350	-5.03(0.033)	-0.66(0.004)	-5.03(0.121)
Korea	-0.0465(9.20E-06)	0.7802	-5.34(0.122)	-0.40(0.009)	-5.34(0.012)
Kuwait	-0.4570(5.72E-04)	0.3841	-2.09(0.012)	-0.45(0.013)	-2.09(0.031)
Norway	-0.1228(3.04E-05)	0.6206	-3.88(0.031)	-0.52(0.007)	-3.88(0.073)
Philippines	-0.0829(3.39E-05)	0.7269	-3.85(0.073)	-0.55(0.019)	-3.85(0.050)
Portugal	-0.1271(5.45E-05)	0.6138	-3.84(0.050)	-0.56(0.006)	-3.84(0.139)

Singapore	-0.2247(5.02E-04)	0.3285	-4.95(0.244)	-0.77(0.006)	-4.95(0.244)
Slovakia	-0.0776(2.31E-05)	0.7238	-4.17(0.067)	-0.83(0.011)	-4.17(0.067)
Spain	-0.0745(7.77E-06)	0.7328	-4.17(0.024)	-0.66(0.004)	-4.17(0.024)
Sri Lanka	-0.2368(3.80E-04)	0.5117	-2.83(0.054)	-0.82(0.013)	-2.83(0.015)
Sweden	-0.1628(3.09E-05)	0.5603	-3.56(0.015)	-0.88(0.004)	-3.56(0.061)
UK	-0.1026(6.11E-06)	0.6691	-3.91(0.009)	-0.65(0.003)	-3.91(0.009)
USA	-0.0397(1.09E-06)	0.8368	-4.48(0.014)	-0.74(0.003)	-4.48(0.014)
Venezuela	-0.1190(7.90E-05)	0.6833	-3.20(0.057)	-0.49(0.011)	-3.20(0.057)

Notes: λ_{ij} is calibrated on aggregate data, defined as $(Y_j - X_j) / (Y_j - X_j + M_j)$. Numbers in parenthesis are standard errors, obtained using the Delta method on a Taylor expansion. ϵ_A denotes import price elasticity estimated with a conventional gravity equation on aggregate data. ϵ_W denotes the one-sector elasticity defined by theory as a weighted average of ϵ^S .

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 circles correspond to the confidence interval, at the 5% level.

Figure 2: Trade Elasticities ϵ_w (US=-4.48)

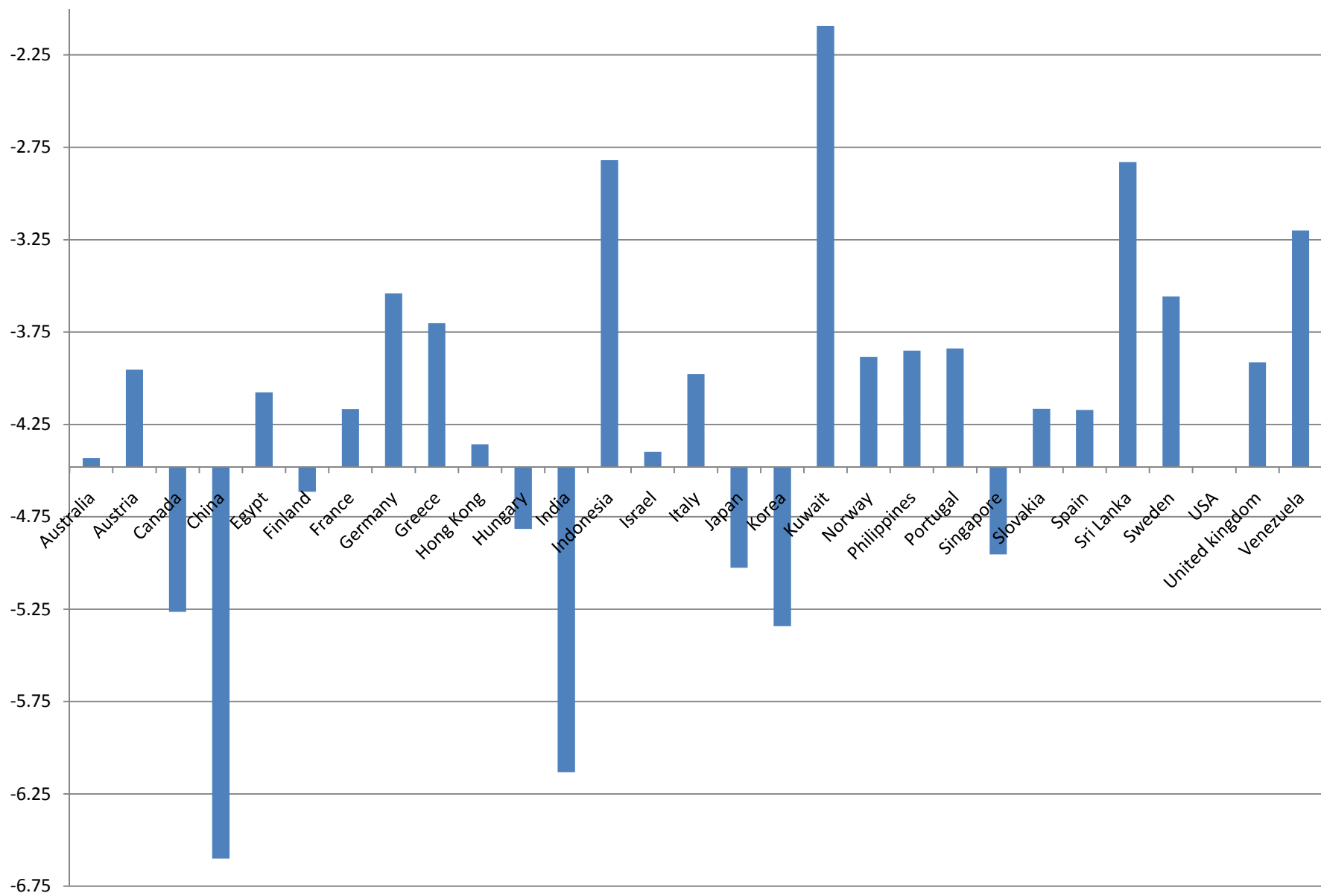


Figure 3: Decomposition of Differences with the US

