Asymmetric Trade Costs, Trade Imbalances, and the World Income Distribution

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March 25, 2012

Abstract

Earlier work by Waugh (2010) suggests that asymmetric market access costs across exporting countries are a major reason for differences in real per-capita income around the globe: $25\% - 50\%$ (depending on the measure) of world income inequality could be explained by cross-country trade cost asymmetries alone. We show that these results were driven by what we call model under-specification, an ill-suited model calibration which fails to match data on per-capita incomes well enough, and a counterfactual experiment inconsistent with the theoretical model and the idea of a reduction in country-specific market access cost asymmetries as such. We use the same (Eaton-Kortum-type) structural model and data, estimate all model parameters while respecting general equilibrium constraints, and calibrate it to the data. The obtained results suggest a largely different picture: the complete abolition of exporter-specific trade cost asymmetries leads to no more than a $1.5\% - 6\%$ reduction in the income inequality around the globe. Moreover, in the presence of large trade imbalances, a reduction in such trade cost asymmetries may even have a detrimental effect on income equalization between poor and rich countries.

Keywords: International trade; Asymmetric trade costs; Real income gap.

JEL-codes: F10; F11; F14; O1.

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Financial support by the Swiss National Science Foundation (SNSF) is gratefully acknowledged.

1 Introduction

Pioneered by Eaton and Kortum (2002), the multi-country version of the Ricardian trade model has gained increasing importance as a tool to conduct comparative static analysis with regard to the relative importance of trade costs (geography and tariffs) and technology (total factor productivity) for trade, factor prices, and welfare (see Alvarez and Lucas, 2007; Dekle, Eaton, and Kortum, 2007; Chor, 2010; Donaldson, 2010; Waugh, 2010; Caliendo and Parro, 2011; Costinot, Donaldson, and Komunjer, 2011; Fieler (2011); Levchenko and Zhang, 2011; Shikher, 2011; for a select list of applications of that model).

One merit of this type of model is its amenity to structural estimation whereby all parameters needed for general-equilibrium-consistent comparative static analysis are estimated from the same data the model is calibrated to. While this approach tends to fit the data well, it requires the imposition of general equilibrium constraints in structural estimation to ensure compliance of the fundamental assumptions of the model with the data. The latter is often, if not typically, ignored in estimation of multi-country trade models. As a result, comparative static analysis may result in quantitatively misleading conclusions accruing to a big gap between predictions of the calibrated theoretical model and the data in a benchmark equilibrium.[1]

A key constraint in all structural multi-country models of bilateral trade is the one on "goods" (potentially encompassing services) market clearance by way of a *multilateral trade balance condition* (as is underlying the models of Alvarez and Lucas, 2007; or Waugh, 2010) or a *multilateral trade imbalance condition* (see Dekle, Eaton, and Kortum, 2007, for an example).[2] We will demonstrate how important this constraint is for estimation and comparative static analysis when analyzing the question about the importance of exporting country-specific goods market access asymmetries for per-capita income differences around the globe. In a recent article, Waugh (2010) examined this question and produced two stunning findings. First, estimated exporter-specific trade costs are highly correlated with real per-capita income across countries. Second, eliminating the asymmetry in such trade costs reduces the world-wide variance in real per-capita income by 25%-50% (depending on the measure). Utilizing the same data-set and theoretical model, we arrive at a starkly different conclusion: eliminating exporter trade cost heterogeneity has only little bearing for real per-capita income variability across countries. This stark difference in the results has two main roots.

[1] In particular, suitable estimates of the vector of bilateral trade costs are vital for a quantification of the magnitude of marginal or discrete changes in trade costs on trade flows and other outcomes not only in a model as the one of Eaton and Kortum (2002) but in all so-called new trade theory models (see Krugman, 1979; Anderson and van Wincoop, 2003; Arcolakis, Costinot, and Rodriguez-Clare, 2010; for examples).

[2] The latter fixes observations on countries' trade imbalances and embeds those imbalances as parameters in the multilateral trade balance condition so that trade is balanced up to some country-specific constant.

First, general equilibrium constraints are *assumed* in the theoretical models of Eaton and Kortum (2002) and Waugh (2010), and this assumption is necessary for the identification strategy of asymmetries in exporter-specific (unobserved) trade costs in Waugh (2010). Yet, Waugh's structural empirical model does not respect this constraint (neither by imposing multilateral trade balances nor trade imbalances of a certain extent across the included countries) so that estimated exporter trade cost asymmetries partly reflect observable trade imbalances rather than trade costs alone. It turns out that poor countries in the sample display systematically greater trade deficits than developed countries so that a significant share of the correlation between the estimated exporter-specific trade costs and data on real per-capita income accrues to issues which lie beyond the utilized theoretical model. By way of general equilibrium constraints, trade imbalances affect factor prices. Both factor prices and exporter market access costs are captured by country fixed effects in estimation. A model which assumes balanced trade or does not explicitly account for imbalances among the included countries results in biased country-specific effects in the sense that they do not jointly predict data on trade flows and income (GDP) without bias but eventually only the former. From the perspective of a model of trade *and* income (when predicting per-capita income differences is at stake), this results in biased effects of exporter-specific trade costs, technology, or factor prices.

Second, the counterfactual experiment conducted in Waugh contradicts the specified model. Given geography which specifies the symmetric level of trade costs in Waugh (2010), the counterfactual level of trade costs can not be generated by country-specific changes in export trade costs. As a consequence, the counterfactual trade cost matrix can only be generated by big country-pair-specific (i.e., preferential) changes in trade cost *levels* which are inconceivable within the the adopted theoretical model and available trade instruments. Moreover, Waugh's experiment does not systematically reduce the *asymmetry* in exporter-specific trade costs, but it mainly changes the *level* of all trade costs.[3] It turns out that reducing the exporter trade cost asymmetry *per se* has only little bearing for the global dispersion of real per-capita income. Hence, offering symmetric world market access to the poor will not close their gap in per-capita income to the richest in a substantial way.

The remainder of the paper is organized as follows. The next section briefly outlines the model. Section 3 characterizes our estimation strategy and reports the results while Section 4 describes our calibration procedure. Section 5 documents findings from a counterfactual analysis and Section 6 concludes.

[3]That reductions of trade friction *levels* are correlated with the income convergence across countries is a well-established result; see Ben-David (1993), Sachs and Warner (1995), Frankel and Romer (1999), Berg and Krueger (2003), Wacziarg and Horn Welch (2003).

2 Model

We employ an Eaton-Kortum type model to infer about the importance of exporter-specific trade cost asymmetries for the dispersion of real per-capita incomes across countries. There are N countries which are indexed by i and endowed with L_i units of labor and K_i units of capital. Labor and capital are mobile domestically (across sectors) but not internationally. Each country produces in two sectors, tradable goods and non-tradables. Bilateral trade is costly and impeded by iceberg costs so that a single unit of consumption in country i requires shipping of $t_{ij} \geq 1$ units from country j.[4]

Tradable sector:

Each country hosts a *continuum* of firms in the tradable sector. Firms in the tradable sector of country i face variable (and marginal) costs of c_{ti} per efficiency unit. The firm producing variety j operates at a total factor productivity $z(j)^{-\theta}$, with $z(j)$ being drawn from a country-specific exponential distribution with mean λ_i. The technology for producing output $q_{ti}(j)$ of variety j in i may be specified as:

$$q_{ti}(j) = z(j)^{-\theta}(r_i^{1-\alpha}k_i^{\alpha})^{\beta}q_{ti}^{1-\beta}, \tag{2.1}$$

where r_i, k_i, and q_{ti} are labor, capital, and composite tradable good inputs, respectively. The composite tradable good q_{ti} is produced via a standard Spence-Dixit-Stiglitz (SDS) CES technology that combines the cheapest available goods that are either produced domestically or imported (subject to trade costs). Using properties of the exponential distribution, it is straightforward to show that the price of the composite tradable good p_{ti} is:

$$p_{ti} = \Omega_t \sum_j \left(\lambda_j \left((w_j^{1-\alpha}r_j^{\alpha})^{\beta}p_{tj}^{1-\beta}t_{ij}\right)^{-\frac{1}{\theta}}\right)^{-\theta} = \Omega_t \left(\sum_j (c_{tj}t_{ij})^{-\frac{1}{\theta}}\right)^{-\theta}, \tag{2.2}$$

where Ω_t is a sector-specific constant, w_j is the wage rate in j, r_j is the returns to capital in j, p_{tj} is the price of tradable inputs in j, and t_{ij} are trade costs on tradables produced in j and shipped to i. Accordingly, $c_{ti} = \lambda_i^{-\theta}(w_i^{1-\alpha}r_i^{\alpha})^{\beta}p_{ti}^{1-\beta}$ are the marginal (and variable) costs of a producer of tradable goods in i.

Nontradable sector:

We assume that each country runs *one* representative firm in the non-tradable sector which operates at an identical Cobb-Douglas technology world-wide. Output of the non-tradable

[4]We make the usual triangularity assumption that rules out opportunities for arbitrage.

final good q_{ni} in country i is produced at

$$q_{ni} = (r_i^{1-\alpha} k_i^{\alpha})^{\gamma} q_{ti}^{1-\gamma}. \tag{2.3}$$

The price of the final non-tradable good is then:

$$p_{ni} = \Omega_n (w_i^{1-\alpha} r_i^{\alpha})^{\gamma} p_{ti}^{1-\gamma} = \Omega_n c_{ni}, \tag{2.4}$$

where Ω_n is a sector-specific constant and c_{ni} are the variable (and marginal) unit costs a firm in country i's nontradable sector faces.

International trade:

Recall that $c_{ti} = \lambda_i^{-\theta} (w_i^{1-\alpha} r_i^{\alpha})^{\beta} p_{ti}^{1-\beta}$ is the marginal (and variable) cost of a producer of tradable goods in i. Use M_{ij} to denote aggregate nominal imports of country i from j and Y_{ti} to denote i's total spending on tradables. Then, through the properties of the exponential distribution, country i spends a share $X_{ij} \equiv M_{ij}/Y_{ti}$ on tradable goods from j, and

$$X_{ij} = \frac{(c_{tj} t_{ij})^{-\frac{1}{\theta}}}{\sum_{\ell} (c_{t\ell} t_{i\ell})^{-\frac{1}{\theta}}}. \tag{2.5}$$

Goods market clearance:

All prices can be expressed in terms of the primitives of the model L_i, K_i, t_{ij}, λ_i and other technology parameters, and a wage vector with typical element w_i. It is customary to solve for w_i via the international goods market clearance condition which simply states that total expenditures of country i on tradables must equal total sales of all countries j (including i) to customers in i. This market clearance condition can be stated for balanced trade between all countries j (see Eaton and Kortum, 2002; Alvarez and Lucas, 2007) or unbalanced trade (see Dekle, Eaton, and Kortum, 2007). Since goods trade is not balanced among the countries included in most data-sets (as the one we use), a natural way to proceed is to assume imbalanced trade upfront.

Let $T_{ti} = \sum_{j=1}^{N} M_{ji}$ denote the $t$otal value of tradable goods produced in i and $Y_{ti} = \sum_{j=1}^{N} M_{ij}$ be country i's total value of consumption of tradables. Then, we may define

$$B_i = \frac{\sum_{j=1}^{N} (M_{ij} - M_{ji})}{\sum_{j=1}^{N} M_{ij}} \equiv \frac{Y_{ti} - T_{ti}}{Y_{ti}} \tag{2.6}$$

as a measure of normalized (by apparent consumption) multilateral trade deficit. Hence, we simply state that the ratio of i's total income to expenditure is some constant $(1 - B_i)$.[5]

At an observed value of B_i and after replacing $M_{ji} = X_{ji}Y_{tj}$ in (2.6), we can derive the aggregate market clearance condition:

$$L_i w_i = \sum_{j=1}^{N} X_{ji} \frac{L_j w_j}{1 - B_j}. \tag{2.7}$$

Of course, at $B_i = 0$ for all countries i, we are in a standard equilibrium with balanced trade as in Eaton and Kortum, Alvarez and Lucas (2007), or Waugh (2010).

We can classify the fundamentals of the model into three categories. The first category consists of the *observable* endowments of labor and capital $\{L_i, K_i\}$. The second category includes unobservable primitives, namely the bilateral trade cost parameters $\{\tau_{ij}, \xi_j\}$, that have to be estimated from a structural stochastic version of equation (2.5). Finally, technology parameters $\{\alpha, \beta, \gamma, \lambda_i, \theta\}$ have to be either estimated from stochastic versions of equations (2.1) and (2.4) or otherwise calibrated to the data. We will discuss in particular the calibration of λ_i in Section 4 and also the choice of the remaining technology parameters in the Appendix.

3 Estimation

In this section, we show how to estimate $\{\tau_{ij}, \xi_j\}$ consistently with all assumptions of the model. If we normalize (2.5) by the share of domestic sales of tradables in total spending on tradables, $X_{ii} \equiv M_{ii}/Y_{ti} = 1 - \sum_{j \neq i}^{N} X_{ij}$, we arrive at the deterministic relative gravity equation:

$$\frac{X_{ij}}{X_{ii}} = \left(\frac{c_{tj} t_{ij}}{c_{ti} t_{ii}} \right)^{-\frac{1}{\theta}}. \tag{3.1}$$

The stochastic counterpart to (3.1) can then be expressed as:

$$\frac{X_{ij}}{X_{ii}} = \exp\left(s_j - s_i + \xi_j + \sum_{k=1}^{K} \beta_k \tau_{k,ij} + u_{ij} \right) \text{ or } \ln \frac{X_{ij}}{X_{ii}} = s_j - s_i + \xi_j + \sum_{k=1}^{K} \beta_k \tau_{k,ij} + v_{ij}, \tag{3.2}$$

where s_j and s_i are country-specific catch-all variables associated with producer prices and technology parameters of countries j and i, respectively. By design, $s_j = s_i$ whenever $i = j$.

[5]Notice that this representation of market clearance is equivalent but not identical to the one in Dekle, Eaton, and Kortum (2007). They fix the absolute amount of nominal trade deficit in country i as a parameter $D_i \equiv \sum_{j=1}^{N}(M_{ij} - M_{ji})$. We do the same such that B_i adjusts properly at given D_i. In any equilibrium, the two representations can be easily transformed one into another.

u_{ij} and v_{ij} are stochastic terms for the models in levels and logs, respectively.[6]

In estimating (3.1), structural empirical work typically adopts three assumptions:

(i) The stochastic counterpart to (3.1) involves either an additive or a multiplicative error term as in (3.2).

(ii) Trade costs $t_{ij}^{-\frac{1}{\theta}}$ are modeled multiplicatively as $t_{ij}^{-\frac{1}{\theta}} = \exp\left(\xi_j + \sum_{k=1}^{K} \beta_k \tau_{k,ij}\right)$, where $\tau_{k,ij}$ is the kth observable measure of log trade costs such as log bilateral distance or an adjacency indicator and ξ_j is a measure of unobservable (net) asymmetric trade costs for exporter j.[7] Let us denote total symmetric trade costs as $\tau_{ij} \equiv \sum_{k=1}^{K} \beta_k \tau_{k,ij}$, then we can express total iceberg trade costs as:

$$t_{ij} \equiv \exp\left(-\theta\xi_j - \theta\tau_{ij}\right). \tag{3.3}$$

(iii) The symmetric component of trade costs τ_{ij} can be specified as

$$\tau_{ij} = \beta_1 adjacency_{ij} + \beta_2 \ln(distance_{1,ij}) + \ldots + \beta_7 \ln(distance_{6,ij}), \tag{3.4}$$

where $distance_{s,ij}$ is an indicator variable referring to the sth sextile of the distribution of the great circle distance between two countries i and j. Alternatively, one could replace $\beta_2 \ln(distance_{1,ij}) + \ldots + \beta_7 \ln(distance_{6,ij})$ by $\psi \ln(distance_{ij})$.

Notice that the model principally assumes that $t_{ij} \geq 1$. However, this assumption is missing from the above list, hence, not translated into an ex-ante constraint in the estimation procedure. As a consequence, there is no guarantee that the predicted total trade costs adhere to this assumption. Some authors such as Waugh (2010) adjust trade costs for those country-pairs ij where $t_{ij} < 1$ ex post. Yet, it is clear that an estimator with a properly specified ex-ante constraint of $(\min \hat{t}_{ij} \geq 1)$ will lead to different estimates t_{ij} than such a procedure. Clearly, the latter involves a gap between estimation and model calibration while the former does not. estimator will that [8]) but that leads to biased marginal effects and is not in line with the general equilibrium approach.

Even more importantly, notice that estimating (3.2) based on $N(N-1)$ observations for country-pairs (N equations on pairs all ii are disregarded) *without* imposing goods market clearance instead of a stochastic version of (2.5) $N(N-1)$ observations subject to N con-

[6]See Santos Silva and Tenreyro (2006) on how the two translate into each other under specific assumptions with exponential-family-type econometric models.

[7]One could include *importer fixed effects* by replacing ξ_j by ξ_i instead. However, we follow Waugh (2010) through including *exporter fixed effects* so that $t_{ij}^{-\frac{1}{\theta}}$ is defined exactly as in (ii).

[8]Waugh (2010) estimates log-linear version of (3.2) and has to manually adjust all $\hat{t}_{ij}$ that are smaller than unity.

straints (2.7) is not the same and will almost always give very different results. The reason for this is that intranational sales X_{ii} are very important and predicting bilateral international trade flows well does not guarantee to predict income or total goods sales T_{ti} or total goods expenditures Y_{ti} well. In some sense, disregarding (2.7) entails a model under-specification. Clearly, an approach which respects the N constraints (2.7) should be able to match model estimates on total income (GDP) and, hence, real per capita income much better than one that does not.

In the context of the notation used in (3.2), we can rewrite the goods market clearance constraint in (2.7) as

$$Y_{ti} = \sum_{j=1}^{N} \frac{\exp(s_i + \xi_i + \tau_{ji})}{\sum_{\ell=1}^{N} \exp(s_\ell + \xi_\ell + \tau_{j\ell})} \frac{Y_{tj},}{1 - B_j}, \tag{3.5}$$

where Y_{tj} are the observed total sales of goods in country j. We argue that an imposition of N constraints (3.5) in conjunction with the $N(N-1)$ observations on international trade in (3.2) will eventually lead to different estimates of s_i, ξ_i, and τ_{ji} than estimating (3.2) without this constraint. In particular, imposition of (3.5) will give the prediction of Y_{ti} (which is proportional to total income and real income per capita) some weight which should ensure that total income and, in turn, per-capita income may be predicted better by the structural model than when disregarding (3.5) in estimation. Hence, we may summarize the two fundamental reasons for why respecting (3.5) is important as follows.

(i) The constraint (3.5) serves as an additional moment condition in (3.5) and allows to use observations on Y_{ti}, which is proportional to real GDP in the current context. As a consequence, the constrained model may predict real per-capita income as the variable of interest considerably better than the unconstrained one.

(ii) Estimating (3.2) subject to (3.5) may lead to different estimates of s_i, ξ_j, τ_{ij}, which affects the quantitative effects of changes of, say, ξ_j or τ_{ij} in general equilibrium.

Formally, we estimate the model in levels in (3.2) by maximizing the corresponding Poisson Pseudo Maximum Likelihood objective function subject to three constraints: market clearance by way of (3.5) where Y_{ti} is total manufacturing absorption (which can be substituted by total real GDP); that the individual ξ_j sum up to zero (as in Waugh, 2010); and that $(\min \hat{t}_{ij} \geq 1)$.

It turns out that there are considerable differences between the estimates of s_i and ξ_j in a model which disregards the constraint in (3.5) versus one that does not. We report estimates of the exporter-specific asymmetric (log) trade costs, $\hat{\xi}_j$, the fixed effects $\hat{s}_i$, and the parameters of the symmetric component of trade costs $\tau_{k,ij}$ for (i) the PPML model which disregards the market clearance constraint (3.5), (ii) constrained PPML which imposes market clearance but assumes $B_i = 0$ for all i, and (iii) PPML which imposes market clearance with $B_i \neq 0$. The coresponding results are summarized in Table 1.

There are large differences in the estimated exporter effects $\hat{\xi}_j$ between the unconstrained and the constrained estimates. For example, unconstrained results as in (i) suggest that $100 \times (e^{-\theta\hat{\xi}_j} - 1)$ – i.e., the observable trade cost amplification factor of unobservable exporter-specific trade costs – of Benin and Rwanda amounts to 78% and 103%, respectively. The corresponding numbers for these two countries in the constrained model as in (ii) are 45% and 62%, respectively. These numbers are 47% and 57%, respectively, in the constrained model as in (iii). Hence, we conclude that ignoring goods market clearance biases coefficients of fundamental interest to a considerable degree.[9]

4 Calibration

Estimating ξ_j and τ_{ij} consistent with the theoretical model is important but not enough. We also need the parameters $\{\alpha, \beta, \gamma, \lambda_i, \theta\}$ in order to conduct counterfactual experiments. Let us relegate the choice of parameters $\{\alpha, \beta, \gamma, \theta\}$ to the Appendix and focus on the calibration of country-specific productivity parameter λ_i, here.

First, given trade flows X_{ji} for all ji and L_j and B_j for all j, we can solve for the vector of wages with element w_i from the market clearance condition:[10]

$$L_i w_i = \sum_{j=1}^{N} X_{ji} \frac{L_j w_j}{1 - B_j}. \tag{4.1}$$

Given w_i, we can express the price of tradables p_{ti} in (2.2) and returns to capital r_i from the first-order conditions of the final good producer. Recall that $\exp(-\theta s_i) \equiv c_{tj} = \lambda_i^{-\theta}(w_i^{1-\alpha} r_i^{\alpha})^{\beta} p_{ti}^{1-\beta}$. Hence, given solutions for w_i, r_i and p_{ti} we can recover λ_i from s_i.[11] The crucial point in the calibration procedure is selecting parameters so as to match data on the key variables of interest in the best possible way.

Balanced Trade

In reality trade is <u>not</u> balanced. In the sample of 77 countries used in this study the average value of B_i is 0.16 with a standard deviation of 0.21. Imbalances are strongly correlated with

[9] Our results do not depend crucially on whether one estimates (3.2) in logs via OLS or in levels via PPML (see Santos Silva and Tenreyro (2006) for a discussion of the difference between the two approaches).

[10] Clearly, for solutions of w_i which are consistent with the estimated model, we need to use predictions $\hat{X}_{ji}$ rather than observations X_{ji} in (4.1).

[11] In practice, this calibration involves an iterative procedure until the fixed point in $\{w_i, r_i, p_{ti}, \lambda_i\}$ is found.

Table 1: PPML ESTIMATES

Country	Unconstrained Estimation PPML				Constrained Estimation PPML, $B_i = 0$				PPML, $B_i \neq 0$			
	s_i		ξ_i		s_i		ξ_i		s_i		ξ_i	
United States	0.91	(0.14)	3.43	(0.29)	1.87	(0.12)	2.61	(0.27)	1.35	(0.11)	2.58	(0.26)
Argentina	0.45	(0.08)	1.21	(0.11)	1.37	(0.08)	-0.06	(0.14)	1.36	(0.10)	-0.50	(0.15)
Australia	-0.09	(0.08)	4.08	(0.42)	-0.02	(0.09)	3.08	(0.44)	-0.49	(0.09)	2.95	(0.46)
Austria	1.03	(0.05)	0.45	(0.14)	1.28	(0.05)	-0.29	(0.15)	0.86	(0.06)	-0.43	(0.15)
Belgium	-0.59	(0.27)	4.22	(0.32)	-0.75	(0.20)	2.91	(0.44)	-1.26	(0.19)	2.79	(0.46)
Benin	-0.85	(0.11)	-5.36	(0.22)	-0.45	(0.11)	-1.42	(0.22)	-0.76	(0.11)	-1.74	(0.24)
Bangladesh	0.71	(0.06)	-1.01	(0.05)	0.97	(0.05)	-1.44	(0.06)	1.31	(0.08)	-1.18	(0.09)
Bolivia	-0.58	(0.08)	-1.57	(0.09)	-0.03	(0.08)	-0.94	(0.10)	-0.36	(0.08)	-0.76	(0.10)
Brazil	1.02	(0.05)	1.66	(0.08)	1.87	(0.06)	0.21	(0.11)	1.78	(0.06)	-0.16	(0.12)
Central African Republic	-0.26	(0.12)	-3.49	(0.15)	0.11	(0.11)	-3.45	(0.14)	-0.12	(0.11)	-2.72	(0.13)
Canada	0.78	(0.57)	2.20	(0.60)	0.92	(0.26)	1.45	(0.32)	0.57	(0.29)	1.31	(0.35)
Switzerland	1.25	(0.20)	0.89	(0.24)	1.32	(0.15)	0.11	(0.21)	0.57	(0.15)	-0.02	(0.19)
Chile	-0.42	(0.09)	1.59	(0.10)	0.51	(0.09)	0.66	(0.12)	0.26	(0.09)	0.45	(0.13)
China-Hong Kong	1.45	(0.13)	3.36	(0.16)	0.54	(0.19)	3.12	(0.24)	1.27	(0.08)	2.01	(0.14)
Cameroon	-0.44	(0.10)	-0.33	(0.16)	-0.08	(0.09)	-1.38	(0.14)	-0.21	(0.09)	-0.58	(0.14)
Colombia	0.41	(0.09)	0.15	(0.21)	1.26	(0.11)	-0.81	(0.26)	1.39	(0.16)	-1.27	(0.31)
Costa Rica	-0.39	(0.19)	0.55	(0.34)	0.04	(0.12)	-0.47	(0.34)	-0.37	(0.11)	-0.85	(0.32)
Denmark	0.52	(0.08)	1.57	(0.16)	0.50	(0.05)	0.83	(0.16)	0.00	(0.06)	0.77	(0.17)
Dominican Republic	-0.78	(0.30)	-1.01	(0.32)	-0.28	(0.23)	0.16	(0.27)	-0.61	(0.23)	0.31	(0.25)
Ecuador	-0.19	(0.09)	-1.29	(0.12)	0.38	(0.08)	-0.84	(0.12)	0.19	(0.08)	-0.70	(0.12)
Egypt	1.12	(0.06)	-2.06	(0.12)	2.09	(0.06)	-2.32	(0.16)	2.17	(0.08)	-2.78	(0.18)
Spain	1.65	(0.16)	1.14	(0.23)	1.92	(0.13)	0.04	(0.21)	1.22	(0.17)	0.14	(0.25)
Ethiopia	-1.14	(0.06)	-4.32	(0.11)	-0.83	(0.06)	-0.36	(0.11)	-1.03	(0.06)	0.41	(0.14)
Finland	1.24	(0.07)	0.70	(0.07)	1.17	(0.06)	-0.25	(0.07)	0.58	(0.04)	-0.34	(0.07)
France	1.58	(0.16)	2.97	(0.64)	0.61	(0.17)	3.03	(0.72)	-0.19	(0.19)	3.18	(0.79)
United Kingdom	0.90	(0.12)	3.19	(0.36)	0.89	(0.09)	2.40	(0.38)	0.36	(0.09)	2.38	(0.39)
Ghana	-1.28	(0.15)	-2.17	(0.18)	-0.89	(0.13)	0.12	(0.16)	-1.19	(0.12)	0.39	(0.16)
Greece	0.90	(0.08)	-0.29	(0.14)	1.43	(0.07)	-0.87	(0.14)	0.85	(0.06)	-0.92	(0.13)
Guatemala	-0.14	(0.14)	1.01	(0.21)	0.12	(0.07)	-0.64	(0.20)	-0.14	(0.07)	-0.44	(0.18)
Honduras	-0.89	(0.27)	-0.56	(0.30)	-0.48	(0.17)	-0.37	(0.20)	-0.82	(0.17)	-0.83	(0.20)
India	1.85	(0.12)	1.57	(0.13)	0.70	(0.18)	1.50	(0.21)	1.88	(0.14)	0.21	(0.15)
Ireland	-0.30	(0.13)	2.14	(0.15)	-0.33	(0.13)	1.41	(0.14)	-0.81	(0.13)	1.29	(0.14)
Iran	1.37	(0.09)	-3.43	(0.09)	2.41	(0.08)	-2.86	(0.10)	2.78	(0.09)	-3.52	(0.12)
Israel	-0.27	(0.09)	1.53	(0.16)	0.25	(0.09)	0.91	(0.18)	-0.21	(0.09)	0.66	(0.19)
Italy	2.03	(0.16)	1.60	(0.27)	1.68	(0.15)	1.09	(0.27)	1.15	(0.19)	1.05	(0.30)
Jamaica	-0.91	(0.30)	0.17	(0.33)	-0.48	(0.21)	-0.25	(0.24)	-0.82	(0.20)	-0.82	(0.23)
Jordan	0.11	(0.06)	-2.77	(0.13)	0.47	(0.05)	-1.53	(0.15)	0.11	(0.06)	-1.93	(0.17)
Japan	2.31	(0.11)	3.05	(0.26)	1.58	(0.10)	2.84	(0.25)	0.42	(0.11)	3.43	(0.26)
Kenya	-1.04	(0.06)	1.92	(0.11)	-0.76	(0.06)	0.29	(0.12)	-0.96	(0.06)	0.75	(0.12)
Republic of Korea	1.72	(0.14)	2.72	(0.25)	0.64	(0.16)	2.70	(0.28)	-0.16	(0.18)	2.66	(0.31)
Sri Lanka	-1.28	(0.06)	0.09	(0.08)	-0.95	(0.06)	0.77	(0.08)	-1.22	(0.06)	1.91	(0.10)
Mexico	0.27	(0.35)	1.96	(0.36)	1.29	(0.35)	0.60	(0.37)	1.28	(0.45)	0.20	(0.47)
Mali	-0.34	(0.15)	-6.21	(0.27)	0.05	(0.14)	-2.47	(0.25)	-0.24	(0.15)	-2.09	(0.28)
Mozambique	-2.42	(0.10)	-0.52	(0.12)	-2.20	(0.14)	0.41	(0.14)	-2.52	(0.15)	0.83	(0.14)
Mauritius	-1.19	(0.06)	0.19	(0.08)	-0.86	(0.07)	0.34	(0.08)	-1.14	(0.07)	0.31	(0.09)
Malawi	-1.33	(0.18)	0.29	(0.22)	-1.01	(0.17)	-0.95	(0.23)	-1.26	(0.17)	-0.54	(0.22)
Malaysia-Singapore	-0.59	(0.15)	4.40	(0.17)	0.09	(0.13)	3.06	(0.15)	-0.92	(0.13)	2.76	(0.17)
Niger	-3.29	(0.38)	-2.49	(0.46)	-2.92	(0.34)	-0.23	(0.40)	-3.23	(0.35)	0.44	(0.39)
Nicaragua	-0.25	(0.10)	-1.06	(0.15)	0.17	(0.08)	-1.87	(0.12)	-0.18	(0.08)	-1.85	(0.13)
Netherlands	0.11	(0.10)	3.20	(0.17)	0.03	(0.09)	2.51	(0.17)	-0.56	(0.09)	2.48	(0.16)
Norway	0.83	(0.07)	0.74	(0.08)	0.96	(0.06)	0.08	(0.08)	0.35	(0.05)	-0.16	(0.09)
Nepal	0.61	(0.09)	-4.68	(0.08)	0.80	(0.10)	-3.49	(0.10)	0.71	(0.12)	-2.24	(0.12)
New Zealand	-0.20	(0.15)	2.41	(0.20)	0.02	(0.11)	1.31	(0.17)	-0.48	(0.10)	1.14	(0.16)
Pakistan	0.49	(0.16)	1.67	(0.21)	0.38	(0.17)	-0.25	(0.33)	0.61	(0.18)	0.32	(0.22)
Panama	-2.50	(0.37)	1.72	(0.39)	-2.08	(0.53)	3.10	(0.59)	-2.39	(0.52)	0.08	(0.58)
Peru	-0.01	(0.07)	-0.06	(0.13)	0.96	(0.07)	-0.77	(0.13)	0.99	(0.09)	-1.10	(0.12)
Philippines	-0.29	(0.14)	1.01	(0.18)	0.26	(0.12)	1.13	(0.16)	0.13	(0.13)	1.00	(0.18)
Papua New Guinea	-2.23	(0.25)	0.57	(0.30)	-1.93	(0.23)	1.04	(0.28)	-2.25	(0.23)	2.06	(0.28)
Portugal	0.84	(0.08)	1.01	(0.21)	1.03	(0.08)	0.07	(0.21)	0.36	(0.06)	0.09	(0.20)
Paraguay	-0.68	(0.10)	-1.23	(0.10)	-0.21	(0.10)	0.13	(0.11)	-0.51	(0.09)	-0.34	(0.11)
Rwanda	0.22	(0.05)	-7.32	(0.05)	0.61	(0.05)	-3.68	(0.06)	0.44	(0.05)	-3.19	(0.08)
Senegal	-0.45	(0.17)	-2.26	(0.22)	-0.06	(0.16)	-1.48	(0.22)	-0.36	(0.16)	-1.24	(0.20)
Sierra Leone	-1.99	(0.22)	-1.53	(0.29)	-1.64	(0.18)	-1.47	(0.25)	-1.96	(0.17)	0.35	(0.23)
El Salvador	-0.91	(0.20)	1.21	(0.25)	-0.53	(0.10)	-0.23	(0.16)	-0.86	(0.09)	0.00	(0.15)
Sweden	0.89	(0.07)	1.88	(0.09)	0.77	(0.06)	1.04	(0.08)	0.13	(0.05)	1.04	(0.09)
Syrian Arab Republic	1.81	(0.05)	-6.39	(0.10)	2.66	(0.04)	-3.98	(0.12)	2.60	(0.06)	-4.86	(0.14)
Togo	-0.78	(0.12)	-5.54	(0.13)	-0.39	(0.11)	-1.90	(0.11)	-0.69	(0.11)	-2.27	(0.11)
Thailand	0.63	(0.11)	0.84	(0.11)	1.44	(0.13)	0.07	(0.13)	1.33	(0.18)	-0.40	(0.18)
Tunisia	0.81	(0.09)	-1.09	(0.14)	0.99	(0.10)	-1.41	(0.15)	0.58	(0.10)	-1.21	(0.15)
Turkey	1.33	(0.09)	-0.40	(0.15)	2.09	(0.07)	-1.53	(0.17)	2.08	(0.08)	-1.93	(0.17)
Uganda	-0.16	(0.07)	-4.35	(0.07)	0.23	(0.06)	-1.98	(0.06)	0.11	(0.08)	-1.59	(0.08)
Uruguay	-0.15	(0.08)	-0.34	(0.10)	0.33	(0.07)	-0.65	(0.11)	0.06	(0.08)	-0.72	(0.12)
Venezuela	-0.41	(0.17)	0.69	(0.19)	0.19	(0.13)	0.30	(0.19)	0.13	(0.19)	0.48	(0.21)
South Africa	0.10	(0.05)	3.53	(0.13)	-0.53	(0.08)	3.20	(0.20)	-0.77	(0.07)	3.07	(0.21)
Democratic Republic of the Congo	-0.93	(0.07)	-0.59	(0.08)	-0.53	(0.07)	-0.79	(0.09)	-0.69	(0.06)	0.08	(0.11)
Zambia	-1.30	(0.17)	-1.66	(0.18)	-1.05	(0.18)	-0.38	(0.23)	-1.35	(0.18)	0.20	(0.24)
Zimbabwe	-0.01	(0.09)	0.91	(0.13)	-0.03	(0.12)	-0.53	(0.25)	-0.25	(0.13)	-0.36	(0.21)
$[0, 375)$	-5.50	(0.20)	-	(0.20)	-4.96	(0.25)	-	-	-5.00	(0.29)	-	-
$[375, 750)$	-5.85	(0.11)	-	(0.11)	-5.22	(0.13)	-	-	-5.12	(0.14)	-	-
$[750, 1500)$	-5.87	(0.07)	-	(0.07)	-5.00	(0.07)	-	-	-5.12	(0.08)	-	-
$[1500, 3000)$	-6.81	(0.13)	-	(0.13)	-5.59	(0.12)	-	-	-5.35	(0.12)	-	-
$[3000, 6000)$	-7.83	(0.12)	-	(0.12)	-6.70	(0.13)	-	-	-6.49	(0.13)	-	-
$[6000, max)$	-8.61	(0.13)	-	(0.13)	-7.38	(0.14)	-	-	-7.13	(0.14)	-	-
Adjacency	0.93	(0.12)	-	(0.12)	1.10	(0.15)	-	-	1.23	(0.16)	-	-
$Pseudo - R^2$	0.79	-	-	-	0.75	-	-	-	0.73	-	-	-

Notes: Standard errors are reported in parentheses and are based on Eicker-White sandwich estimates. The reported $Pseudo - R^2$ corresponds to the correlation between observed and predicted values of the dependent variable.

real GDP per capita (the correlation coefficient is -0.54) and are higher for less developed countries. For some countries, trade imbalances are as high as 80% of total absorption in manufacturing.

However, it may be convenient to assume that trade is balanced ($B_i = 0$ for all i) nonetheless for two reasons. First, the solution of the model becomes slightly easier computationally. Second, disregarding imbalances may result in comparative static effects which are, on average, not too dissimilar from the ones of a model which allows for imbalanced trade. For the data at hand, assuming multilaterally balanced trade does not entail a severe bias if this assumption is imposed in the estimation and calibration procedures. Assuming balanced trade in estimation of (3.2) is straightforward. We simply modify the constraint in (3.5) by setting $B_i = 0$ for all i so as to obtain

$$Y_{ti} = \sum_{j=1}^{N} \frac{\exp(s_i + \xi_i + \tau_{ji})}{\sum_{\ell=1}^{N} \exp(s_\ell + \xi_\ell + \tau_{j\ell})} Y_{tj}. \tag{4.2}$$

Notice that in the presence of imbalanced trade in the data the imposition of balanced trade in equilibrium renders the use of $\hat{X}_{ji}$ instead of X_{ji} particularly pertinent in (4.1). In what follows, we refer to the approach which assumes balanced trade in the estimation of (3.2) subject to (3.5) given data on Y_{ti} and the corresponding calibration of λ_i as **Model A**.

Unbalanced Trade

The data on B_i are observable. With this approach, we estimate (3.2) subject to (3.5) given the data on B_i and Y_{ti}. Here, we can use data X_{ij} rather than predictions $\hat{X}_{ij}$ in (4.1) to recover λ_i, since B_i inter alia absorbs all differences between X_{ij} and $\hat{X}_{ij}$ in the aggregate for a country. In what follows, we refer to this approach as **Model B**.

We argue that Models A and B significantly outperform more common estimation/calibration approaches such as running (3.2) without using the information in (3.5) and calibrating the model to the data on trade shares ignoring imbalances in (4.1) (**model C**) as in Waugh

(2010),[12] or running (3.2) without using the information in (3.5) and calibrating the model to the data on trade shares and imbalances in (4.1) (**model D**).

4.1 Quantitative comparison of Models A-D

To demonstrate that Models A and B are preferable over Models C and D, we compare the fit of calibration between these models with respect to the central variable of interest, namely real GDP per capita, y_i. In Figure 1, we plot the deviation of the prediction of y_i from its true value (we define it as a ratio of true y_i to the model's prediction) against the data on imbalances B_i using Waugh's original estimates by way of Model C (left panel) and Model A (right panel).

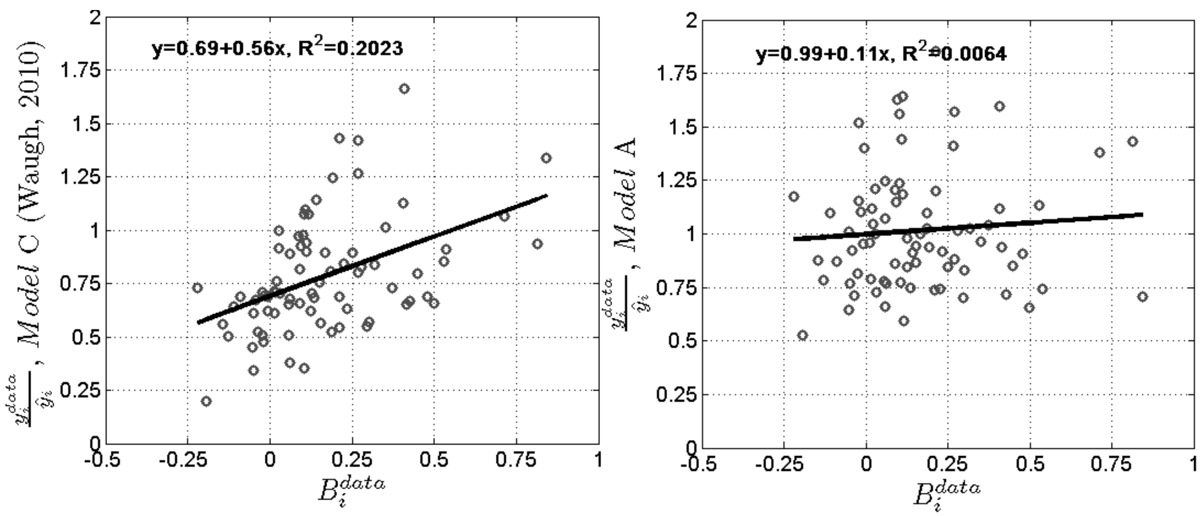

Figure 1: Fit of Calibration: Waugh (2010) versus Model A

The left panel clearly indicates that y_i is consistently overpredicted in Waugh (2010) for those countries with higher trade imbalances and underpredicted for those with lower imbalances. On the other hand, consistent with the assumption of balanced trade Model A's error in predicting y_i is purely random with respect to the data on imbalances B_i.

[12] Waugh (2010) assumes that the trade is balanced in his theoretical model. However, he estimates a log-linear version of (3.2) without the constraint in (3.5) and calibrates the model based on $B_i = 0$ for all i to the data on trade shares X_{ij}. This is exactly what is done in Model C. Accordingly Model C is based on OLS, while Models A, B, and D are based on PPML. Moreover, Models A, B, and D use a calibration which fits the data better than the one used by Waugh (2010) and Model C (see the Appendix for details). Hence, there is a discrepancy between Model C and the other ones which relates also to the estimator and parametrization used rather than the structural model implementation alone. However, we can show that neither using OLS versus PPML nor using one or the other calibration is elemental for the differences in the comparative static results between those models. But rather the main drivers are the lack of imposition of (3.5) which leads to biased estimates of s_j, ξ_j, and τ_{ij} and the use of unbalanced trade *data* X_{ij} rather than balanced model predictions $\hat{X}_{ij}$ in calibrating λ_i through (4.1). Hence, there are fundamental gaps between the theoretical model and its implementation which lead to predictions of y_i that are biased as will be shown below.

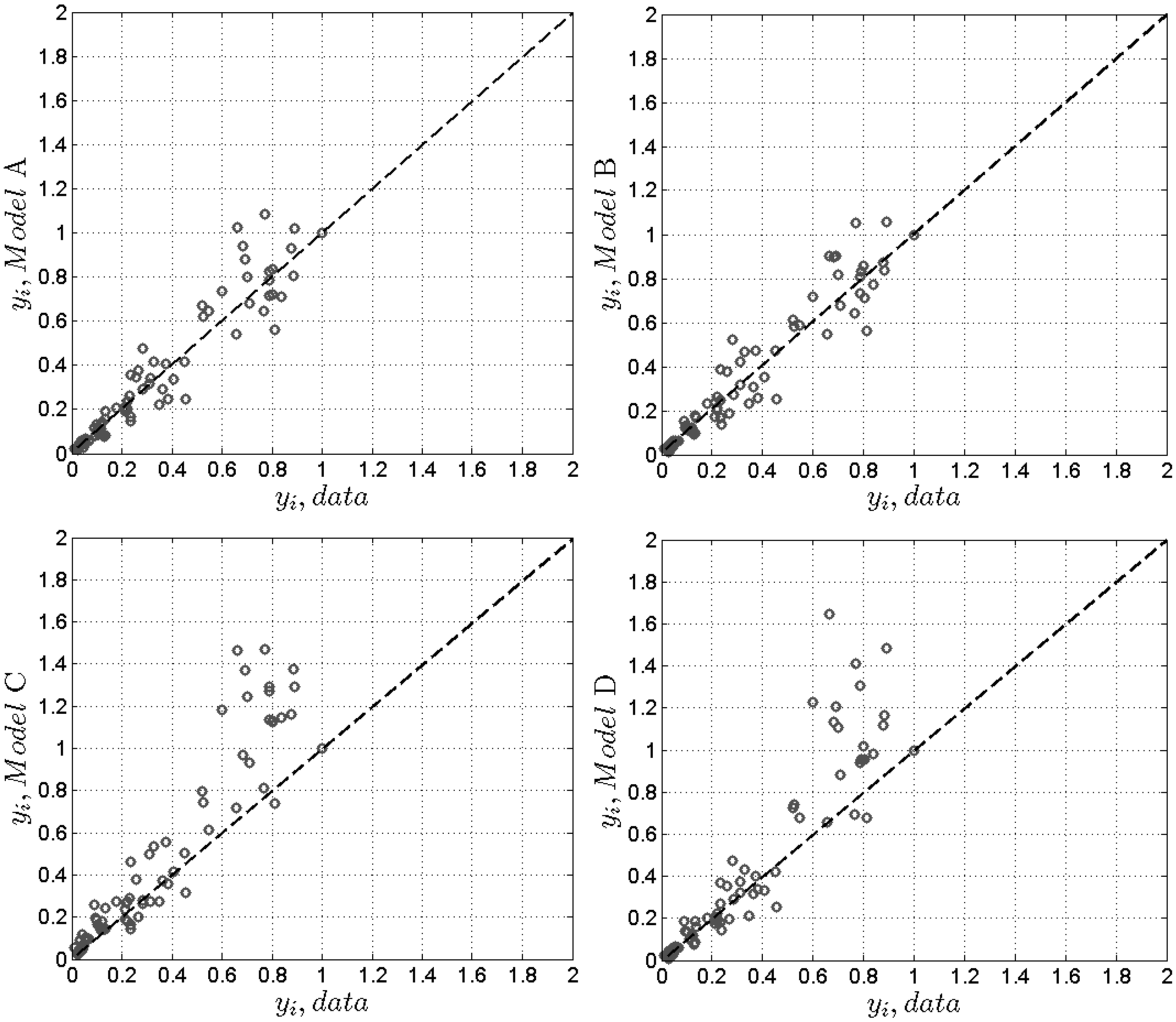

Figure 2: Fit of Calibration: Models A-D

The difference becomes even more transparent in Figure 2, where we plot the data on y_i versus predictions of Models A-D. First, predictions of Model A and Model B are very close to each other and predict the data well for both poor and rich countries. Model C (Waugh's, 2010, benchmark model) consistently overpredicts y_i for both rich and poor countries but by a considerably larger margin for rich countries. Finally, Model D does better than Model C in terms of poor countries and worse in terms of rich countries.

We also compare the predictions of these four models with respect to y_i in Table 5, where we report the first two moments of the distribution of y_i and two inequality measures that we will employ in the counterfactual experiments below: the variance in log real per-capita income across countries ($\text{var}(\ln y_i)$) and the 90-to-10 centile gap in real per capita income (y_{90}/y_{10}).

Quite obviously, Models A and B significantly outperform Model C (Waugh, 2010) and Model D in terms of matching the data on y_i.

Table 2: Fit of Calibration

Statistics	**Data**	Model A	Model B	Model C	Model D
Mean	**0.34**	0.35	0.36	0.47	0.43
Variance	**0.09**	0.10	0.10	0.20	0.19
$\text{var}(\ln y_i)$	**1.38**	1.39	1.37	1.30	1.67
y_{90}/y_{10}	**25.6**	20.9	21.3	25.7	29.3

5 Comparative static analysis

In this section, we ask how the world income distribution would change if all exporters were granted identical market access so that trade costs were bilaterally symmetric. First, we show that Waugh's (2010) results were partly driven by the aforementioned problems with estimation and calibration. Second, we demonstrate that Waugh's counterfactual experiments largely contradict one of the key assumptions of the model and cannot answer the question at hand. Then, we conduct two novel counterfactual experiments and show that a complete elimination of asymmetries would have a negligible impact on the world income distribution.

The central counterfactual exercise in Waugh (2010) boils down to setting counterfactual total trade costs to

$$t_{ij}^c = \min(\hat{t}_{ij}, \hat{t}_{ji}). \tag{5.1}$$

Waugh (2010) claims that this allows one to quantify the effect of an elimination of export market access asymmetries on world income inequality.

We set $t_{ij}^c = \min(\hat{t}_{ij}, \hat{t}_{ji})$ for Model A and Model B and compare the response of $\log y_i$ and y_{90}/y_{10} implied by these models to Waugh's (2010) results. The results are reported in Table 3. We report two results for Model B. They reflect two different assumptions on the response of B_i to changes in trade costs. First, we assume that trade imbalances do not respond to a reduction in trade costs - $B_i^c = B_i$. We view this as a lower bound and denote it by Model B(L). Second, we assume that trade imbalances are elastic to the level of trade costs and $B_i^c = 0$ for all i. We consider this as an upper bound and denote this model by Model B(U). We discuss why these two extremes constitute lower and upper bounds in the following section.

Table 3: World Income Distribution and Trade Costs

	$\log y_i$			y_{90}/y_{10}		
Model	Benchmark	$\min(t_{ij}, t_{ji})$	**change in %**	Benchmark	$\min(t_{ij}, t_{ji})$	**change in %**
Waugh (2010)	1.30	1.05	$-$**24**%	25.7	17.3	$-$**49**%
Model A	1.39	1.22	$-$**14**%	20.9	17.92	$-$**17**%
Model B(U)	1.37	1.23	$-$**11**%	21.3	18.3	$-$**17**%
Model B(L)	1.37	1.24	$-$**11**%	21.3	18.4	$-$**16**%

Table 3 reveals that Waugh's (2010) claims are largely driven by the aforementioned problems in the model estimation and calibration. Waugh (2010) overestimates the impact of the changes in trade costs on international income differences by a factor of two for $\log y_i$ and by a factor of three for y_{90}/y_{10}.

Yet, we argue that setting $t^c_{ij} = \min(\hat{t}_{ij}, \hat{t}_{ji})$ is only vaguely related to a reduction of asymmetric exporter-specific trade costs and the results in Table 3 should be interpreted with caution. Notice that this counterfactual experiment entails an extremely large exogenous reduction in *all* trade costs, not only asymmetric exporter-specific ones. To see this, consider the following. Recall that total trade costs from (3.2) are

$$\hat{t} = \exp\left(-\theta\hat{\xi}_j - \theta\hat{\tau}_{ij}\right), \tag{5.2}$$

where $\hat{\xi}_j$ is the exporter-specific asymmetric component and $\hat{\tau}_{ij}$ is the parameterized bilateral symmetric component of trade costs. However, $\hat{t}^c_{ij} = \min(\hat{t}_{ij}, \hat{t}_{ji})$ in the context of the theoretical model and at constant $\hat{\tau}^c_{ij} = \hat{\tau}_{ij}$ (distance, etc., is held fixed) requires that the estimated exporter fixed effect $\hat{\xi}_j$ – a unilateral, single-indexed variable – must take an ij-(double) index in counterfactual equilibrium to enable $\hat{t}^c_{ij} = \min(\hat{t}_{ij}, \hat{t}_{ji})$. Hence, at constant $\hat{\tau}_{ij}$ as assumed in Waugh (2010) the counterfactual equilibrium is inconsistent with the specifications in (3.2) and (5.2).

Notice that one key identification assumption in Waugh (2010) is that ξ_j must be exporter-specific and cannot vary across importers. Let us fix $\hat{\tau}^c_{ij} = \hat{\tau}_{ij}$ and decompose $\hat{t}^c_{ij}$ as follows:

$$\hat{t}^c_{ij} = \min(\hat{t}_{ij}, \hat{t}_{ji}) = \exp\left(-\theta\tau_{ij} - \theta\max(\hat{\xi}^c_i, \hat{\xi}^c_j)\right). \tag{5.3}$$

Hence, we should be able to recover $\hat{\xi}^c_j$ as follows:

$$\hat{\xi}^c_j = \max(\hat{\xi}_i, \hat{\xi}_j) = -\frac{1}{\theta}\ln(\hat{t}^c_{ij}) - \hat{\tau}_{ij} \tag{5.4}$$

In order for $\hat{\xi}^c_j$ to satisfy (5.2) – a key identification assumption – it cannot vary across importers and, hence, it must be the case that

$$\hat{\xi}^c_j = \max(\hat{\xi}_k, \hat{\xi}_j) \text{ for all } k \tag{5.5}$$

Notice that (5.5) can hold if $\hat{\xi}_j = \max(\hat{\xi}_k)$ or $\hat{\xi}_j = \hat{\xi}_k$ for all k. In either case, the variance of $\hat{\xi}_j$ would have to be zero. Certainly, this is not the case as can be seen from Waugh's (2010) estimates.

To give an example, we illustrate in Figure 3 to which extent Zimbabwe's counterfactual $\hat{\xi}^c_{ZWE}$

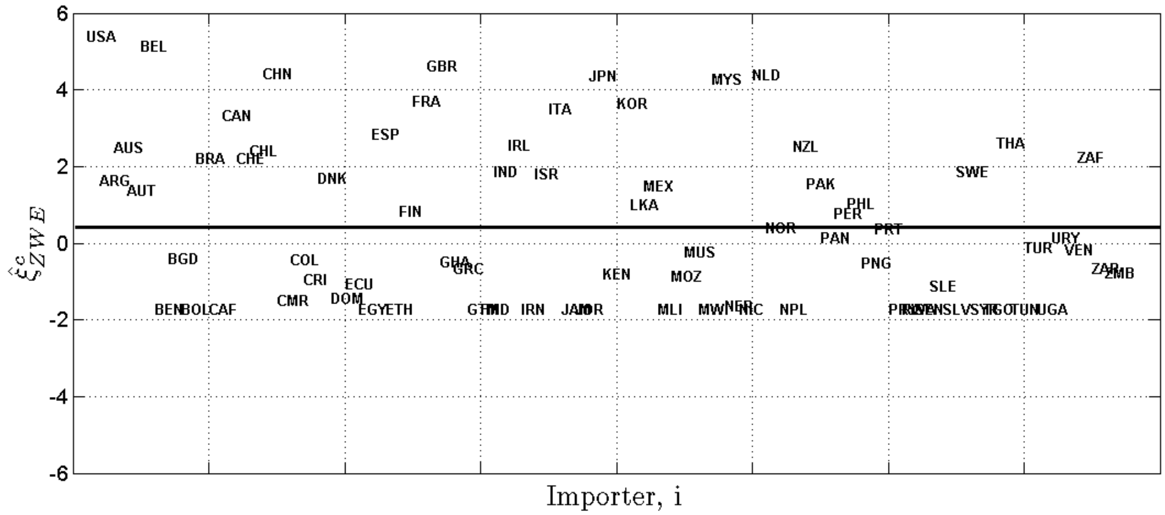

Figure 3: $\hat{\xi}^c_{ZWE}$ across importers i

has to vary across importers to accommodate (5.3). Obviously, $\hat{\xi}^c_{ZWE}$ varies tremendously across importers i which is inconsistent with the identification assumption of the model. In the figure, we plot the average $\hat{\xi}^c_{ZWE} = N^{-1}\sum_{i=1}^{N}\hat{\xi}^c_{ZWE,i} = 0.45$ as a solid horizontal line. The standard deviation of $\hat{\xi}^c_{ZWE,i}$ amounts to 2.14. A variance decomposition of $\hat{\xi}_j$ derived as in (5.4) versus $\hat{\xi}^c_j$ for all exporting countries in the data is given in the upper bloc of Table 4.

Table 4: Variance Decomposition of $\hat{\xi}_j$, $\hat{\xi}^c_j$, $\hat{t}_{ij}$, and $\hat{t}^c_{ij}$

Variance component	$\hat{\xi}_j$	$\hat{\xi}^c_j$
	Sum of Squares	**Sum of Squares**
Exporter-specific	41851.63	11723.08
Residual (Pair-specific	0	16777.78
Total	41851.63	28500.86
Adj. R^2	1	0.41
Observations (Pairs)	5,852	5,852
	$\hat{t}_{ij}$	$\hat{t}^c_{ij}$
	Sum of Squares	**Sum of Squares**
Exporter-specific	33016.86	4877.70
Residual (Pair-specific	6359.02	10873.40
Total	39375.88	15751.09
Adj. R^2	0.84	0.30
Observations (Pairs)	5,852	5,852

By construction, all of the variation of $\hat{\xi}_j$ is exporter-specific in the outset. Yet, in counterfactual equilibrium $\hat{\xi}^c_j$ is not exporter-specific and only about 40% of the variance in ξ^c_j is exporter-specific, while the rest is attributed to the variance across country-pairs. Notice that if $\hat{\xi}_j$ had been constructed like that, it could not have been identified! The reason is that, unlike $\hat{\tau}_{ij}$, which is based on observable variables, nothing about $\hat{\xi}_j$ is unobservable. Hence, beyond the problems associated with the lack of imposition of market clearance, Waugh's (2010) counterfactual model is inconsistent with the identifying assumptions that have to be met in order to estimate ξ_j consistently. This also shows in the variance decomposition of $\hat{t}_{ij}$ and $\hat{t}^c_{ij}$ in the lower bloc of Table 4.

We plot total trade costs in the benchmark versus their respective counterfactual values in Figure 4. To facilitate an easy comparison between the benchmark and counterfactual trade costs we plot them in the familiar iceberg form. We also plot benchmark average trade

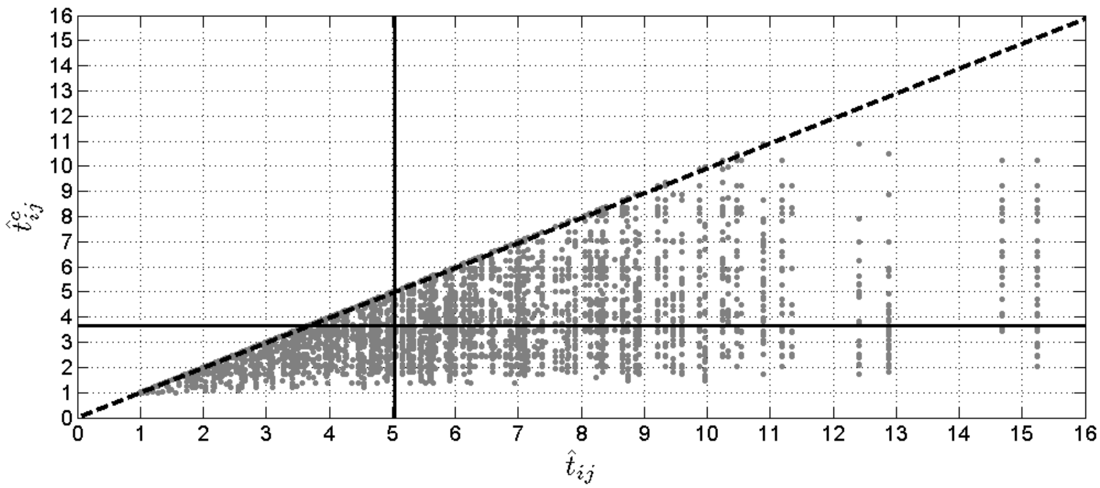

Figure 4: $\hat{t}_{ij}$ VERSUS $\hat{t}^c_{ij}$

costs and their counterfactual counterpart as solid vertical and horizontal lines, respectively. Notice that <u>all</u> bilateral trade costs have unambiguously decreased from the benchmark to the counterfactual equilibrium as is clear from setting $t^c_{ij} = \min(\hat{t}_{ij}, \hat{t}_{ji})$. The average trade cost fell by more than one-third! However, what is interesting to see is that a blunt reduction of all trade costs by one-third would not be enough to generate the effects on trade flows in Waugh (2010). It turns out that it is necessary to completely eliminate *exporter-specific trade costs* and to reduce *all bilateral trade costs* by more than 50% in order to generate the desired impact on bilateral exports. Hence, symmetric exporter-specific market access can in no way induce the effect found in Waugh (2010).

5.1 Asymmetric trade costs and income differences

To assess the importance of the asymmetric component of trade costs on the world income distribution, let us use Models A and B. There is a clear trade off between these two models in terms of matching the data on trade imbalances and assumptions about B_i.

An advantage of Model A is that we do not have to make assumptions about the response of B_i to changes in ξ_i. However, this model cannot match the data on trade imbalances. On the other hand, Model B helps predicting further features of the data (i.e., trade imbalances) at the cost of additional, potentially restrictive assumptions with regard to the response of B_i to exogenous shocks.

To maximize the accuracy of our counterfactual outcomes we employ both models. With regard to Model B and the response of B_i to changes in ξ_j we adopt two alternative assumptions. First, we assume that trade imbalances are completely inelastic to ξ_j. This approach is in the spirit of Caliendo and Parro (2011) and the counterfactual outcome calculated under this assumption is our *lower bound*, referred to as Model B(L) above. To establish the *upper bound* – dubbed Model B(U) above – we assume that under symmetric trade costs trade imbalances converge to zero.[13].

In order to asses the importance of asymmetries in trade costs for international income differences one has to keep *average* trade costs unaffected while reducing the *asymmetric* trade cost component. In our view, there are two types of comparative static experiments that adhere to this requirement. For instance, one could eliminate all differences in exporter-specific unobservable trade costs, ξ_j. Notice that $\hat{\xi}_j$ is centered around zero so that setting counterfactual values $\xi_j^c = 0$ for all exporters j does not affect the average *level* of trade costs in the world economy. Below, we refer to this as *Experiment 1.* Alternatively, one could eliminate trade cost heterogeneity at large by setting all counterfactual bilateral trade costs to the average level, $t_{ij}^c = \bar{t} \equiv [N(N-1)]^{-1} \sum_{j=1}^{N} \sum_{i \neq j} t_{ij}$. By definition, also the latter does not affect the average *level* of trade costs in the world economy. Below, we refer to this as *Experiment 2.* We will contrast these experiments with the outcome of a gradual and symmetric reduction of trade cost *levels after* the asymmetries in the two experiments had been removed. This will illustrate that the stark response of real per-capita income in Waugh (2010) is mainly due to the change in trade cost *levels* rather than their *asymmetry* across exporters.

Experiment 1:

Experiment 1 consists of two steps. In the *first step*, we gradually eliminate the heterogeneity across $\hat{\xi}_j$ *ceteris paribus.* Here, we stepwise reduce unobservable exporter-specific trade costs by defining counterfactual values of $\xi_{\kappa,ij}^1$ for each exporter of the form:

$$\xi_{\kappa,j}^1 = \hat{\xi}_j(1-\kappa) \text{ , where } \kappa \in \{0, 0.01, \ldots, 1\}. \tag{5.6}$$

[13]This assumption may seem arbitrary but there is some evidence that trade cost asymmetries are correlated with imbalances. The correlation between ξ_i and B_i is positive and roughly equals 0.50. In any case, we use this assumption only for establishing the upper bound of the response.

We eliminate all asymmetries in ξ_j at $\xi^1_{1,j} = 0$. As κ increases, the variance of the distribution of fixed exporter-specific trade costs degenerates. In Figure 5, we illustrate the response of two measures of real per-capita income dispersion – the variance, $\text{var}(\ln y_i)$, on the left and the 90-to-10 percentile gap, y_{90}/y_{10}, on the right – to an increase in κ (less heterogeneity in unobservable exporter-specific trade costs, $\hat{\xi}^1_j$) as predicted by Model A (broken line), Model B(L) (solid line with delimiter) and Model B(U) (solid line).

For the lower bound response in Model B we keep all counterfactual $B^1_i = B_i$. For the upper bound we adjust them using κ such that $B^1_{\kappa,i} = B_i(1-\kappa)$.

At the vertical delimiter, $\xi^1_j = 0$ for all j and $\hat{t}^1_{ij}$ is perfectly symmetric *bilaterally* for all ij.[14] It turns out that a complete elimination of the exporter-specific unobservable trade costs displays a negligible impact on international real per-capita income inequality. Depending on the preferred model y_{90}/y_{10} changes by $[-6.18\%, 0.56\%]$ and $\text{var}(\ln y_i)$ by only $[-1.51\%, 2.58\%]$.

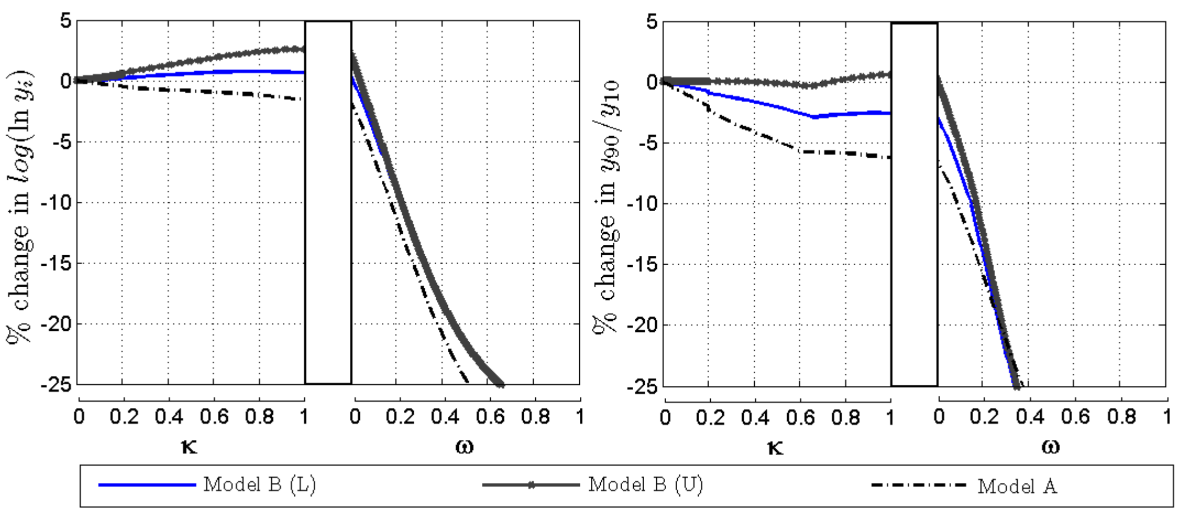

Figure 5: Comparative Static Analysis: Experiment 1

Notice that only Model A predicts negligible negative effect. Predictions of Model B are in the non-negative interval. The reason for this is twofold. First, consistent with Dekle, Eaton and Kortum (2007) a reduction in trade deficits leads to a reduction in earnings in countries with relatively higher trade deficits and vice versa. It is established that poor countries have higher trade deficits. Hence, they will experience a relatively higher decrease in wages. Second, even though poor countries gain in terms of getting symmetric access to foreign markets, they lose in terms of prices of tradables because large exporters such as the

[14] Notice that observable trade costs t^1_{ij} are still heterogenous across countries after setting $\xi^1_j = 0$ for all j due to the differences in distances and adjacency variables. Hence, raising ω in Experiment 1 reduces trade cost asymmetries across exporters *and* trade cost levels in t^c_{ij}, unlike in Experiment 2. Hence, Experiment 2 is cleaner than Experiment 1 in terms of disentangling effects of trade cost *levels* from those of *asymmetries*.

United States now have relatively higher export costs.

Let us denote counterfactual trade costs at $\xi^1_{1,j} = 0$ by t^1_{ij}. In the *second step*, we gradually reduce t^1_{ij} until $\min(t^1_{ij} : i \neq j) = 1$, defining values of $t^1_{\omega,ij}$ for each country pair of the form:

$$t^1_{\omega,ij} = t^1_{ij}(1-\omega) \text{ , where } \omega \in \{0, 0.01, \ldots, 1 : \min(t^1_{\omega,ij} : i \neq j) \geq 1\}. \tag{5.7}$$

Hence, as ω rises, t^1_{ij} approaches unity. The lines to the right of the delimiter in Figure 5 plot the response of $\mathrm{var}(\ln y_i)$ and y_{90}/y_{10} to the changes in ω. Reducing observable trade costs t^1_{ij} gradually leads to relatively bigger responses in the income dispersion than reducing unobservable exporter-specific trade cost heterogeneity. However, we have to reduce t^1_{ij} by more than 50% in order to reduce income inequality by 25% in case of $\mathrm{var}(\ln y_i)$. We have to admit that the second step of Experiment 1 is likely to be irrelevant in terms of policy making. However, the purpose of the step is to show that in order to arrive at Waugh's (2010) results, one would have to reduce general trade costs by an unattainable margin.

Experiment 2

Experiment 2 allows us to disentangle changes in the *asymmetry* versus the *level* of overall (observable *and* unobservable) trade costs. Here, we stepwise reduce trade costs by defining counterfactual values $t^2_{\kappa,ij}$ for each pair of countries:

$$t^2_{\kappa,ij} = \hat{t}_{ij} + \kappa(\bar{t} - \hat{t}_{ij}) \text{ , where } \kappa \in \{0, 0.01, \ldots, 1\}. \tag{5.8}$$

Notice that $t^2_{0,ij}$ corresponds to the benchmark estimate of bilateral trade costs, whereas $t^2_{1,ij} = \bar{t}$.

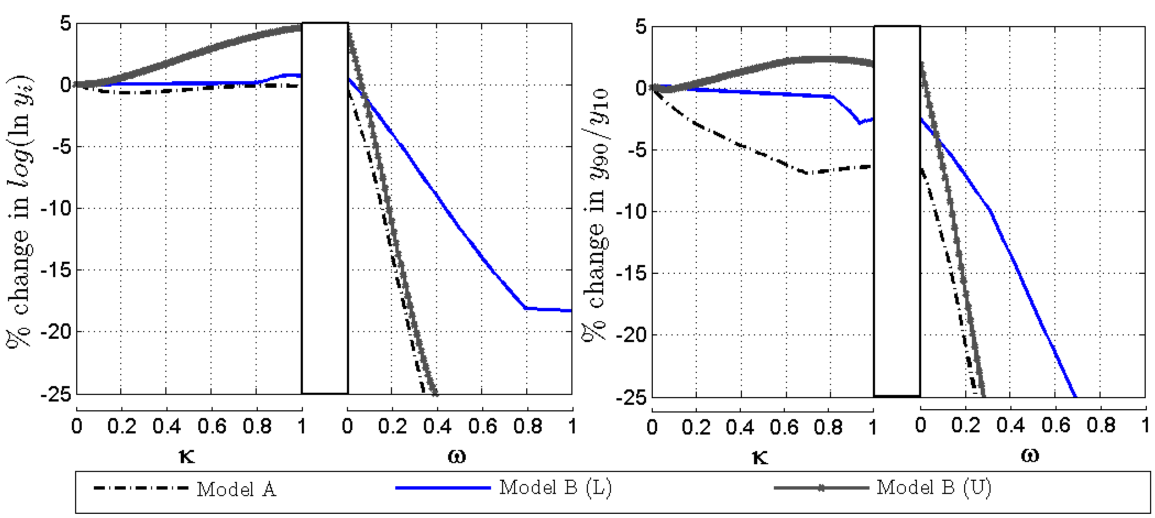

Figure 6: Comparative Static Analysis: Experiment 2

After having eliminated completely trade cost asymmetries at $\kappa = 1$ and $t^2_{1,ij} = \bar{t}$, we can gradually reduce $\bar{t}$ towards unity in order to see how the dispersion in real per-capita incomes responds to changes in trade cost *levels* versus trade cost *asymmetry*:

$$t^2_{\omega,ij} = \bar{t}(1-\omega) \text{ , where } \omega \in \{0, 0.01, \ldots, 1\}. \tag{5.9}$$

As ω approaches unity, $t^2_{\omega,ij}$ converges towards unity so that $t^2_{0,ij} = \bar{t}$ and $t^2_{1,ij} = 1$. Both the responses of $\mathrm{var}(\ln y_i)$ and of y_{90}/y_{10} to changes in κ and ω are summarized in Figure 6 in the same fashion as in Experiment 1. We treat B^c_i exactly as discussed in Experiment 1.

Again, eliminating trade cost asymmetries has little bearing for the dispersion of real per-capita income around the globe. Depending on the model, a complete abolition of trade cost asymmetry leads to only minor change in $\mathrm{var}(\ln y_i)$ – $[-0.15\%, 4.56\%]$, and in y_{90}/y_{10} – $[-6.36\%, 1.91\%]$. Consistent with the results of Experiment 1, these results suggest that in the presence of trade imbalances a mere reduction in trade cost asymmetries may actually increase income inequality around the globe.

The results of Experiment 2 confirm that to achieve a reduction in real per-capita income dispersion by about 25% for $\mathrm{var}(\ln y_i)$ and by 50% for y_{90}/y_{10} as in Waugh (2010), one would have to reduce average trade cost *levels* (beyond exporter trade cost asymmetry) on the globe by a magnitude of at least 50% which of course is not attainable.

6 Conclusion

We propose a way to translate all fundamental assumptions of Eaton-Kortum-type general equilibrium models into a structural estimation of unobservable parameters to avoid creating a wedge between the theoretical model and its structural implementation. We show that ignoring this link may lead to highly biased results.

We revisit Waugh's (2010) analysis to assess the importance of asymmetric trade costs for the world real income distribution and come to starkly different conclusions. We show that a complete abolition of asymmetries in trade costs has close-to-negligible effects on the equalization of real per-capita income across countries. To do that we disentangle trade-imbalances from pure trade-cost-asymmetries and conduct counterfactual experiments that do not violate assumptions of the theoretical model. Our experiments allow us to compare the effect of the reduction in asymmetric versus symmetric general trade costs. We conclude that a reduction in general symmetric trade costs might significantly reduce differences in real per-capita incomes (an old wisdom from the literature), but these reductions would have to be implausibly large to counteract the effects of geographical and cultural barriers to trade.

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Appendix

Data

With one exception, the data underlying this study are the same as the ones in Waugh (2010), available at the AER website. The only other data we use are simply averaged bilateral tariff rates from Mayer, Paillacar, and Zignano (2008) for the average year between 1995-1997 (the reference year of all other data is 1996). The advantage of utilizing such data is that θ can then be identified without using disaggregated price data. We discuss the estimation of θ in the next paragraph.

Choice of parameters

We mostly use the same parameters to calibrate the model as Waugh (2010). However, we use alternative values of θ and γ. The latter allows us to match first and second moments of the distribution of real per-capita income better than Waugh's calibration does. In particular, using a value of $\gamma = 0.60$ instead of $\gamma = 0.75$ as in Waugh helps improving the fit of the variance in real per-capita income, $\text{var}(\ln y_i)$, as well as the 90-to-10 percentile ratio, y_{90}/y_{10}. Our choice of γ seems to be well in line with the data.

For instance, according to the OECD's Structural Analysis Database (STAN) average value added in the non-tradable sector among OECD countries in 2000 amounted to about 55% of total value added which is consistent with a value of $\gamma = 0.55$. Since our sample includes OECD and non-OECD countries, it is plausible that the corresponding value is slightly higher than within the OECD.

Waugh (2010) estimates $(1 - \beta) = 1/3$ from data in the examined country sample. The choice of $(1 - \alpha) = 1/3$ is consistent with conventional values used in the literature and the same as in Waugh.

In order to identify θ, we add a measure of tariff barriers to the specification of t_{ij}. Then, the coefficient on the log of one plus the average bilateral tariff rate, the only observable ad-valorem measure of trade costs in $\ln t_{ij}$, is θ^{-1}. The corresponding estimates under this specification have to be interpreted with caution. Since, tariffs are asymmetric,[15] part of the exporter-specific effects $\hat{\xi}_j$ as estimated in Table 1 now will accrue to the asymmetry in tariffs. The corresponding estimation results for the cum-tariffs model are summarized in Table 5. The coefficient on the tariff measure is -7.36, so that $\hat{\theta} \approx 0.14$. The relatively good

[15] Developed countries consistently faced higher tariffs in manufacturing in 1996 than developing countries did.

Table 5: PPML estimates with tariffs

country	s_i		ξ_i		country	s_i		ξ_i	
United States	1.91	(0.13)	1.79	(0.23)	Republic of Korea	0.39	(0.14)	2.73	(0.27)
Argentina	1.38	(0.08)	0.15	(0.13)	Sri Lanka	1.01	(0.07)	1.43	(0.18)
Australia	0.22	(0.1)	2.65	(0.47)	Mexico	1.15	(0.3)	0.89	(0.3)
Austria	1.28	(0.06)	1.03	(0.18)	Mali	0.15	(0.15)	2.94	(0.28)
Belgium	0.78	(0.21)	2.19	(0.45)	Mozambique	2.1	(0.15)	0.18	(0.15)
Benin	0.38	(0.11)	1.59	(0.3)	Mauritius	0.79	(0.07)	1.08	(0.1)
Bangladesh	0.94	(0.04)	0.68	(0.43)	Malawi	0.91	(0.17)	0.87	(0.22)
Bolivia	0.09	(0.07)	1.13	(0.1)	MalaysiaSingapore	0.22	(0.15)	3.24	(0.2)
Brazil	1.88	(0.06)	0.19	(0.11)	Niger	2.83	(0.34)	0.38	(0.42)
Central African Republic	0.18	(0.12)	3.05	(0.17)	Nicaragua	0.15	(0.08)	2.2	(0.16)
Canada	0.86	(0.26)	0.83	(0.37)	Netherlands	0.03	(0.09)	1.75	(0.2)
Switzerland	1.12	(0.15)	0.49	(0.21)	Norway	0.77	(0.08)	0.52	(0.11)
Chile	0.42	(0.09)	0.5	(0.13)	Nepal	0.77	(0.09)	2.76	(0.18)
ChinaHong Kong	0.39	(0.2)	4.3	(0.32)	New Zealand	0.04	(0.14)	0.68	(0.19)
Cameroon	0.1	(0.09)	0.9	(0.17)	Pakistan	0.27	(0.17)	1.87	(0.54)
Colombia	1.22	(0.1)	0.78	(0.22)	Panama	2.1	(0.55)	2.81	(0.6)
Costa Rica	0.07	(0.12)	0.81	(0.35)	Peru	0.9	(0.06)	0.68	(0.12)
Denmark	0.49	(0.06)	0.08	(0.2)	Philippines	0.17	(0.12)	1.21	(0.15)
Dominican Republic	0.31	(0.23)	0.27	(0.25)	Papua New Guinea	1.86	(0.23)	1.42	(0.28)
Ecuador	0.38	(0.09)	0.86	(0.1)	Portugal	1.03	(0.09)	0.71	(0.25)
Egypt	2.16	(0.07)	0.38	(0.51)	Paraguay	0.04	(0.09)	0.21	(0.12)
Spain	1.93	(0.14)	0.75	(0.24)	Rwanda	0.72	(0.05)	3.1	(0.15)
Ethiopia	0.71	(0.06)	0.19	(0.16)	Senegal	0.07	(0.15)	1.57	(0.28)
Finland	1.17	(0.07)	1.03	(0.12)	Sierra Leone	1.54	(0.19)	2.3	(0.24)
France	0.62	(0.17)	2.25	(0.71)	El Salvador	0.55	(0.1)	0.52	(0.17)
United Kingdom	0.91	(0.09)	1.6	(0.39)	Sweden	0.78	(0.07)	0.23	(0.13)
Ghana	0.92	(0.12)	0.33	(0.16)	Syrian Arab Republic	2.65	(0.06)	3.29	(0.2)
Greece	1.43	(0.08)	1.66	(0.2)	Togo	0.31	(0.11)	2.09	(0.15)
Guatemala	0.07	(0.08)	0.83	(0.21)	Thailand	1.26	(0.13)	1.15	(0.23)
Honduras	0.48	(0.17)	0.59	(0.2)	Tunisia	0.93	(0.1)	0.31	(0.26)
India	0.55	(0.18)	3.22	(0.38)	Turkey	1.9	(0.07)	1.75	(0.15)
Ireland	0.35	(0.13)	0.69	(0.18)	Uganda	0.31	(0.06)	2.17	(0.06)
Iran	2.27	(0.08)	2.97	(0.1)	Uruguay	0.52	(0.09)	1.2	(0.12)
Israel	0.16	(0.09)	0.43	(0.19)	Venezuela	0.18	(0.14)	0.3	(0.17)
Italy	1.68	(0.15)	0.31	(0.29)	South Africa	0.61	(0.09)	3.21	(0.21)
Jamaica	0.5	(0.2)	0.2	(0.22)	Democratic Republic of the Congo	0.47	(0.07)	1.44	(0.09)
Jordan	0.44	(0.06)	0.76	(0.2)	Zambia	0.94	(0.18)	0.17	(0.24)
Japan	1.46	(0.11)	2.29	(0.23)	Zimbabwe	0	(0.13)	0.19	(0.24)
Kenya	0.72	(0.06)	1.19	(0.23)					
$[0, 375)$	4.19	(0.27)							
$[375, 750)$	4.44	(0.16)							
$[750, 1500)$	4.17	(0.13)							
$[1500, 3000)$	4.71	(0.22)							
$[3000, 6000)$	5.72	(0.17)							
$[6000, max)$	6.39	(0.19)							
Adjacency	1.08	(0.16)							
Tariff	**7.36**	**(1.31)**							
Pseudo R^2	0.74								

Notes: $\theta = 0.14$, Standard errors are reported in parentheses and are based on Eicker-White sandwich estimates. The reported Pseudo R^2 corresponds to the correlation between observed and predicted values of the dependent variable.

fit of the proposed calibration relative to Waugh's is robust to reasonable perturbations of θ and γ.