International Trade and Income Differences: Re-examination

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Abstract

In a recent article, Waugh (2010) employs an Eaton-Kortum-type structural model of bilateral trade to show that developing countries face much higher relative export costs than developed countries, and that this asymmetry in trade costs accounts for a significant part of the variance in relative real per-capita incomes across countries. Equalizing the cost of export market access across economies would help equalizing real per-capita incomes around the globe. We reassess this analysis using the same data as Waugh and come to starkly different conclusions. Respecting the assumed general equilibrium constraints and calibrating the model suitably to the original data, we illustrate that reducing exporter trade cost asymmetry alone leads to a reduction in the variance of log real per-capita incomes by not more than 1.5% rather than by almost one-third as in Waugh. From this, we conclude that there is little hope for the poor to close the real income gap significantly from getting symmetric access to world markets in comparison to developed countries.

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

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

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1 Introduction

In a recent article, Waugh (2010) employs a variant of the model of Eaton and Kortum (2002) to examine how exporter cost asymmetries across countries affect international differences in real per-capita income. Two findings from this analysis stand out. 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 almost one-third.

We utilize the same data-set and theoretical model as Waugh but arrive at starkly different results: eliminating exporter trade cost heterogeneity has only little bearing for real per-capita income variability across countries. The difference between the results in this paper in comparison to Waugh (2010) has two main roots.

First, multilateral trade balance is *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. Yet, Waugh's structural empirical model does not respect this constraint so that estimated exporter trade cost asymmetries partly reflect 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.

Second, the counterfactual experiment conducted in Waugh does not systematically reduce the *asymmetry* in exporter-specific trade costs, but it mainly changes the *level* of trade costs. Hence, Waugh's result can not be interpreted as to reflect the consequences of reduced exporter-specific trade cost asymmetries. We illustrate that reducing 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.[1]

The remainder of the paper is organized as follows. The next section briefly outlines the model underlying the structural estimation of bilateral as well as exporter-specific trade costs. Section 3 summarizes the model estimates, while Section 4 documents findings from a counterfactual analysis. Section 5 provides a sensitivity analysis with respect to the assumption of multilaterally balanced trade, and the last section concludes.

[1] That a reduction of trade friction levels would have big consequences for an equalization of per-capita incomes 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

In the model of Eaton and Kortum (2002) as utilized by Waugh (2010) each one of N countries hosts a *continuum* of firms in the tradable sector and *one* representative firm in the non-tradable sector. Firms in the tradable sector of country j face variable (and marginal) costs per efficiency unit of c_{tj} and the firm in the nontradable sector faces variable (and marginal) unit costs of c_{nj}. 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. In that model, country i spends a share $X_{ij} \equiv M_{ij}/Y_{ti}$ on tradable goods from j, where M_{ij} are aggregate nominal imports of i from j and Y_{ti} denotes i's total spending on tradables. Then,

$$X_{ij} = \left(\frac{c_{tj} t_{ij}}{\Psi p_{ti}} \right)^{-\frac{1}{\theta}}, \tag{2.1}$$

where Ψ is a constant and p_{ti} is the price index of tradable goods that will be defined later on. If we normalize (2.1) 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:

$$\frac{X_{ij}}{X_{ii}} = \left(\frac{c_{tj} t_{ij}}{c_{ti} t_{ii}} \right)^{-\frac{1}{\theta}}. \tag{2.2}$$

In estimating (2.2), structural empirical work typically adopts five assumptions: (i) the stochastic counterpart to (2.2) involves either an additive or a multiplicative error term; (ii) trade costs $t_{ij}^{-\frac{1}{\theta}}$ are modeled multiplicatively as $t_{ij}^{-\frac{1}{\theta}} = e^{\xi_j + \sum_{k=1}^{K} \beta_k \tau_{k,ij}}$, where $\tau_{k,ij}$ is the kth observable measure of log trade costs – such as log bilateral distance or an adjacency indicator – and $\beta_k = -\beta_{k0}/\theta$ are corresponding parameters on those observable trade costs while $\xi_j = -\xi_{j0}/\theta$ is a measure of unobservable (net) asymmetric trade costs for exporter j;[2] (iii) there is free intra-national trade with $t_{ii} = 1$; and (iv) trade is multilaterally balanced in the sense of $\sum_{j=1}^{N}(M_{ij} - M_{ji}) = 0$ for all i which corresponds to[3]:

$$Y_{ti} - \sum_{j=1}^{N} (Y_{tj} X_{ji}) = 0. \tag{2.3}$$

[2]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).

[3]Dekle, Eaton, and Kortum (2007) dispense with this assumption by imposing trade balance up to a country-specific constant. However, we do not follow their approach here since further assumptions would have to be made as to the response of trade imbalances to fundamentals in counterfactual equilibrium. We will shed light on such a model variant in Section 5.

When collecting terms in (2.2) the stochastic model can the be expressed as

$$\frac{X_{ij}}{X_{ii}} = e^{s_j - s_i + \xi_j + \sum_{k=1}^{K} \beta_k \tau_{k,ij}} + u_{ij} \quad \text{or} \quad \ln \frac{X_{ij}}{X_{ii}} = s_j - s_i + \xi_j + \sum_{k=1}^{K} \beta_k \tau_{k,ij} + v_{ij}, \tag{2.4}$$

where s_j and s_i are country-specific catch-all variables associated with prices and technology parameters of countries j and i, respectively. By design, $s_j = s_i$ whenever $i = j$. u_{ij} and v_{ij} are stochastic terms for the models in levels and logs, respectively.[4] As long as multilateral trade balance is imposed, $\xi_j \neq 0$ at least for some exporters j implies that unobservable country-specific market access (or trade) costs are asymmetric and ξ_j is a complete measure thereof.

With regard to assumption (i) above, we follow Santos Silva and Tenreyro (2006) by mainly using Poisson pseudo-maximum likelihood to estimate a modified version of (2.2). This has the two advantages of being applicable with log-additive and level-additive error models and of including observations where $M_{ij} = 0$. With regard to assumption (ii), we broadly follow Eaton and Kortum (2002) and Waugh (2010) by assuming

$$-\frac{1}{\theta} \ln t_{ij} \quad = \quad \beta_1 adjacency_{ij} + \beta_2 \ln(distance_{1,ij}) + \ldots + \beta_7 \ln(distance_{6,ij}), \tag{2.5}$$

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. The formulation of trade costs in (2.5) is in some regards more flexible than one that is based on the log distance between countries i and j. With respect to assumption (iii), we normalize intra-national trade costs to unity ($t_{ii} = 1$) as is common in the literature. Concerning assumption (iv), we impose multilateral trade balance as in (2.3) throughout which is not the case in Waugh (2010).

With one exception, the data underlying this study are the same as the ones in Waugh (2010) so that we may economize on the description of their sources. 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.[5]

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

[5] Highly disaggregated price data are necessary for accurate estimation of θ (e.g. see Simonovska and Waugh, 2011). One can estimate θ directly by including the log of one plus the average bilateral tariff rate as a regressor in (2.5). The corresponding parameter is an estimate of $-1/\theta$. Notice that none of the results depends crucially on this deviation from Waugh's strategy.

3 Estimation

3.1 Multilateral trade imbalance

Multilateral trade balance is one of the central assumptions in both Eaton and Kortum (2002) and Waugh (2010) and elemental for ξ_j to measure (unobservable) exporter-specific trade costs. Empirically, trade is unbalanced multilaterally for most countries. Moreover, it turns out that multilateral trade imbalances are systematically related to real per-capita levels across countries. To see this, let us define

$$B_i = \frac{\sum_{j=1}^{N}(M_{ij} - M_{ji})}{Y_{ti}} \tag{3.1}$$

as a measure of normalized (by apparent consumption) multilateral trade deficit and plot it against real per-capita income ($RGDPPC$) in Waugh's data in Figure 1.

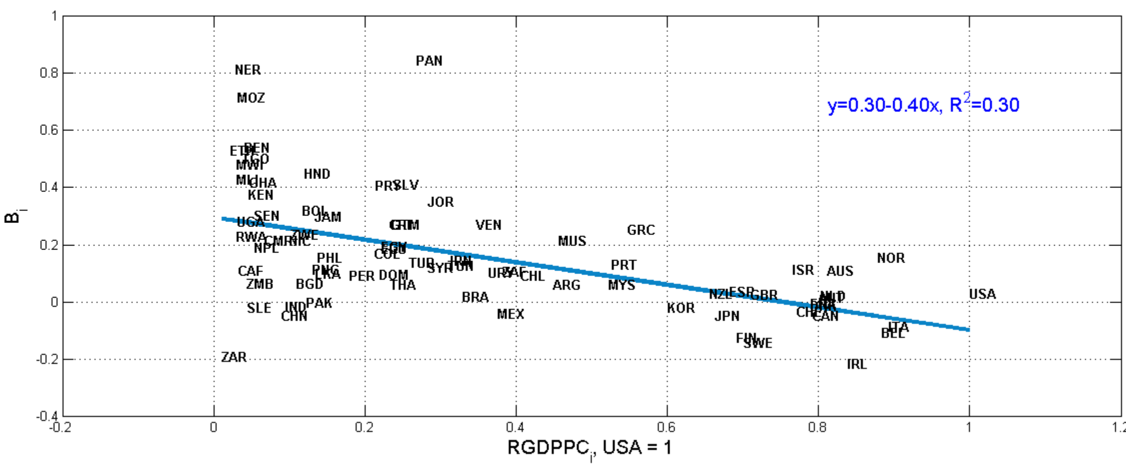

Figure 1: Normalized Trade Imbalance (B_i) and Income per Capita ($RGDPPC_i$)

The data suggest two things. First, trade between rich and poor countries is highly unbalanced. Second, B_i is systematically related to real per-capita income. While this does not necessarily invalidate an analysis along Waugh's lines, it requires one of two strategies: imposing multilateral trade balance onto the data so as to disentangle ξ_j as a measure of pure exporter cost asymmetry from multilateral trade balance which then becomes part of the disturbance term in (2.4); or modeling trade imbalances explicitly.[6] Unless multilateral

[6] We follow the first strategy here but emphasize that Dekle, Eaton, and Kortum (2007) offer a model to explicitly account for trade imbalances in estimation. We follow their approach only in Section 5 in the interest of keeping the analysis as close as possible to Waugh's (2010). In any case, when accounting for trade imbalances with a comparative static analysis in mind as we do, one has to put additional assumptions about the response of trade imbalances to shocks in, say, trade costs.

trade balance is imposed properly on the data or respected otherwise, the impact of trade cost asymmetry through ξ_j on the variance in real per-capita income may be confounded by effects which accrue to trade imbalances rather than trade cost asymmetry per se.

3.2 Waugh's approach

The parameter estimates obtained by Waugh (2010) are not generally consistent with multilateral trade balance, in contrast to the adopted assumptions. To see this, note that Waugh estimates (2.4) subject to[7]:

$$\sum_{j=1}^{N} s_j = 0 \quad \text{and} \quad \sum_{j=1}^{N} \xi_j = 0 \tag{3.2}$$

while disregarding the multilateral trade balance constraint in (2.3).[8] We can visualize this problem by means of Figure 2. There, we plot the normalized data on trade imbalances B_j against Waugh's original estimates of ξ_j in the left panel.[9] In the right panel, we plot the predicted B_j from Waugh's log-linear OLS model rather than data on B_j against the corresponding predictions of ξ_j. The relationship between these variables as measured in terms of the regression coefficient in the figure should be zero, at least in the right panel, which is not the case.

Figure 2 suggests that estimation of (2.2) without respecting the multilateral trade balance conditions obtains estimates of ξ_j which should not be interpreted as measures of exporter trade cost asymmetry.

3.3 Imposing multilateral trade balance

When estimating the model in (2.2) with a specification of observable trade costs as in (2.5) by Poisson pseudo-maximum likelihood subject to (2.3) and the latter constraint in (3.2), one obtains results as summarized in Table 1. In particular, we report estimates of the exporter-specific asymmetric (log) trade costs, $\hat{\xi}_j$, the percentage-change in trade through

[7] Waugh(2010) imposes those constraints implicitly by first coding s_i, s_j, and ξ_j as deviations from the United State as $s_i - s_{USA}$, $s_j - s_{USA}$, $\xi_i - \xi_{USA}$, getting estimates for all $i \neq USA$ and then recovering s_{USA} and ξ_{USA} from (3.2).

[8] By centering the country fixed effects s_j and ξ_j around zero, the set of constraints in (3.2) ensures that a country's average log bilateral exports equals its average log bilateral imports, which is not the same as multilateral trade balance.

[9] He reports OLS estimates of the log-linear version of normalized bilateral trade flows in (2.4) which eliminates observations with zero bilateral imports. The correlation is only slightly lower for the respective estimates of Poisson pseudo-maximum likelihood where zero observations on the dependent variable are not dropped. Details are in the Appendix.

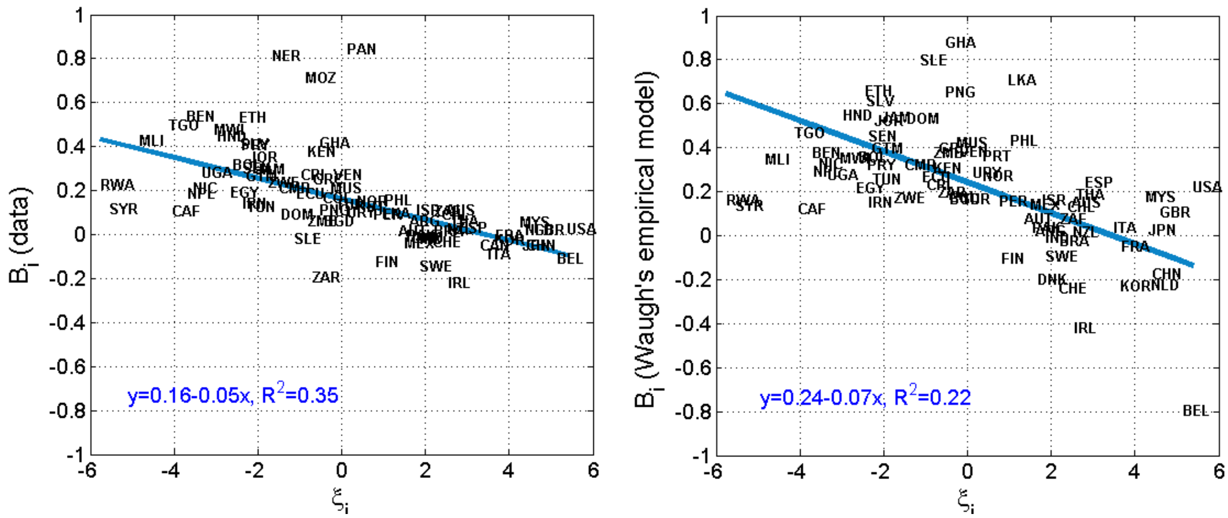

Figure 2: Exporter Fixed Effects (Waugh, 2010) and Trade Imbalances

asymmetric exporter-specific costs alone, $100 \times (e^{\hat{\xi}_{0j}} - 1)$, the fixed effects $\hat{s}_j$, and a relative productivity measure $\left(\frac{\hat{\lambda}_{USA}}{\hat{\lambda}_i}\right)^{\hat{\theta}}$ which we will utilize and discuss only in the next section but report in Table 1 to economize on space. Notice that the latter estimate and $100 \times (e^{\hat{\xi}_{0j}} - 1)$ with $\hat{\xi}_{0j} = -\hat{\theta}\hat{\xi}_j$ employ an estimate $\hat{\theta}$ which is obtained and discussed in the next section and the Appendix. Moreover, the table summarizes parameters on the observable trade costs $\tau_{k,ij}$ in the bloc at the bottom.

Most importantly, there are large differences in the estimated exporter effects $\hat{\xi}_j$ and $\hat{\xi}_{0j}$ between Table 1 and Waugh's (2010) findings. For example, Waugh suggests that $100 \times (e^{\hat{\xi}_{0j}} - 1)$ – i.e., the observable trade cost amplification factor of unobservable exporter-specific trade costs – of Benin and Rwanda amounts to 96% and 184%, respectively. The corresponding numbers for these two countries in Table 1 are only 22% and 67%.

There are two central questions at this point. First, how are the obtained estimates of ξ_j correlated with real per-capita incomes relative to Waugh's? Second, how would real per capita incomes change if we completely abolished trade cost asymmetries in comparison to Waugh's analysis? The subsequent analysis is devoted to answering these questions.

4 Trade cost asymmetries and real per-capita income differences

For an assessment of exporter trade cost asymmetry on outcome, we need two ingredients: estimates of t_{ij} in general and of ξ_{0j} in specific; predictions of real per-capita income on the

Table 1: PPML estimates

Country	$\hat{\xi}_j$		$100 \times (e^{\hat{\xi}_{0j}} - 1)$	$\hat{s}_j$		$\left(\frac{\hat{\lambda}_{USA}}{\hat{\lambda}_i}\right)^{\hat{\theta}}$
United States	2.61	(0.27)	-30.62	1.87	(0.12)	1.00
Argentina	-0.06	(0.14)	0.84	1.37	(0.08)	1.83
Australia	3.08	(0.44)	-35.01	-0.02	(0.09)	1.84
Austria	-0.29	(0.15)	4.12	1.28	(0.05)	1.16
Belgium	2.91	(0.44)	-33.44	-0.75	(0.20)	1.42
Benin	-1.42	(0.22)	22.02	-0.45	(0.11)	13.22
Bangladesh	-1.44	(0.06)	22.36	0.97	(0.05)	3.49
Bolivia	-0.94	(0.10)	14.11	-0.03	(0.08)	4.72
Brazil	0.21	(0.11)	-2.93	1.87	(0.06)	1.56
Central African Republic	-3.45	(0.14)	62.00	0.11	(0.11)	4.63
Canada	1.45	(0.32)	-18.32	0.92	(0.26)	1.26
Switzerland	0.11	(0.21)	-1.48	1.32	(0.15)	0.95
Chile	0.66	(0.12)	-8.77	0.51	(0.09)	1.92
China-Hong Kong	3.12	(0.24)	-35.35	0.54	(0.19)	2.62
Cameroon	-1.38	(0.14)	21.24	-0.08	(0.09)	5.39
Colombia	-0.81	(0.26)	12.01	1.26	(0.11)	3.01
Costa Rica	-0.47	(0.34)	6.80	0.04	(0.12)	3.12
Denmark	0.83	(0.16)	-10.92	0.50	(0.05)	1.30
Dominican Republic	0.16	(0.27)	-2.15	-0.28	(0.23)	2.68
Ecuador	-0.84	(0.12)	12.45	0.38	(0.08)	3.27
Egypt	-2.32	(0.16)	38.40	2.09	(0.06)	3.20
Spain	0.04	(0.21)	-0.49	1.92	(0.13)	1.03
Ethiopia	-0.36	(0.11)	5.14	-0.83	(0.06)	13.86
Finland	-0.25	(0.07)	3.54	1.17	(0.06)	0.95
France	3.03	(0.72)	-34.54	0.61	(0.17)	1.21
United Kingdom	2.40	(0.38)	-28.53	0.89	(0.09)	1.26
Ghana	0.12	(0.16)	-1.62	-0.89	(0.13)	6.09
Greece	-0.87	(0.14)	13.03	1.43	(0.07)	1.95
Guatemala	-0.64	(0.20)	9.42	0.12	(0.07)	4.66
Honduras	-0.37	(0.20)	5.33	-0.48	(0.17)	4.96
India	1.50	(0.21)	-18.95	0.70	(0.18)	3.23
Ireland	1.41	(0.14)	-17.91	-0.33	(0.13)	1.22
Iran	-2.86	(0.10)	49.33	2.41	(0.08)	3.12
Israel	0.91	(0.18)	-12.01	0.25	(0.09)	1.56
Italy	1.09	(0.27)	-14.14	1.68	(0.15)	0.87
Jamaica	-0.25	(0.24)	3.62	-0.48	(0.21)	3.35
Jordan	-1.53	(0.15)	23.95	0.47	(0.05)	3.25
Japan	2.84	(0.25)	-32.79	1.58	(0.10)	0.81
Kenya	0.29	(0.12)	-4.04	-0.76	(0.06)	9.05
Republic of Korea	2.70	(0.28)	-31.52	0.64	(0.16)	1.15
Sri Lanka	0.77	(0.08)	-10.24	-0.95	(0.06)	4.54
Mexico	0.60	(0.37)	-8.04	1.29	(0.35)	1.59
Mali	-2.47	(0.25)	41.26	0.05	(0.14)	11.59
Mozambique	0.41	(0.14)	-5.53	-2.20	(0.14)	13.90
Mauritius	0.34	(0.08)	-4.67	-0.86	(0.07)	2.75
Malawi	-0.95	(0.23)	14.18	-1.01	(0.17)	12.98
Malaysia-Singapore	3.06	(0.15)	-34.89	0.09	(0.13)	1.26
Niger	-0.23	(0.40)	3.21	-2.92	(0.34)	22.16
Nicaragua	-1.87	(0.12)	30.00	0.17	(0.08)	3.77
Netherlands	2.51	(0.17)	-29.65	0.03	(0.09)	1.39
Norway	0.08	(0.08)	-1.07	0.96	(0.06)	1.29
Nepal	-3.49	(0.10)	62.93	0.80	(0.10)	5.93
New Zealand	1.31	(0.17)	-16.71	0.02	(0.11)	1.60
Pakistan	-0.25	(0.33)	3.49	0.38	(0.17)	3.53
Panama	3.10	(0.59)	-35.20	-2.08	(0.53)	7.34
Peru	-0.77	(0.13)	11.36	0.96	(0.07)	2.62
Philippines	1.13	(0.16)	-14.62	0.26	(0.12)	2.85
Papua New Guinea	1.04	(0.28)	-13.57	-1.93	(0.23)	5.00
Portugal	0.07	(0.21)	-0.94	1.03	(0.08)	1.46
Paraguay	0.13	(0.11)	-1.76	-0.21	(0.10)	5.01
Rwanda	-3.68	(0.06)	67.37	0.61	(0.05)	11.93
Senegal	-1.48	(0.22)	23.09	-0.06	(0.16)	4.67
Sierra Leone	-1.47	(0.25)	22.84	-1.64	(0.18)	6.54
El Salvador	-0.23	(0.16)	3.21	-0.53	(0.10)	5.71
Sweden	1.04	(0.08)	-13.51	0.77	(0.06)	1.03
Syrian Arab Republic	-3.98	(0.12)	74.65	2.66	(0.04)	2.69
Togo	-1.90	(0.11)	30.49	-0.39	(0.11)	8.84
Thailand	0.07	(0.13)	-0.95	1.44	(0.13)	1.69
Tunisia	-1.41	(0.15)	21.75	0.99	(0.10)	1.85
Turkey	-1.53	(0.17)	23.88	2.09	(0.07)	2.01
Uganda	-1.98	(0.06)	32.02	0.23	(0.06)	6.96
Uruguay	-0.65	(0.11)	9.57	0.33	(0.07)	2.25
Venezuela	0.30	(0.19)	-4.06	0.19	(0.13)	3.93
South Africa	3.20	(0.20)	-36.14	-0.53	(0.08)	2.88
Democratic Republic of the Congo	-0.79	(0.09)	11.65	-0.53	(0.07)	5.43
Zambia	-0.38	(0.23)	5.53	-1.05	(0.18)	4.85
Zimbabwe	-0.53	(0.25)	7.66	-0.03	(0.12)	4.71
$\ln(distance)$:						
$[\ln min, \ln 375)$	-4.96	(0.25)	100.12	-	-	-
$[\ln 375, \ln 750)$	-5.22	(0.13)	107.76	-	-	-
$[\ln 750, \ln 1500)$	-5.00	(0.07)	101.25	-	-	-
$[\ln 1500, \ln 3000)$	-5.59	(0.12)	118.71	-	-	-
$[\ln 3000, \ln 6000)$	-6.70	(0.13)	155.32	-	-	-
$[\ln 6000, \ln max)$	-7.38	(0.14)	181.10	-	-	-
adjacency	1.10	(0.15)	-14.26	-	-	-
Pseudo R^2	0.75					

Notes: We utilize $\hat{\theta} \approx 0.14$ as estimated in the Appendix rather than one of $\theta = 0.18$ as in Waugh (2010) to estimate $\hat{\xi}_{0j} = -\hat{\theta}\hat{\xi}_j$ and to calculate $\left(\frac{\hat{\lambda}_{USA}}{\hat{\lambda}_i}\right)^{\hat{\theta}}$ from the estimates of $\hat{s}_j$. The latter will be relevant only for the counterfactual analysis and is discussed, there. Standard errors are reported in parenthesis 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.

basis of the general equilibrium model in the benchmark situation of estimated trade costs and possibly some counterfactual values thereof. The latter requires more details on the model structure than provided above.

4.1 Model structure and calibration

As said before, we use data on bilateral tariffs to estimate the parameter θ in a separate model in the Appendix. The corresponding regression obtains an estimate $\hat{\theta} \approx 0.14$ which has already been used in Table 1. That estimate is lower than the one of $\theta = 0.18$ used in Waugh (2010) and closer to the benchmark values used in Eaton and Kortum (2002) and Alvarez and Lucas (2006). Yet, the results of interest here do not crucially depend on that parameter.

Each country is endowed with two internationally immobile factors: aggregate labor force L_i and aggregate capital K_i. These factors and the bundle of tradable intermediate goods are employed in production of any firm in any sector. All factors are mobile nationally across sectors so that their costs – w_i for labor, r_i for capital, and p_{ti} for the bundle of tradable intermediates – do not bear a sector index. Denote the average variable (and marginal) costs of producing one unit of output of differentiated tradables and homogeneous non-tradables in country i by:

$$c_{ti} = \lambda_i^{-\theta}(w_i^{1-\alpha} r_i^{\alpha})^{\beta} p_{ti}^{1-\beta} \quad \text{and} \quad c_{ni} = (w_i^{1-\alpha} r_i^{\alpha})^{\gamma} p_{ti}^{1-\gamma}, \tag{4.1}$$

where λ_i is a parameter determining the extreme-value (exponential) distribution about total factor productivity $z_i(g)$ across varieties $g \in [0, 1]$ of the continuum of tradable goods in i.

At given K_i/L_i we can find r_i as a function of w_i from the first-order conditions of the final good producers:

$$r_i = w_i \frac{\alpha}{1-\alpha} \frac{K_i}{L_i}. \tag{4.2}$$

Domestic consumption of traded goods requires prior aggregation via a Spence-Dixit-Stiglitz technology so that p_{ti} is a constant-elasticity-of-substitution price index. Using properties of the exponential distribution we obtain

$$p_{ti} = \Omega \sum_j \left(\lambda_i \left((w_j^{1-\alpha} r_j^{\alpha})^{\beta} p_{tj}^{1-\beta} t_{ij} \right)^{-\frac{1}{\theta}} \right)^{-\theta}, \tag{4.3}$$

where Ω is a constant. Notice that (4.3) can be solved implicitly for p_{ti} if λ_i, w_j, α, β, t_{ij}, and θ are known.[10]

Notice that the estimates of s_j in Table 1 reflect $\theta^{-1}\ln\hat{c}_{tj}$. Hence, we can recover p_{ti} using e^{s_j} and t_{ij} from (4.3). With the estimates of p_{ti}, t_{ij} and s_j at hand we can predict trade shares X_{ij} from (2.1). Substituting expressions in (2.3), we can rewrite the multilateral trade balance condition as

$$L_i w_i \sum_i X_{ij} = \sum_j L_j w_j X_{ji} \text{ or} \tag{4.4}$$

$$L_i w_i \sum_i e^{s_j} \left(\frac{t_{ij}}{p_{ti}}\right)^{-\frac{1}{\theta}} = \sum_j L_j w_j e^{s_i} \left(\frac{t_{ji}}{p_{tj}}\right)^{-\frac{1}{\theta}} \tag{4.5}$$

and solve it for w_i, where we choose w_{USA} as the numéraire. Finally, we recover λ_i from (4.1) using the estimates of s_i, w_i, r_i, and p_{ti}.

The central difference between our model solution and calibration and the ones of Waugh (2010) lies in how the multilateral trade balance condition is employed. Waugh (2010) uses *data* on trade shares X_{ij} in (4.4) to solve for w_i. This is inconsistent with general equilibrium given the estimates for two reasons. First, as illustrated above, *data* on X_{ij} do not support multilaterally balanced trade. Second, w_i is then a function of the stochastic term (i.e., the discrepancy between *data* and model predictions of X_{ij}) which does not have an interpretation in the rest of the model.[11] In contrast to Waugh, we use estimates of s_i and t_{ij} as in (2.4) to pin down w_i from (4.5) as the model suggests.

In Figure 3, we illustrate that Waugh's (2010) approach and the one taken in this paper differ starkly in their relative performance to predict moments of the distribution of real per-capita income.

Quite obviously, the calibration used in this paper matches the two first moments of the distribution of real per-capita income much more closely than Waugh's calibration does. In comparison to our approach, Waugh's (2010) systematically under-predicts real per-capita income levels of poor countries and over-predicts income levels of rich countries as shown in Figure 3. When employing Waugh's estimates in solving (4.5) for w_i, the benchmark model fits the data even considerably worse than when using (4.4) instead of (4.5). This is documented in more detail in the Appendix.

[10] As in Waugh (2010), we set $\alpha = \beta = 1/3$. Further details on our choice of parameters can be found in the Appendix.

[11] An additional problem with Waugh's benchmark calibration is the following. His OLS estimates of the coefficients in (2.5) predict some t_{ij} which are inconsistent with zero trade flows and some which are inconsistent with the assumption of $\min(t_{ij}) \geq 1$. Waugh (2010) then employs alternative values t^a_{ij} instead of estimates of t_{ij} from (2.5) such that: (i) $t^a_{ij} = 100$ whenever $X_{ij} = 0$ and (ii) $t^a_{ij} = 1$ whenever $t_{ij} < 1$.

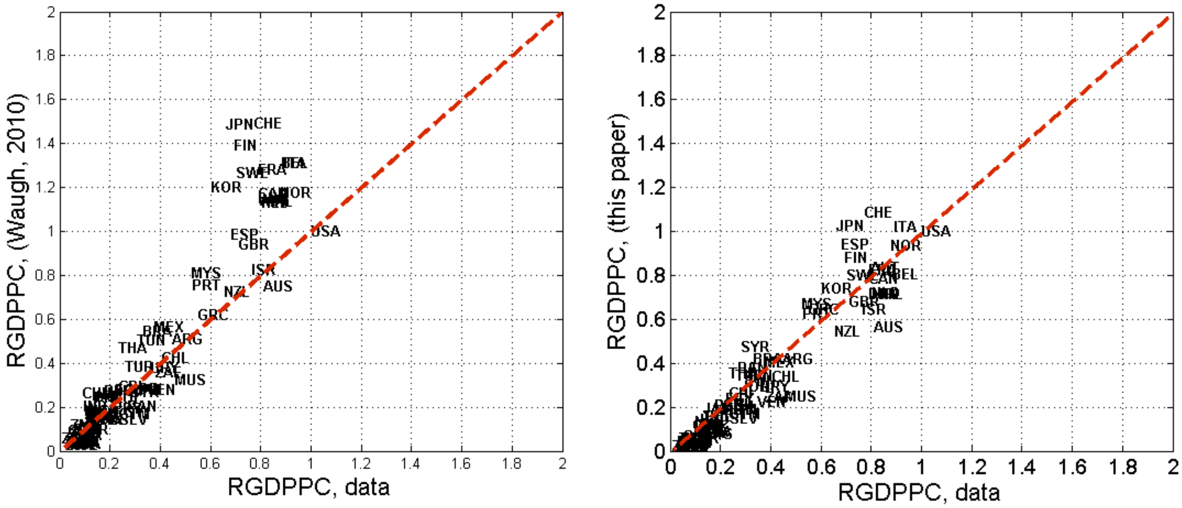

Figure 3: Fit of Calibration: Waugh (2010) versus This Paper

4.2 Correlation of $e^{-\hat{\xi}_{0j}}$ with real per-capita incomes

We provide an assessment of the correlation of the estimates $e^{\hat{\xi}_{0j}}$ with real per-capita incomes in Figure 4. The figure contains two panels. The one on the left employs *data* on real per-capita income and *estimates* of $e^{\hat{\xi}_{0j}}$ which are based on log-linear OLS model estimates of (2.4) subject to (3.2) as in Waugh.[12] This model does not impose trade balance, and the figure exactly reproduces Waugh's result in his Figure 2.[13] According to the left panel of Figure 4, marginally higher (exponentiated) asymmetric exporter trade costs translate almost one-for-one into marginally lower real per-capita income. The overall R^2 of the linear regression in the left panel is as high as 0.43.

Yet, plotting *data* on real per-capita incomes against *estimates* of (exponentiated) asymmetric exporter trade costs is not quite appropriate. Then, part of the correlation may accrue to confounding factors which are beyond the theoretical and estimated model. We should rather correlate *predictions* of per-capita income according to the employed structural model with *estimates* $e^{\hat{\xi}_{0j}}$.[14] This is done in the right panel of Figure 4.[15] Notice that the relationship at stake is dramatically weaker in that panel: the slope coefficient drops to almost one-third of the original value and the R^2 drops by about one-half.

The difference between our result and the one in Waugh accrues to three issues: (i) the

[12] The results are also true for PPML estimates. Details can be found in the Appendix.

[13] Recall that Waugh uses a value of $\theta = 0.18$ to calculate $\hat{\xi}_{0j}$ and calibrate the model.

[14] We can express nominal GDP of country i as: $GDP_i = L_i w_i + K_i r_i$. Define $k_i \equiv K_i/L_i$ and normalize GDP_i by L_i and the final good price index p_{ni} to obtain real per-capita income as $RGDPPC_i = \frac{w_i}{p_{ni}} + k_i \frac{r_i}{p_{ni}}$.

[15] There, we use $\hat{\theta} = 0.14$. However, the results of interest are not very sensitive to the value of θ. The correlation coefficient between $e^{-\hat{\xi}_{0j}}$ and real per-capita income is only slightly higher for higher values of θ.

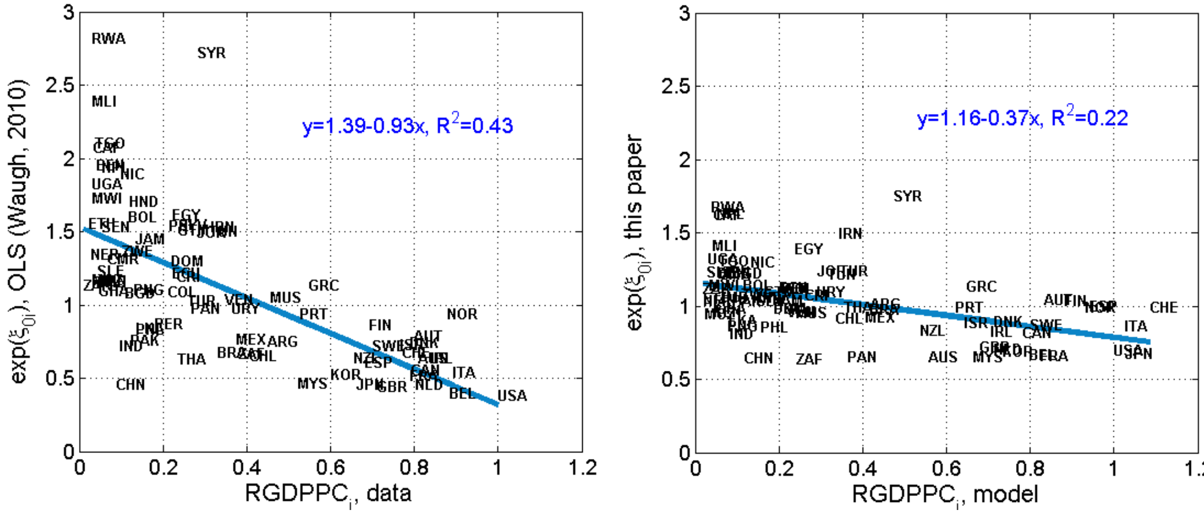

Figure 4: Exporter Fixed Effects and RGDPPC

imposition of multilaterally balanced trade in this paper unlike in Waugh; (ii) the comparison of trade cost estimates with model predictions rather than data on real per-capita income; and (iii) the use of nonlinear rather than linear model estimates so that zero trade flows are not omitted in this paper. Issues (i) and (ii) are crucial for the outcome.

Notice that correlations as in Figure 4 have only limited value for an assessment of trade cost asymmetry effects on real per-capita income in a nonlinear general equilibrium model. The reason is that real per-capita income also depends on other fundamentals than observable and unobservable trade costs. We need to flesh out effects of trade costs by way of comparative static analysis.

4.3 Comparative static analysis

In what follows we investigate Waugh's question, namely to which extent exporter trade cost asymmetry impacts the variance in real per-capita incomes, by way of comparative static analysis. For this, we consider the variance in log real per-capita income, $\mathrm{var}(\ln RGDPPC_i)$, and the 90-to-10 percentile gap in the distribution of real per-capita income $RGDPPC_{90}/RGDPPC_{10}$ at benchmark trade costs and at counterfactual trade costs.

Waugh intends to reduce (or eliminate) exporter trade cost asymmetry and consider the response of real per-capita income, $RGDPPC_i$. For this, he sets counterfactual trade costs at $t^c_{ij} = \min(t_{ij}, t_{ji})$. Certainly, this equalizes *bilateral* trade costs between countries i and j. Yet, it does neither ensure a reduction in the *asymmetry* of ξ_j across *exporters* j, nor does it leave the average *level* of trade costs unaffected. Accordingly, the conducted experiment in Waugh can not help answering the question at stake.[16]

In our view, there are two types of comparative static experiments to shed light on the question at stake. 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 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 $\xi_{\kappa,ij}^c$ for each exporter of the form:

$$\xi_{\kappa,j}^c = \hat{\xi}_j (1-\kappa) \text{ , where } \kappa \in \{0, 0.01, \ldots, 1\}. \tag{4.6}$$

While $\xi_{0,j}^c$ corresponds to the benchmark estimate $\hat{\xi}_j$, we eliminate all asymmetries in ξ_j at $\xi_{1,j}^c = 0$. As κ increases, the variance of the distribution of fixed exporter-specific trade costs degenerates. Let us denote counterfactual trade costs at $\xi_{1,ij}^c = 0$ by t_{ij}^c. Subsequently, we gradually reduce t_{ij}^c to unity by defining values of $t_{\omega,ij}^c$ for each country pair of the form:

$$t_{\omega,ij}^c = t_{ij}^c (1-\omega) \text{ , where } \omega \in \{0, 0.01, \ldots, 1\}. \tag{4.7}$$

Hence, as ω rises, t_{ij}^c approaches unity. In Figure 5, we illustrate the response of two measures of real per-capita income dispersion – the variance, var($\ln RGDPPC_i$), at the top and the 90-to-10 percentile gap, $RGDPPC_{90}/RGDPPC_{10}$, at the bottom – to an increase in κ (less heterogeneity in unobservable exporter-specific trade costs, $\hat{\xi}_j$) and in ω (lowering

[16]To see that Waugh's experiment does not necessarily reduce trade cost asymmetry, consider the following example of three exporters and importers. Assume that bilateral trade costs in the benchmark situation, $\{t_{21}, t_{31}, t_{12}, t_{32}, t_{13}, t_{23}\}$, take on values of $\{1, 2, 2, 1, 2, 1\}$. Then, with $t_{ij}^c = \min(t_{ij}, t_{ji})$, the counterfactual set of trade costs is $\{1, 2, 1, 1, 2, 1\}$. Obviously, average exporter-specific trade costs $\bar{t}_j = 0.5(t_{ij} + t_{kj})$ for all countries are $\{1.5, 1.5, 1.5\}$ in the benchmark case but $\{1.5, 1, 1.5\}$ in the counterfactual situation. Average exporter-specific trade costs $\bar{t} = \sum_{j=1}^{3} \bar{t}_j$ are 1.5 in the outset but 4/3 in the counterfactual. Hence, the variance of $\bar{t}_j$ is zero initially but larger than zero after the change. This proves the claim by example.

observable trade costs in counterfactual equilibrium, $\hat{t}^c_{ij}$), respectively. At the vertical delimiter, $\xi^c_j = 0$ for all j and $\hat{t}^c_{ij}$ is perfectly symmetric *bilaterally* for all ij.[17] In the figure, we conduct this experiment for Waugh's (indicated by solid loci) and our calibration of the model (indicated by broken loci).[18] 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. $RGDPPC_{90}/RGDPPC_{10}$ declines by 6.2% when using our calibration and var($\ln RGDPPC_i$) by only 1.5%. In either case, the response is somewhat larger when using Waugh's calibration, which is inferior in capturing key moments of the data as shown above.[19]

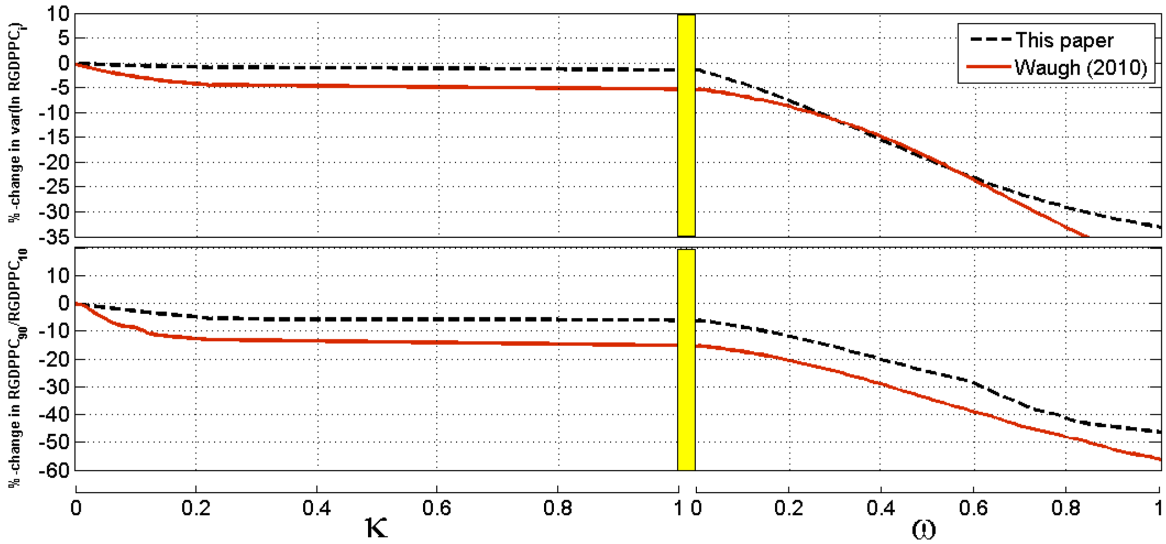

Figure 5: Comparative Static Analysis: Experiment 1

The right panel in Figure 5 suggests that reducing observable trade costs t^c_{ij} gradually leads to relatively bigger responses in the income dispersion than reducing unobservable exporter-specific trade cost heterogeneity. However, we have to drastically reduce t^c_{ij} in order to reduce income inequality by 30%. This suggests that removing the asymmetry in trade costs does not lead to significant average reductions in cross-country differences in real incomes. Hence, better access to international markets for developing countries will induce insignificant improvements in their per-capita incomes relative to those of the developed countries.

[17] Notice that observable trade costs t^c_{ij} are still heterogenous across countries after setting $\xi^c_j = 0$ 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*.

[18] Notice that Waugh (2010) assumed $t_{ij} = A$ with A being an arbitrary large number for all country pairs where $X_{ij} = 0$. This is inconsistent with data on observable trade costs and estimates of the corresponding parameters and unobservable exporter-specific trade costs in (2.5) and (2.4).

[19] Waugh's calibration leads to a decline of var($\ln RGDPPC_i$) and $RGDPPC_{90}/RGDPPC_{10}$ in response to the elimination of exporter-specific trade costs ($\xi^c_j = 0$) by 5.3% and 15.2%, respectively. This is larger than with the preferred calibration but still much smaller than Waugh's result.

Experiment 2

Experiment 2 allows us to disentangle changes in the *asymmetry* versus the *level* of overall (observable *and* unobservable) trade costs more appropriately than Experiment 1. Here, we stepwise reduce trade costs by defining counterfactual values $t^c_{\kappa,ij}$ for each pair of countries:

$$t^c_{\kappa,ij} = \hat{t_{ij}} + \kappa(\hat{\bar{t}} - \hat{t_{ij}}) \text{ , where } \kappa \in \{0, 0.01, \ldots, 1\}. \tag{4.8}$$

Notice that $t^c_{0,ij}$ corresponds to the benchmark estimate of bilateral trade costs, whereas $t^c_{1,ij} = \bar{t}$.[20] After eliminating completely trade cost asymmetries at $\kappa = 1$ and $t^c_{1,ij} = \bar{t}$, 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*:

$$\bar{t}^c_{\omega} = \hat{\bar{t}}(1 - \omega) \text{ , where } \omega \in \{0, 0.01, \ldots, 1\}. \tag{4.9}$$

As ω approaches unity, $\bar{t}^c_{\omega}$ converges towards unity so that $\bar{t}^c_0 = \hat{\bar{t}}$ and $\bar{t}^c_1 = 1$. Both the responses of $\text{var}(\ln RGDPPC_i)$ and of $RGDPPC_{90}/RGDPPC_{10}$ to changes in κ and ω are summarized in Figure 6 for our calibration and Waugh's.

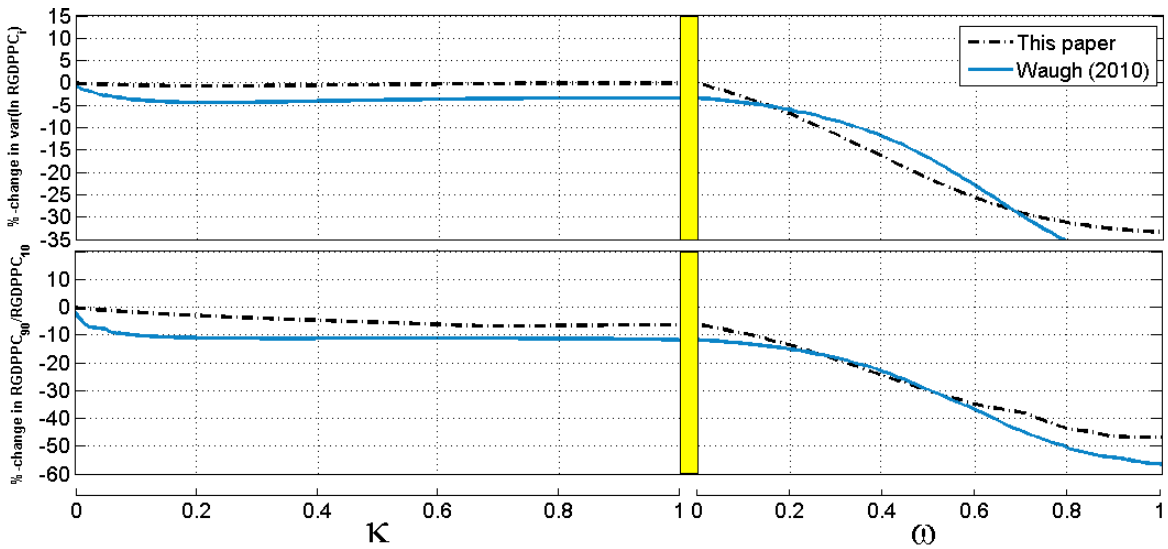

Figure 6: Comparative Static Analysis: Experiment 2

Again the left panels of Figure 6 suggest that eliminating trade cost asymmetries has little bearing for the dispersion of real per-capita income around the globe. While Waugh's calibration leads to somewhat bigger responses than ours, the predicted changes in $\text{var}(\ln RGDPPC_i)$ and $RGDPPC_{90}/RGDPPC_{10}$ to a complete abolition of trade cost asymmetry amount to 3.4% and 11.6%, respectively. Our calibration suggests corresponding responses of almost zero percent and about 6.4%, respectively. Quite obviously, a reduction

[20] In our benchmark model $\hat{\bar{t}} = 2.6$ while it is $\hat{\bar{t}} = 5.1$ in Waugh's (2010) model.

in income levels would have much greater effects on the outcomes of interest. To achieve a reduction in real per-capita income dispersion by about 30% 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 50% (in the lower panel of Figure 6) to 70% (in the upper panel of Figure 6).

5 Sensitivity analysis: a Dekle, Eaton, and Kortum (2007) model version with unbalanced trade

Dekle, Eaton, and Kortum (2007) suggest solving for wages from multilateral trade balance conditions for tradable goods subject to observed trade deficits. We calibrate and solve a version of their model to conduct counterfactual experiments as an alternative to Waugh's model which assumes multilaterally balanced trade. However, for the question at stake one has to make specific assumptions about the response of trade imbalances to shocks in trade costs. In what follows, we approach this assumption from the relationship of estimated fixed exporter-specific trade costs and data on trade imbalances.

Above, we assumed that normalized trade imbalances as defined in (3.1) were zero such that $B_i = 0$ for all i. This was consistent with the regular trade balance condition in (4.4). Here, we relax the assumption of multilaterally balanced trade.[21]

Let T_{ti} denote the $t$otal value of tradable goods produced in i. Next, suppose that country i's consumption of tradables, Y_{ti}, does not equal its production and recall our measure of trade imbalances from Section 1:

$$B_i = \frac{\sum_{j=1}^{N}(M_{ij} - M_{ji})}{Y_{ti}}. \tag{5.1}$$

Also recall that $M_{ij} = X_{ij}Y_{ti}$ and $M_{ji} = X_{ji}Y_{tj}$. We substitute this expression for M_{ij} into (5.1) to arrive at the trade balance condition in terms of trade shares and total spending on tradables:

$$(1 - B_i)Y_{ti} = \sum_{j=1}^{N} X_{ji}Y_{tj}. \tag{5.2}$$

Unlike (4.5) which we used to solve for wages in Section 4, we relax the assumption that $B_i = 0$ or $Y_{ti} = T_{ti}$ for all i and use (5.2) to establish the new identity between consumption

[21] To obtain new estimates of trade costs $t_{ij}^{-\frac{1}{\theta}}$ and of $c_{tj}^{-\frac{1}{\theta}}$, we estimate (2.4) subject to $\sum_{j=1}^{N}(M_{ij} - M_{ji}) = D_i$, where D_i is the nominal trade deficit of country i as observed in the data. Details on the corresponding results can be found in Table 2 of the Appendix.

and production of tradables in i by way of $(B_i - 1)T_{ti} = Y_{ti}$. Then, the new trade (im)balance condition that is solved for wages reads

$$L_i w_i = \sum_{j=1}^{N} X_{ji} \frac{L_j w_j}{1 - B_j}. \tag{5.3}$$

While this model helps predicting further features of the data (i.e., trade imbalances) in comparison to the benchmark estimates, its predictions on real GDP per capita do not match the data as well as the benchmark estimates: as expected, predictions of real per-capita incomes in developed countries (displaying trade surpluses) exceed but ones in developing countries (displaying trade deficits) fall short of the levels in the data.[22] Altogether, this model outperforms Waugh's (2010) but not our benchmark model in terms of matching data on trade flows and real per-capita incomes.

As said before, additional assumptions have to be adopted with regard to the response of B_i to changes in trade costs. Consistent with Figure 2, we assume that B_i is sensitive to trade cost asymmetries, and we assume that B_i approaches zero as ξ_i approaches zero for all i.[23]

We can now assess the sensitivity of the results with regard to eliminating trade cost asymmetries obtained from Experiments 1 and 2 to the use of the Dekle, Eaton, and Kortum (2007) version of the model in comparison to the benchmark estimates. It turns out that reducing trade cost asymmetries so that $B_i = 0$ for all i in counterfactual equilibrium actually leads to an *increase* in var($\ln RGDPPC_i$) of 2% in Experiment 1 and of 4.9% in Experiment 2.[24] The reason for this is the following. There are two forces that simultaneously affect the international income distribution: elimination of trade imbalances and elimination of asymmetries in trade costs. The former has a negative effect on income equality because a reduction of the absolute value of B_i leads to a reduction in earnings in countries with relatively higher trade deficits and vice versa.[25] These results are in line with Dekle, Eaton, and Kortum (2007). On the other hand, elimination of trade cost asymmetries exerts a minor effect towards income equalization. However, the latter is dominated by the former.

Suppose we assumed instead that B_i is completely insensitive to changes in trade costs instead in this model cum trade imbalances. Then, var($\ln RGDPPC_i$) and $RGDPPC_{90}/RGDPPC_{10}$ would decline when eliminating the asymmetry in trade costs,

[22]The cum-imbalances model also fits the data worse than the benchmark model in terms of the two inequality measures we employ. More details can be found in the Appendix.

[23]Notice that Figure 2 indicates that $B_i \approx 0.2$ at $\xi_i = 0$. However, by Walras' law I_i can be set at any constant number for all i with the same consequences for var($\ln RGDPPC_i$) and $RGDPPC_{90}/RGDPPC_{10}$ in counterfactual equilibrium.

[24] The results are nearly identical when we measure inequality through $RGDPPC_{90}/RGDPPC_{10}$.

[25]Figure 1 suggests that developing countries display relatively higher trade deficits than developed countries.

as in the benchmark case in Section 4. Yet, the responses would again be negligible: var(ln $RGDPPC_i$) would decrease by 1.8% in Experiment 1 and by only 0.3% in (the cleaner) Experiment 2. Hence, our original conclusions remain intact: abolishing trade cost asymmetries has only little bearing for the dispersion in real per-capita incomes around the globe.

6 Conclusion

This paper disentangles trade-imbalances from pure trade-cost-asymmetries to asses the importance of the latter in explaining international income differences. In a recent article, Waugh (2010) found that abolishing the asymmetry of exporter-specific trade costs would lead to a reduction in the global dispersion of real per-capita incomes by almost one-third. Using the same theoretical model and data, we provide starkly different results and conclude that eliminating exporter-specific trade cost heterogeneity on a global level would have negligible effects on real per-capita income dispersion. A general reduction in trade costs might significantly reduce differences in real per-capita incomes, but these reductions would have to be severe and largely counteract geographical or cultural barriers to trade.

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Appendix

Difference between OLS and PPML in Waugh (2010)

The key differences between the results in this paper and Waugh (2010) are *not* driven by the use of PPML versus OLS estimation. To see this, we present PPML counterparts to the OLS results in the left panels of Figures 2 and 4 below.

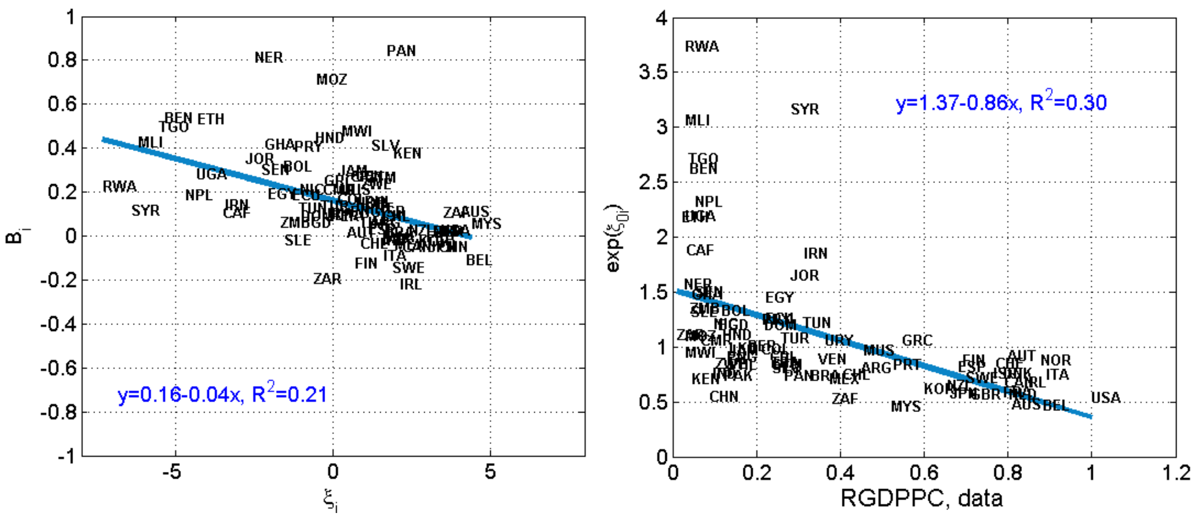

Figure 7: PPML Estimates of Waugh (2010)

The left panel of Figure 7 confirms that the OLS-based results regarding the relationship between B_j and $\hat{\xi}_j$ in Figure 2 are unaltered when using PPML instead. The right panel of Figure 7 suggests that the relationship between $e^{\hat{\xi}_{0j}}$ and data on real per-capita income ($RGDPPC_i$) in Figure 4 does not accrue to the omission of country pairs with zero bilateral trade flows and OLS estimation of the log-linearized normalized gravity equation.

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 $\gamma = 0.85$ as in Waugh helps improving the fit of the variance in real per-capita income, var$(\ln RGDPPC_i)$, as well as the 90-to-10 percentile ratio, $RGDPPC_{90}/RGDPPC_{10}$. The fit of the calibration in these respects is summarized in Table 3.

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 $\beta = 1/3$ from data in the examined country sample. The choice of $\alpha = 1/3$ 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,[26] 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 2. The coefficient on the tariff measure is -7.36, so that $\hat{\theta} \approx 0.14$.

The relatively good fit of the proposed calibration relative to Waugh's is robust to reasonable perturbations of θ and γ. To see this, consider Figure 8 where we plot counterparts to Figure 3: In the left panel, we employ Waugh's (2010) estimates together with $\theta = 0.14$ and $\gamma = 0.60$, and in the right panel we employ our estimates together with $\theta = 0.18$ and $\gamma = 0.75$.

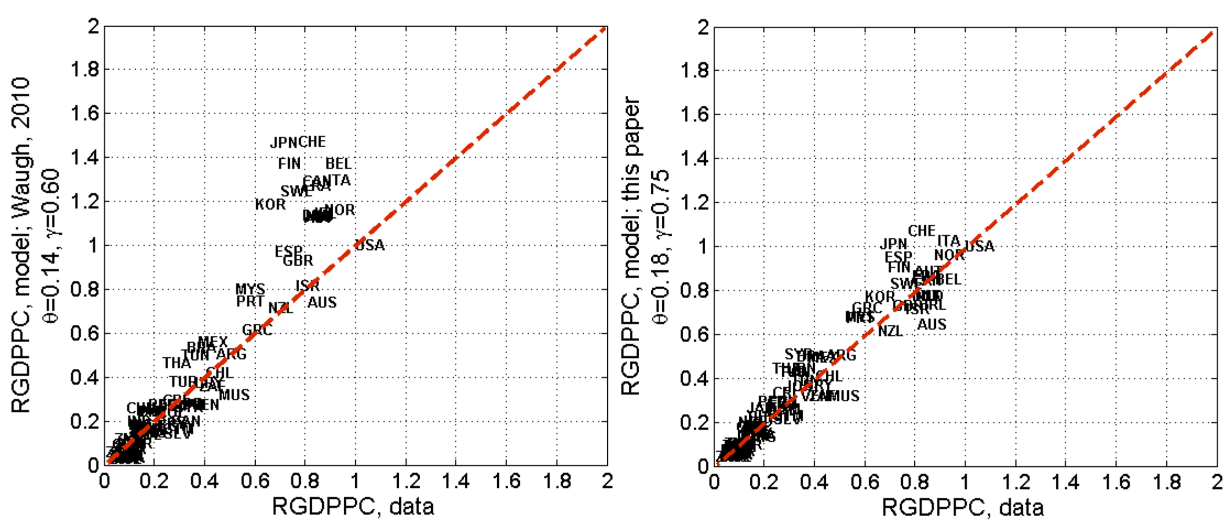

Figure 8: Fit of Calibration of Waugh (2010) Versus This Paper ($\theta = 0.14$,$\gamma = 0.60$) versus this paper ($\theta = 0.18$,$\gamma = 0.75$)

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

Table 2: PPML estimates (i) with tariffs and (ii) unbalanced trade

	Model with Tariffs				Model with Unbalanced Trade			
country	ξ_i		s_i		ξ_i		s_i	
United States	1.79	(0.23)	1.91	(0.13)	3.66	(0.30)	1.93	(0.14)
Argentina	-0.15	(0.13)	1.38	(0.08)	0.84	(0.13)	1.30	(0.08)
Australia	2.65	(0.47)	-0.22	(0.10)	4.15	(0.45)	0.11	(0.10)
Austria	-1.03	(0.18)	1.28	(0.06)	0.34	(0.16)	1.82	(0.05)
Belgium	2.19	(0.45)	-0.78	(0.21)	3.73	(0.39)	-0.24	(0.22)
Benin	-1.59	(0.30)	-0.38	(0.11)	-4.09	(0.22)	-0.35	(0.11)
Bangladesh	0.68	(0.43)	0.94	(0.04)	-0.58	(0.06)	1.34	(0.05)
Bolivia	-1.13	(0.10)	-0.09	(0.07)	-2.68	(0.11)	0.40	(0.10)
Brazil	0.19	(0.11)	1.88	(0.06)	1.06	(0.11)	1.98	(0.06)
Central African Republic	-3.05	(0.17)	0.18	(0.12)	-2.55	(0.14)	0.26	(0.11)
Canada	0.83	(0.37)	0.86	(0.26)	2.11	(0.35)	1.21	(0.28)
Switzerland	-0.49	(0.21)	1.12	(0.15)	0.70	(0.21)	1.96	(0.15)
Chile	0.5	(0.13)	0.42	(0.09)	1.43	(0.11)	0.54	(0.09)
China-Hong Kong	4.3	(0.32)	0.39	(0.20)	3.50	(0.19)	1.55	(0.15)
Cameroon	-0.9	(0.17)	-0.1	(0.09)	-0.79	(0.15)	0.02	(0.09)
Colombia	-0.78	(0.22)	1.22	(0.10)	-0.11	(0.29)	0.98	(0.08)
Costa Rica	-0.81	(0.35)	0.07	(0.12)	0.02	(0.36)	-0.02	(0.09)
Denmark	0.08	(0.20)	0.49	(0.06)	1.49	(0.17)	1.04	(0.07)
Dominican Republic	0.27	(0.25)	-0.31	(0.23)	1.07	(0.29)	-0.17	(0.24)
Ecuador	-0.86	(0.10)	0.38	(0.09)	-0.51	(0.16)	0.47	(0.08)
Egypt	0.38	(0.51)	2.16	(0.07)	-2.26	(0.13)	1.67	(0.04)
Spain	-0.75	(0.24)	1.93	(0.14)	0.80	(0.25)	2.28	(0.17)
Ethiopia	0.19	(0.16)	-0.71	(0.06)	-4.66	(0.11)	-0.46	(0.07)
Finland	-1.03	(0.12)	1.17	(0.07)	0.61	(0.08)	1.84	(0.05)
France	2.25	(0.71)	0.62	(0.17)	3.56	(0.70)	1.23	(0.17)
United Kingdom	1.6	(0.39)	0.91)	(0.09)	3.02	(0.39)	1.41	(0.10)
Ghana	0.33	(0.16)	-0.92	(0.12)	0.32	(0.18)	-0.74	(0.13)
Greece	-1.66	(0.20)	1.43	(0.08)	-0.45	(0.15)	1.33	(0.07)
Guatemala	-0.83	(0.21)	0.07	(0.08)	-0.12	(0.26)	0.08	(0.08)
Honduras	-0.59	(0.20)	-0.48	(0.17)	-0.73	(0.22)	-0.40	(0.17)
India	3.22	(0.38)	0.55	(0.18)	2.12	(0.19)	1.44	(0.17)
Ireland	0.69	(0.18)	-0.35	(0.13)	2.39	(0.17)	0.17	(0.13)
Iran	-2.97	(0.10)	2.27	(0.08)	-3.54	(0.09)	1.95	(0.08)
Israel	0.43	(0.19)	0.16	(0.09)	1.71	(0.17)	0.53	(0.09)
Italy	0.31	(0.29)	1.68	(0.15)	1.19	(0.28)	2.75	(0.16)
Jamaica	-0.2	(0.22)	-0.5	(0.20)	0.14	(0.26)	-0.37	(0.21)
Jordan	-0.76	(0.20)	0.44	(0.06)	-2.20	(0.13)	0.68	(0.04)
Japan	2.29	(0.23)	1.46	(0.11)	2.81	(0.26)	2.87	(0.10)
Kenya	1.19	(0.23)	-0.72	(0.06)	0.77	(0.16)	-0.79	(0.07)
Republic of Korea	2.73	(0.27)	0.39	(0.14)	3.11	(0.29)	1.57	(0.19)
Sri Lanka	1.43	(0.18)	-1.01	(0.07)	1.86	(0.09)	-0.72	(0.06)
Mexico	0.89	(0.30)	1.15	(0.30)	1.37	(0.51)	1.66	(0.50)
Mali	-2.94	(0.28)	0.15	(0.15)	-6.38	(0.26)	0.25	(0.15)
Mozambique	0.18	(0.15)	-2.1	(0.15)	-0.30	(0.14)	-1.99	(0.12)
Mauritius	1.08	(0.10)	-0.79	(0.07)	1.11	(0.09)	-0.69	(0.07)
Malawi	-0.87	(0.22)	-0.91	(0.17)	-1.18	(0.33)	-0.93	(0.16)
Malaysia-Singapore	3.24	(0.20)	0.22	(0.15)	3.81	(0.17)	0.57	(0.16)
Niger	-0.38	(0.42)	-2.83	(0.34)	-0.94	(0.41)	-2.78	(0.34)
Nicaragua	-2.2	(0.16)	0.15	(0.08)	-1.46	(0.16)	0.25	(0.08)
Netherlands	1.75	(0.20)	0.03	(0.09)	3.15	(0.17)	0.56	(0.09)
Norway	-0.52	(0.11)	0.77	(0.08)	0.62	(0.08)	1.28	(0.06)
Nepal	-2.76	(0.18)	0.77	(0.09)	-3.05	(0.09)	1.16	(0.10)
New Zealand	0.68	(0.19)	-0.04	(0.14)	2.27	(0.19)	0.29	(0.14)
Pakistan	1.87	(0.54)	0.27	(0.17)	0.88	(0.31)	0.76	(0.18)
Panama	2.81	(0.60)	-2.1	(0.55)	0.81	(0.61)	-2.00	(0.54)
Peru	-0.68	(0.12)	0.9	(0.06)	-0.14	(0.15)	0.94	(0.07)
Philippines	1.21	(0.15)	0.17	(0.12)	1.83	(0.17)	0.65	(0.13)
Papua New Guinea	1.42	(0.28)	-1.86	(0.23)	2.18	(0.28)	-1.73	(0.23)
Portugal	-0.71	(0.25)	1.03	(0.09)	0.89	(0.21)	1.19	(0.07)
Paraguay	-0.21	(0.12)	-0.04	(0.09)	-4.78	(0.12)	0.11	(0.11)
Rwanda	-3.1	(0.15)	0.72	(0.05)	-12.83	(0.06)	0.74	(0.05)
Senegal	-1.57	(0.28)	-0.07	(0.15)	-1.28	(0.22)	0.10	(0.16)
Sierra Leone	-2.3	(0.24)	-1.54	(0.19)	-0.16	(0.26)	-1.49	(0.19)
El Salvador	-0.52	(0.17)	-0.55	(0.10)	-0.13	(0.21)	-0.51	(0.08)
Sweden	0.23	(0.13)	0.78	(0.07)	1.72	(0.08)	1.53	(0.05)
Syrian Arab Republic	-3.29	(0.20)	2.65	(0.06)	-5.08	(0.10)	2.35	(0.04)
Togo	-2.09	(0.15)	-0.31	(0.11)	-8.48	(0.12)	-0.15	(0.11)
Thailand	1.15	(0.23)	1.26	(0.13)	0.63	(0.15)	1.92	(0.14)
Tunisia	-0.31	(0.26)	0.93	(0.10)	-0.90	(0.15)	1.32	(0.09)
Turkey	-1.75	(0.15)	1.9	(0.07)	-0.63	(0.16)	1.92	(0.08)
Uganda	-2.17	(0.06)	0.31	(0.06)	-3.06	(0.06)	0.38	(0.07)
Uruguay	-1.2	(0.12)	0.52	(0.09)	-0.03	(0.11)	0.45	(0.08)
Venezuela	0.3	(0.17)	0.18	(0.14)	0.64	(0.18)	0.18	(0.11)
South Africa	3.21	(0.21)	-0.61	(0.09)	4.11	(0.17)	-0.34	(0.08)
Democratic Republic of the Congo	-1.44	(0.09)	-0.47	(0.07)	0.78	(0.09)	-0.38	(0.08)
Zambia	-0.17	(0.24)	-0.94	(0.18)	0.62	(0.22)	-0.86	(0.18)
Zimbabwe	0.19	(0.24)	0	(0.13)	0.16	(0.28)	-0.03	(0.12)
Distance:								
$[0, 375)$	-4.19	(0.27)	-	-	5.54	(0.23)	-	-
$[375, 750)$	-4.44	(0.16)	-	-	-5.71	(0.12)	-	-
$[750, 1500)$	-4.17	(0.13)	-	-	-5.74	(0.07)	-	-
$[1500, 3000)$	-4.71	(0.22)	-	-	-6.59	(0.12)	-	-
$[3000, 6000)$	-5.72	(0.17)	-	-	-7.66	(0.14)	-	-
$[6000, max)$	-6.39	(0.19)	-	-	-8.45	(0.15)	-	-
Adjacency	1.08	(0.16)	-	-	1.05	(0.13)	-	-
Tariff	-7.36	(1.31)	-	-	-	-	-	-
Pseudo R^2	0.74				0.75			

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. The model with tariffs refers to the benchmark model as in Table 1 plus tariffs, while the model based on unbalanced trade refers to the version of the Dekle, Eaton, and Kortum (2007) model described and employed in Section 5.

In any case, our proposed calibration fits the data better in terms of the first two moments of the cross-country real per-capita income distribution but also in terms of the two inequality measures used in the counterfactual experiments (see Table 3).

Table 3: Fit of Calibration

			This paper[a]	
Statistics	Data	Waugh (2010)	Balanced trade	Unbalanced trade
θ	0.14	-	-	-
γ	0.60	-	-	-
var($\ln RGDPPC_i$)	1.38	2.16	1.39	1.60
$RGDPPC_{90}/RGDPPC_{10}$	25.59	71.45	20.99	28.24
Moments of $RGDPPC_i$				
Mean	0.34	0.44	0.35	0.38
Variance	0.09	0.26	0.10	0.15
θ	0.18	-	-	-
γ	0.75	-	-	-
var($\ln RGDPPC_i$)	1.38	1.31	0.91	-
$RGDPPC_{90}/RGDPPC_{10}$	25.59	26.55	12.60	-
Moments of $RGDPPC_i$				
Mean	0.34	0.47	0.40	-
Variance	0.09	0.20	0.10	-

Notes: a: The model with balanced trade corresponds to the one outlined in Sections 2-4, while the one based on unbalanced trade refers to the version of the Dekle, Eaton, and Kortum (2007) model in Section 5.