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17 March 2023

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March 17, 2023

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Abstract

In seminal studies, Feenstra (1994) and Broda and Weinstein (2006) develop an econometric method to quantify the welfare gains from trade associated with increases in product varieties. While it is the most widely respected structural approach, it relies on a contentious assumption: the orthogonality of import demand and supply residuals. This paper demonstrates theoretically that including additional controls for exporter-year factors increases the likelihood that the orthogonality condition is satisfied in practice. In particular, the extended model adresses the potential correlation between product quality (demand shocks) and exporter characteristics (supply shocks). The paper then implements the benchmark and the extended model in international trade data. Three main results emerge from the analysis. First, F-tests support the extended model over the constraint (benchmark) model. Second, the mean and standard deviation of the estimated elasticities are both smaller in the extended model compared to the benchmark model. Third, the correlation between the two sets of estimates is positive (as expected) but small, at 0.33. Overall, these empirical results suggest that it is important to control for exporter-year factors when estimating import-demand elasticities.

Keywords: Export supply, Import demand, International trade, Structural estimation, Trade elasticities.

JEL Classification Numbers: F1.

1 Introduction

A series of theoretical models shows that trade liberalization increases welfare by providing access to new differentiated product varieties (e.g., Armington (1969), Krugman (1980), Arkolakis et al. (2008), and Feenstra (2018)). In seminal studies, Feenstra (1994) and Broda and Weinstein (2006) (henceforth F/BW) develop and implement an econometric method that estimates the welfare gains associated with increases in product varieties. Numerous subsequent studies apply (or build) on F/BW (e.g., Broda and Weinstein (2010), Soderbery (2015), Hottman et al. (2016), Feenstra and Weinstein (2017), Soderbery (2018), and Redding and Weinstein (2020)).

Quantifying welfare gains from product varieties requires industry-level import-demand elasticities, which are notoriously difficult to estimate. F/BW's approach is the most widely respected structural method to address the classical simultaneity issue associated with estimating these elasticities in international trade data. Using the assumption that demand and supply shocks are orthogonal, their method combines two structural equations, one for import-demand and one for import-supply, into a single estimating equation. The estimated coefficients are then used to recover measures for the underlying structural parameters of the model – one of which is the import-demand elasticity needed for welfare estimation.

A significant criticism directed towards F/BW is that the orthogonality condition required for identification may not hold (e.g., Soderbery (2015)). Beginning with Linder (1961), an empirical literature in international trade argues that there is a correlation between the quality of exported products and the characteristics of the exporting country. High income per capita countries are more productive (high supply shocks) *and* are more likely to produce high quality goods (high demand shocks) compared to lower income per capita countries (e.g., Hallak (2006), Hallak (2010), and Khandelwal (2010)). This suggests that the demand and supply shocks are positively correlated and, as a result, that the F/BW orthogonality condition does not hold in practice.

The first contribution of this paper is to suggest a simple amendement to F/BW's method that increases the likelihood that the orthogonality condition is satisfied. In a nutshell, F/BW suggest an heteroskedastic estimator that identifies coefficients using second moments of bilateral trade shares and prices (Rigobon (2003)).[1] The transformation of the data into variances and covariances implicitly controls for the impact of importer-time and time-invariant factors. However, it does not address the potential impacts of exporter-time factors such as exporter productivity and product quality. This paper shows that extending the benchmark model to include additional covariates solves this problem.

The second contribution of the paper is to compare estimates from the benchmark to the extended model. Three main results emerge from this analysis. First, F-tests provide strong support to the extended model over the constraint (benchmark) model. Second, the mean and standard deviation of the estimated elasticities of substitution are both smaller in the extended model compared to the benchmark model. Third, the correlation between the two sets of estimates is positive (as expected) but small, at 0.33. Overall, these empirical results suggest that it is important to control for exporter-year factors when estimating import-demand elasticities.

To the best of my knowledge, this is the first paper to assess the impact of controlling for exporter-time factors in F/BW's method. In his seminal study, Feenstra (1994) did not raise the issue because his estimation uses US import data only. In that case, there is not enough variation to control for exporter-time factors. However, subsequent studies (including Broda and Weinstein (2006)) include numerous importing countries. In that context, it is possible to exploit the richer panel structure of the data to control for exporter-time factors, in addition to importer-time, and time-invariant factors. Yet, the issue remains unaddressed.

The remainder of the paper proceeds as follow. Section 2 briefly reviews F/BW's method and motivates the inclusion of additional controls in the estimating equation. Section 3

[1] Starting from bilateral data on import prices and trade shares, F/BW take a first-difference over time to remove time-invariant factors, then another first-difference with respect to a reference exporter to remove importer-year factors, and finally average over time within each importer-exporter pair. After transformation, the data set contains variances and covariances of trade shares and prices.

presents empirical estimates using the benchmark estimation method F/BW and three richer specifications that control for the impact of exporter-time factors. Section 4 briefly concludes.

2 Econometric Method

This section begins with a brief review of the F/BW method. It then presents a simple extension that controls for exporter-time factors.

2.1 A Review of F/BW

Let i, j, g, and t index, respectively, exporters, importers, industries, and time periods. Following the literature, I define the structural import-demand and import-supply equations as follow:[2]

$$\Delta^k \ln s_{ijg,t} = -(\sigma_g - 1)\Delta^k \ln p_{ijg,t} + \varepsilon_{ijg,t}, \tag{1}$$

$$\Delta^k \ln p_{ijg,t} = \left(\frac{\omega_g}{1+\omega_g}\right)\Delta^k \ln s_{ijg,t} + \xi_{ijg,t}, \tag{2}$$

where $s_{ijg,t}$ denotes the bilateral import value as a share of overall import value, $p_{ijg,t}$ is the unit price, and $\varepsilon_{ijg,t} \equiv \Delta^k \ln b_{ijg,t}$ and $\delta_{ijg,t} \equiv \Delta^k \ln \eta_{ijg,t}$ are, respectively, demand and supply residuals that depend on demand and supply shocks, respectively denoted $b_{ijg,t}$ and $\eta_{ijg,t}$. The operator Δ^k denotes a double-difference with respect to time and to a reference country k, such that $\Delta^k \ln s_{ijg,t} = (\ln s_{ijg,t} - \ln s_{ijg,t-1}) - (\ln s_{kjg,t} - \ln s_{kjg,t-1})$. The parameter σ_g is the constant elasticity of substitution (CES) between varieties, while ω_g governs the import-supply elasticity.

[2] These equations are equivalent to equations (14) and (15) in Broda and Weinstein (2006), and (3) and (4) in Soderbery (2015).

As shown in F/BW, we obtain a single estimating equation by multiplying equations (1) and (2) and rearranging:

$$Y_{ijg,t} = \theta_{1g} X_{1ijg,t} + \theta_{2g} X_{2ijg,t} + \mu_{ijg,t}, \tag{3}$$

where $Y_{ijg,t} \equiv \left(\Delta^k \ln p_{ijg,t}\right)^2$, $X_{1ijg,t} \equiv \left(\Delta^k \ln s_{ijg,t}\right)^2$, $X_{2ijg,t} \equiv \left(\Delta^k \ln p_{ijg,t} \Delta^k \ln s_{ijg,t}\right)$, $\mu_{ijg,t} = \varepsilon_{ijg,t}\xi_{ijg,t}$,

$$\theta_{1g} = \frac{\omega_g}{(1+\omega_g)(\sigma_g - 1)}, \qquad \text{and} \quad \theta_{2g} = \frac{1 - \omega_g(\sigma_g - 2)}{(1+\omega_g)(\sigma_g - 1)}. \tag{4}$$

Equation (4) clearly shows that the two structural parameters of the model, σ_g and ω_g, completely determine the two unknown coefficients, θ_{1g} and θ_{2g}, in equation (3).

As explained in F/BW, the coefficients θ_{1g} and θ_{2g} cannot be consistently estimated from equation (3) because the error term $\mu_{ijg,t}$ is correlated with the explanatory variables that depend on prices and expenditure shares. To obtain consistency, F/BW exploit the time-series dimension of the data set and the assumption that the unobserved demand and supply shocks are independent:

F/BW Moment Condition : $\mathbb{E}(\mu_{ijg,t}) \equiv \mathbb{E}(\Delta^k \ln b_{ijg,t} \Delta^k \ln \eta_{ijg,t}) = 0.$

F/BW implement the moment condition by taking the average over time of equation (3):

$$\overline{Y}_{ijg} = \theta_{1g} \overline{X}_{1ijg} + \theta_{2g} \overline{X}_{2ijg} + \overline{\mu}_{ijg}, \tag{5}$$

where $\overline{Z}_{ijg} = T^{-1} \sum_t Z_{ijg,t}$ for all $Z \in \{Y, X_1, X_2, \mu\}$. Because all variables are in double-differences, the regressors are second moments of the data. Therefore, in addition to the moment condition, identification requires across-country differences in the relative variance of the demand and supply shocks (see Feenstra (1994) equation (12) p.164).

2.2 An extension of F/BW

To better understand the restrictions implied by the F/BW moment condition, let's decompose the demand and supply shocks into time-invariant components ($b_{g,t}$ and $\eta_{g,t}$), exporter-time components ($b_{ig,t}$ and $\eta_{ig,t}$), importer-time components ($b_{jg,t}$ and $\eta_{jg,t}$), and residual components ($\nu_{ijg,t}$ and $\gamma_{ijg,t}$) as follows:

$$b_{ijg,t} = b_{g,t} b_{ig,t} b_{jg,t} \nu_{ijg,t}, \qquad \text{and} \qquad \eta_{ijg,t} = \eta_{g,t} \eta_{ig,t} \eta_{jg,t} \gamma_{ijg,t}. \tag{6}$$

The first difference with respect to time eliminates the time-invariant components, while the first-difference with respect to the reference exporter k eliminates the importer-time components. However, the exporter-time components and the residual components remain. As shown in Online Appendix A, when we explicitly distinguish the exporter-time components and the residual components, the estimating equation takes the following form:

$$\overline{Y}_{ijg} = \theta_{1g} \overline{X}_{1ijg} + \theta_{2g} \overline{X}_{2ijg} + \theta_{3ig} \overline{X}_{3ijg} + \theta_{4ig} \overline{X}_{4ijg} + \theta_{5ig} + \bar{\tilde{\mu}}_{ijg}, \tag{7}$$

where $\overline{X}_{3ijg} \equiv T^{-1} \sum_t \Delta^k \ln s_{ijg,t}$, $\overline{X}_{4ijg} \equiv T^{-1} \sum_t \Delta^k \ln p_{ijg,t}$, and $\bar{\tilde{\mu}}_{ijg} \equiv \Delta^k \ln \nu_{ij,t} \Delta^k \ln \gamma_{ij,t}$.

The extended model in equation (7) nests the benchmark model in equation (3) as a special case because the first two regressors, $\overline{X}_{1ijg}$ and $\overline{X}_{2ijg}$, and their corresponding coefficients, θ_{1g} and θ_{2g}, are the same in both models. The three additional regressors are: (i) a set of interaction terms between exporter fixed-effects and trade shares, $\theta_{3ig}\overline{X}_{3ijg}$, that controls for the impact of the exporter-specific components of demand and supply shocks on demand; (ii) a set of interaction terms between exporter fixed-effects and trade prices, $\theta_{4ig}\overline{X}_{4ijg}$, that controls for the impact of the exporter-specific components of supply and demand shocks on supply; and (iii) a set of exporter fixed effects, θ_{5ig}, that controls for the covariances between the exporter-time components of demand and supply shocks, and their correlation with trade shares and prices. Formal definitions for the additional coefficients (θ_{3ig}, θ_{4ig}, and θ_{5ig}) are presented in Online Appendix A.

The moment condition required for identification in the extended model is

$$\textbf{New Moment Condition} : \mathbb{E}(\tilde{\mu}_{ijg}) = \mathbb{E}(\Delta^k \ln \nu_{ij,t} \Delta^k \ln \gamma_{ij,t}) = 0.$$

In addition to the "New Moment Condition", the benchmark F/BW method also requires that $\theta_{3ig} = \theta_{4ig} = \theta_{5ig} = 0$. These additional assumptions are quite strong and unlikely to hold in the data. Furthermore, because it is straightforward to include the additional controls $\overline{X}_{3ijg}$, $\overline{X}_{4ijg}$, and θ_{5ig} to the model, they are not necessary.

3 Empirical Implementation

In this section, I report the results from implementing the estimation method developed in the previous section in international trade data.

3.1 Data

Estimating equation (7) requires data on trade shares and import prices (inclusive of trade costs) for each exporter-importer-industry-year observation. Information on the value of bilateral trade flows come from the United Nation's Comtrade database. It collects import values that include the transaction value of the goods (inclusive of the value of services performed to deliver goods to the border of the exporting country) and the value of the services performed to deliver the goods from the border of the exporting country to the border of the importing country.[3] Measures of policy barriers come from a dataset compiled by Feenstra and Romalis (2014). Using these measures, I can compute bilateral import values inclusive of transportation costs and tariffs from which trade shares can be calculated.

[3] As in Feenstra and Romalis (2014), I drop observations where the ratio of the c.i.f. unit value reported by the importer to the corresponding f.o.b. unit value reported by the exporter is less than 1 or exceeds 10 in an attempt to minimize the effect of measurement error on the estimation.

The Comtrade database does not contain direct information on trade prices. However, it contains information on the volume of trade in quantity unit. As common in this literature, I estimate prices using average unit values defined as total value over total quantity.[4] The final sample covers 778 four-digits Standard International Trade Classification (SITC) industries from 1984 to 2011.

3.2 Estimation

As typical in this literature, I estimate the model separately for each industry g in my sample. Because it would not be useful to report hundreds of parameters, I only present the distribution of the point estimates in Table 1.[5] The table reports results from estimating four different specifications: (i) the benchmark model (5) labeled "FBW" (corresponds to the case where $\theta_{3ig} = \theta_{4ig} = \theta_{5ig} = 0$ in equation (7)); (ii) a model with fixed effects but no interaction terms labeled "FE" (corresponds to the case where $\theta_{3ig} = \theta_{4ig} = 0$ in equation (7)); (iii) a model with interactions but no fixed effects labeled "INTER" (corresponds to the case where $\theta_{5ig} = 0$ in (7)); and (iv) the full model described in equation (7) labeled "FULL". For brevity, I follow Broda and Weinstein (2006) and only report results for the elasticity of substitution (σ_g), the parameter required for welfare analysis. The results for the elasticity of transformation (ω_g) are presented in Online Appendix B.

As reported in Table 1, the median estimated elasticity of substitution for the FBW model is 4.48. This result is in line with comparable estimates available in the literature.[6] Comparing FBW estimates to the FE and INTER estimates suggests that the interaction

[4]When possible, I convert physical units of measurement to a common denominator (e.g., "Thousands of items" to "Items"). For industries with multiple units of measurement, I keep only observations with the unit of measurement that accounts for the largest share of import value.

[5]These statistics exclude industries for which the estimates imply inadmissible values for the parameters (i.e., $\sigma_g < 1$ or $\gamma_g < 0$). About 35 percent of the estimates from Broda and Weinstein (2006) imply inadmissible values. In contrast, less than 10 percent of the industries (60 out of 778) in my sample imply inadmissible values for at least one of the four models presented in Table 1.

[6]For instance, Broda and Weinstein (2006) report medians elasticity of substitution ranging from 2.2 to 3.7, Broda and Weinstein (2010) report a median of 7.5, and Feenstra and Romalis (2014) report a median of 5.8. The variation across studies can be explained by changes in estimation procedure, level of aggregation, and sample composition.

terms may have a greater impact on the point estimates compared to the fixed effects. The median estimated elasticity of substitution is 4.58 for FE, whereas it is 3.28 for INTER. Interestingly, the FULL estimates seem to suggest that the impact of the exporter's fixed effects dominates the impact of the interaction terms. When all controls are included the estimated median of 4.37, which is very close to the benchmark FBW median of 4.48.

TABLE 1
DISTRIBUTION OF ESTIMATED ELASTICITY OF SUBSTITUTION

Percentile	FBW	FE	INTER	FULL
1	2.58	2.66	2.33	2.70
5	3.06	3.08	2.52	3.09
10	3.26	3.29	2.63	3.26
25	3.75	3.79	2.89	3.67
50	4.48	4.58	3.28	4.37
75	5.77	6.06	3.83	5.45
90	8.45	8.74	4.76	7.23
95	11.63	12.63	5.39	9.76
99	22.36	30.73	7.43	17.97
Mean	6.11	7.25	3.53	5.20
St. Dev.	10.53	26.00	1.02	3.67
Correlation	1.00	0.51	0.35	0.33

Notes: This table presents the distributions of the estimated structural parameters of the model obtained from estimating the theoretical model separately for each industries in the sample using four different models, as listed above each column. The bottom of the table reports the mean and the standard deviation for each set of estimates, as well as the correlation with the benchmark FBW estimates. The summary statistics are reported for the 718 (out of 778) industries for which all four estimates are consistent with the theoretical model (i.e., $\sigma_g > 1$ and $\omega_g > 0$).

However, the distributions in Table 1 mask important variations across models in industry-level estimates. As seen at the bottom of the table, the mean and the standard deviation of the FULL estimates are both smaller compared to those for the benchmark FBW estimates. Figure 1 presents a scatter plot of the FULL estimates against the benchmark FBW estimates. The figure clearly shows that the FULL estimates are generally smaller (most dots are to the right of the 45 degree line) and are more densely distributed compared to the benchmark estimates. Furthermore, the correlation between FULL and FBW estimates reported at the bottom of Table 1 is positive (as expected) but small, at 0.33. This suggests that for some industries the difference between the FBW and FULL estimates is significant. A quick look at Figure 1 confirms that in some cases the dots are quite far from the 45 degree line.

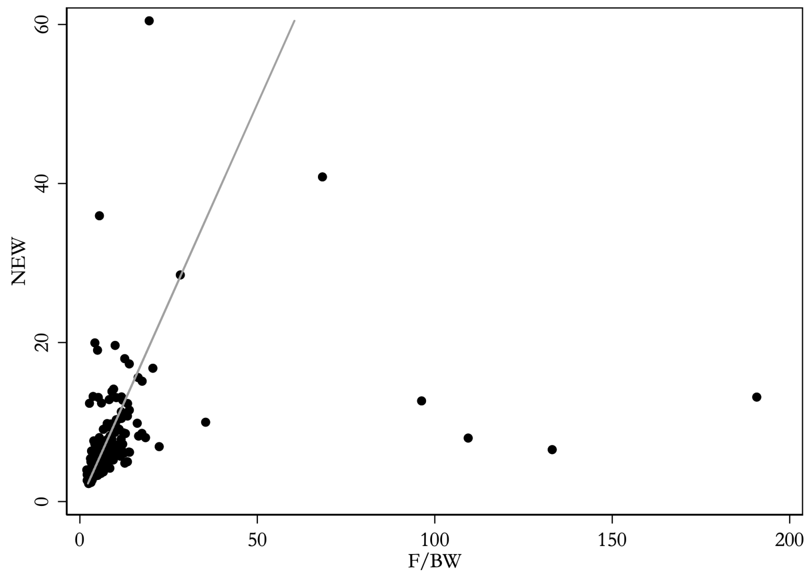

Figure 1: Scatter plots of benchmark against new estimates

The benchmark (FBW) estimating equation (5) is a constrained version of the extended model (FULL) in equation (7). To test the empirical validity of these restrictions, I perform an F-test for each industry in my sample. The distribution of the F-statistics for the null hypothesis that $\theta_{3ig} = \theta_{4ig} = \theta_{5ig} = 0$ is reported in Figure 2 (additional statistics are presented in Online Appendix Table C.1). The figure shows that the F-statistics are generally large. Combined with the large number of restrictions, the F-test rejects the null hypothesis 97 percent of the time at the 1 percent level (697 out of 718 industries). Overall, the results presented in this section suggest that controlling for exporter-time factors is statistically important when estimating import-demand elasticities.

4 Conclusion

In seminal studies, Feenstra (1994) and Broda and Weinstein (2006) develop an econometric method that estimates the import-demand elasticity. The current paper argues that the moment condition required for identification may fail because their method does not control

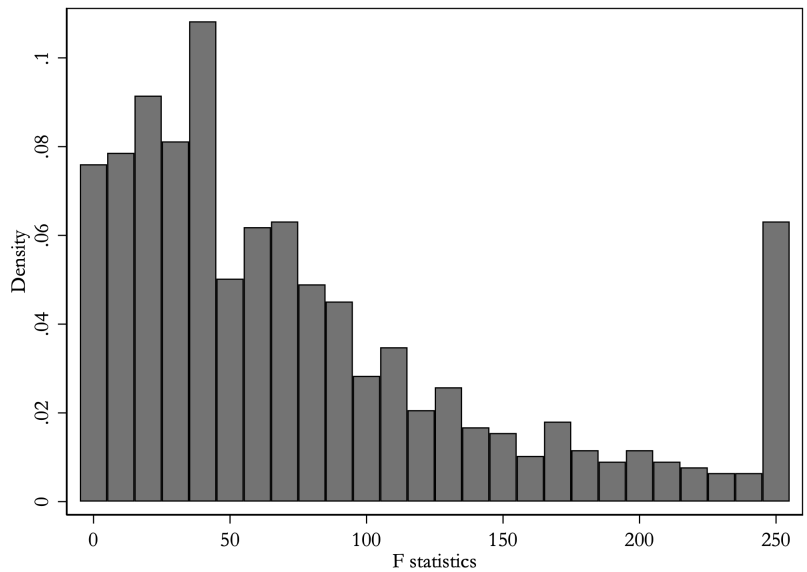

Figure 2: Histogram of F-statistics

for exporter-time effects. The paper then suggests an extended model that includes additional regressors and nests the benchmark model as a special case. The empirical results provide support to the extended model.

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Online Appendices

A Estimating Equation

This appendix briefly describes the steps required to obtain the estimating equation for the extended model (equation (7) in the main text).

Taking the log double-difference of the demand and supply shocks, defined in equation (6), eliminates the importer-year and the time-invariant factors:

$$\begin{aligned} \Delta^k \ln b_{ijg,t} &= \Delta^k \ln b_{ig,t} + \Delta^k \ln \nu_{ijg,t}, \\ \Delta^k \ln \eta_{ijg,t} &= \Delta^k \ln \eta_{ig,t} + \Delta^k \ln \gamma_{ijg,t}. \end{aligned} \tag{A.1}$$

Substituting with these results in equations (1) and (2), respectively, yields:

$$\begin{aligned} \Delta^k \ln s_{ijg,t} &= -(\sigma_g - 1)\Delta^k \ln p_{ijg,t} + \Delta^k \ln b_{ig,t} + \tilde{\varepsilon}_{ijg,t}, \\ \Delta^k \ln p_{ijg,t} &= \left(\frac{\omega_g}{1+\omega_g}\right) \Delta^k \ln s_{ijg,t} + \Delta^k \ln \eta_{ig,t} + \tilde{\xi}_{ijg,t}, \end{aligned} \tag{A.2}$$

where $\tilde{\varepsilon}_{ijg,t} \equiv \Delta^k \ln \nu_{ijg,t}$ and $\tilde{\xi}_{ijg,t} \equiv \Delta^k \ln \gamma_{ijg,t}$.

Taking the product of the two equations in (A.2) and rearranging yields:

$$\begin{aligned} \left(\Delta^k \ln p_{ijg,t}\right)^2 = &\left[\frac{\omega(\sigma-1)-2}{(\sigma-1)(\omega+1)}\right] \Delta^k \ln s_{ijg,t} \Delta^k \ln p_{ijg,t} + \frac{\omega}{(\sigma-1)(\omega+1)} \left(\Delta^k \ln s_{ijg,t}\right)^2 \\ &+ A_{ig,t}\, \Delta^k \ln s_{ijg,t} + B_{ig,t}\, \Delta^k \ln p_{ijg,t} - \frac{\Delta^k \ln b_{ig,t} \Delta^k \ln \eta_{ig,t}}{\sigma - 1} \\ &+ \frac{\tilde{\varepsilon}_{ijg,t}\tilde{\xi}_{ijg,t}}{\sigma-1}, \end{aligned} \tag{A.3}$$

where I introduced two new variables, $A_{ig,t}$ and $B_{ig,t}$, defined as:

$$\begin{aligned} A_{ig,t} &\equiv \frac{\Delta^k \ln \eta_{ig,t}}{\sigma-1} - \frac{\omega \Delta^k \ln b_{ig,t}}{(\sigma-1)(\omega+1)}, \\ B_{ig,t} &\equiv \frac{\Delta^k \ln b_{ig,t}}{\sigma-1} - \Delta^k \ln \eta_{ig,t}. \end{aligned} \tag{A.4}$$

Taking the expectation of equation (A.3) and using the fact that:

$$\begin{aligned}
\mathbb{E}(A_{ig,t}\Delta^k \ln s_{ijg,t}) &= \mathbb{E}(A_{ig,t})\mathbb{E}\left(\Delta^k \ln s_{ijg,t}\right) + \operatorname{cov}\left(A, \Delta^k \ln s_{ijg,t}\right), \\
\mathbb{E}(B_{ig,t}\Delta^k \ln p_{ijg,t}) &= \mathbb{E}(B_{ig,t})\mathbb{E}\left(\Delta^k \ln p_{ijg,t}\right) + \operatorname{cov}\left(B, \Delta^k \ln p_{ijg,t}\right),
\end{aligned} \tag{A.5}$$

we obtain estimating equation (7) in the main text:

$$\overline{Y}_{ijg} = \theta_{1g}\overline{X}_{1ijg} + \theta_{2g}\overline{X}_{2ijg} + \theta_{3ig}\overline{X}_{3ijg} + \theta_{4ig}\overline{X}_{4ijg} + \theta_{5ig} + \overline{\overline{\mu}}_{ijg}, \tag{A.6}$$

where the coefficients are defined as follows:

$$\begin{aligned}
\theta_{3ig} &\equiv \mathbb{E}(A_{ig,t}) = \frac{1}{\sigma-1}\mathbb{E}(\Delta^k \ln \eta_{ig,t}) - \frac{\omega}{(\sigma-1)(\omega+1)}\mathbb{E}(\Delta^k \ln b_{ig,t}), \\
\theta_{4ig} &\equiv \mathbb{E}(B_{ig,t}) = \frac{1}{\sigma-1}\mathbb{E}(\Delta^k \ln b_{ig,t}) - \mathbb{E}(\Delta^k \ln \eta_{ig,t}), \\
\theta_{5ig} &\equiv \operatorname{cov}\left(A_{ig,t}, \Delta^k \ln s_{ijg,t}\right) + \operatorname{cov}\left(B_{ig,t}, \Delta^k \ln p_{ijg,t}\right) - \frac{\mathbb{E}(\Delta^k \ln b_{ig,t}\Delta^k \ln \eta_{ig,t})}{\sigma-1}.
\end{aligned} \tag{A.7}$$

B Elasticity of Transformation

Table B.1 reports the estimation results for the elasticity of transformation. The general layout of the table is the same as in Table 1 (in the main text).

TABLE B.1
Distribution of estimated elasticity of transformation

Percentile	FBW	FE	INTER	FULL
1	0.71	0.70	0.23	0.60
5	1.07	1.17	0.45	1.13
10	1.30	1.42	0.59	1.36
25	1.79	1.91	0.89	1.84
50	2.80	2.93	1.39	2.73
75	4.76	5.00	2.26	4.68
90	8.59	8.37	3.94	8.65
95	13.15	13.92	5.93	12.33
99	41.96	43.53	10.25	33.57
Mean	4.82	6.22	2.19	5.24
St.Dev.	10.33	28.52	5.87	19.42
Correlation	1.00	0.57	0.20	0.81

Notes: This table presents the distributions of the estimated structural parameters of the model obtained from estimating the theoretical model separately for each industries in the sample as well as the correlations between the estimates and the benchmark FBW estimates. The summary statistics are reported for the 718 (out of 778) industries for which all four sets of estimates are consistent with the theoretical model (i.e., $\sigma_g > 1$ and $\omega_g > 0$).

The median estimates for the benchmark FBW model is 2.80. The extant literature provides few estimates of the elasticity of transformation for comparison. In particular, Broda and Weinstein (2006) and Broda and Weinstein (2010) do not report their estimates. Baier and Bergstrand (2001) report a point estimate of 8.6 with a 90% confidence interval of 1.4 to 15.8. This result is based on aggregate trade flows, so it is not directly comparable. Feenstra (1994) estimates the elasticity in disaggregated trade data for 8 products, with median values of 2.85 (if we consider the 4^{th} estimate as the median) or 7.78 (if we use the 5^{th} estimate). The most recent evidence comes from Hottman et al. (2016). They use scanner code data to estimate an elasticity of marginal production costs with respect to output of 0.18. While the point estimates vary across studies, there is significant overlap between the current estimates and those reported in prior works.

Comparing the benchmark FBW and the extended model FULL seems to suggest that the estimates are almost the same. The median estimates are 2.80 and 2.73, respectively. However, comparing the means and variances reveal that there are important differences across the two sets of estimates. The mean and the variance are both larger for the FULL estimates. This can also be seen in Figure B.1, which reports a scatter plot of the FULL against the FBW estimates. The figure shows that a large fraction of the dots lie to the left of the 45 degree line and that, for a number of industries, the FULL estimates are much larger than the corresponding FBW estimates. As indicated at the bottom of Table B.1, the correlation between the two sets of estimates remains quite large, at 0.81. These results are driven, in part, by an outlier seen in the upper region of Figure B.1. The correlation drops to 0.69 when removing the largest outlier.

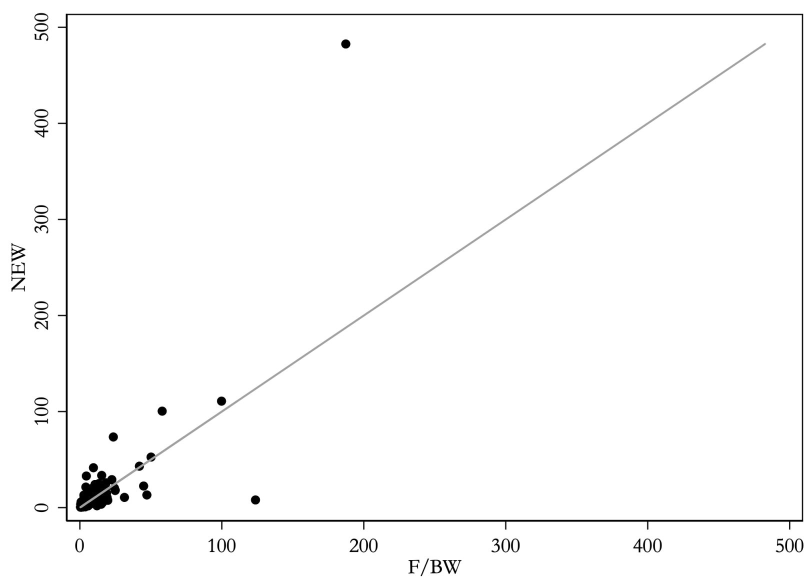

Figure B.1: Scatter plots of benchmark against new estimates

C F-tests

TABLE C.1
Distribution of f-test components

Percentile	N	df1	df2	F	p-value
1	236	2	114	2.01	0.00
5	472	2	298	7.76	0.00
10	787	2	580	13.32	0.00
25	1518	2	1253	30.07	0.00
50	2668	2	2345	59.58	0.00
75	3976	2	3589	108.51	0.00
90	5236	2	4748	179.80	0.00
95	6012	2	5525	232.35	0.00
99	7258	2	6737	378.52	0.13
Mean	2849	2	2520	81.98	0.00
St. Dev.	1688	2	1597	79.81	0.03

Notes: This table presents the distribution of the number of observations (N), the number of parameters in the constraint F/BW model (df1), the number of constraints (df2), the F-statistics, and the implied p-values obtained from estimating the models separately for each industries in the sample. The bottom of the table reports the mean and the standard deviation for each set of estimates. The summary statistics are reported for the 718 (out of 778) industries for which all four estimates are consistent with the theoretical model (i.e., $\sigma_g > 1$ and $\omega_g > 0$).