Measuring the Impact of Agricultural Production Shocks on International Trade Flows *

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Abstract

The purpose of this study is to measure the sensitivity of trade volumes and trade unit values to agricultural production shocks. We derive a gravity model of trade with two new features. First, the model assumes perfectly inelastic supply, which captures the short-run nature of year-to-year production shocks and has important implications for levels of regression coefficients and how they can be used to measure the elasticity of substitution. Second, the presence of per-unit trade costs implies that, in percentage terms, unit values based on importing country data (including trade costs) should react less to production shocks than unit values based on exporting country data (excluding trade costs). Using bilateral trade flow data for a large sample of countries and agricultural commodities we find empirical support for the predictions of the model, with relatively high substitutability between varieties of crops differentiated by country of origin and quantitatively large per-unit trade costs. Our framework provides a new method for measuring substitution elasticities and per-unit trade costs using international trade data. Furthermore, our results suggest that trade frictions or substitution with other goods diminish the role of international trade as way of coping with production volatility.

JEL Classification Codes: F14, F18, Q11, Q17, Q18 .

Keywords: Food price volatility, pass-through, agricultural trade

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The recent volatility in food prices has brought issues of food security to the forefront of the policy debate (Anderson et al., 2014). On the supply side, agricultural production is sensitive to weather conditions and it is negatively affected by extreme weather events. As climate change progresses, many types of extreme weather events such as heat waves, droughts and floods are expected to become more frequent (IPCC, 2013) which may contribute to food-price volatility in the years to come.[1]

International trade in food is one potential mechanism to mitigate the effects of production volatility.[2] Countries that experience a negative shock to agricultural production can reduce volatility in consumption by importing the products that they need. With this in mind, we will here consider how the effects of agricultural production shocks are propagated between countries through trade via their impact on quantities traded and the prices at which they trade, commonly referred to as "trade unit values".

In order to guide our empirical analysis we set up a general-equilibrium model of trade whereby production shocks in exporting countries affect both traded quantities and trade unit values. The model yields regression equations similar to standard gravity models. Two features of the model are adapted for our current analysis. First, our model assumes a perfectly inelastic short-run supply, with the prediction that a one percent change in the quantity produced translates into a one percent change in traded quantities and a $1/\sigma$ percent change in trade unit values, where σ is the elasticity of substitution. In contrast, a standard gravity model of international trade with iceberg trade costs and perfectly elastic supply predicts that a one percent change in productivity (unit costs) leads to a one percent change in trade unit values and a σ percent change in traded quantities. When considering short-run fluctuations in production of the kind we do, it seems more reasonable to consider production as the exogenous variable and then let prices and traded quantities adjust. Assuming that there is scope for substitution (i.e. a relatively large σ), the difference in prediction is sizable and the coefficient estimates can validate

[1] The connection between climate change and food production is a well researched topic within crop science. For example, Schlenker and Roberts (2009) find that temperatures above a certain threshold are very harmful to corn, soybean and cotton yields. IPCC (2014) provide an overview of the main results and a more detailed synthesis can be found in the full IPCC report. The importance of the connection from an economics perspective has also been long recognized. There are a number of studies investigating the role of trade as a means of adaption (see, e.g., Reilly and Hohmann, 1993; Rosenzweig and Parry, 1994; Tsigas et al., 1997; Randhir and Hertel, 2000).

[2] At least with well-functioning markets. See Gars and Spiro (2016) for an example of how trade may exacerbate shocks if production relies on renewable resources with imperfect property rights.

our modeling approach.

Second, we add per-unit trade costs to the model, which allows us to consider the difference in effects of production shocks on CIF ("Cost Insurance Freight"-including trade costs) and FOB ("Free On Board"-excluding trade costs) unit values. In particular, the presence of per-unit trade costs implies that the relative change in CIF unit values following production shocks should be smaller since the per-unit trade costs are not affected and dampen the effect in percentage terms. In contrast, a standard gravity model with only iceberg trade costs would imply that the elasticity of prices with respect to production should be the same for CIF and FOB unit values (since they are then proportional). Per-unit trade costs have been found to be quantitatively large in international trade of manufacturing goods (Irarrazabal et al., 2015), and given the bulky nature of agricultural commodities we expect that per-unit trade costs are a nascent feature of international agricultural trade.

Using data on bilateral trade of agricultural commodities, our goal is to estimate the sensitivity of trade volumes and trade unit values to fluctuations in production. In our baseline approach we exploit year-to-year changes in production, using yield per acre as our instrument for production. While yields can be influenced by prices in the medium- to long-term (Miao et al., 2016), Roberts and Schlenker (2013) find that detrended yield shocks appear to stem mainly from random weather shocks, with little risk that short-run yield fluctuations are endogenously determined by prices. We exploit the year-to-year variation in production between crops in the same country, which allows us to disentangle the effect of production on trade from other factors that vary by country and year, such as macroeconomic shocks. Since we do not observe actual prices in the trade data, we calculate the average price of trade flows, which we refer to as trade unit values in the rest of the article.

Our regression results yield two main findings. First, the elasticity of traded quantities with respect to production is around 0.5, while the effects on unit values are significantly smaller. The estimates thus support our modeling of short-run supply as being inelastic, with some scope for substitutability between varieties of a given crop differentiated by country of origin. Given that trade in agricultural commodities is small compared to total production, the result for traded quantities suggests that trade is relatively unresponsive to production shocks. The estimates thus suggest that there is room for improving trade as a mechanism for coping with food-production volatility. Second, there is a difference between effects of productivity shocks on CIF and FOB unit values, which implies the presence of large per-unit trade costs. The difference in the point estimates for the CIF and

FOB unit-value regressions allows us to calculate the average size of per-unit trade costs in the data, and the point estimates indicate per-unit trade costs that are 14-29% of import unit values.

Our use of production shocks and trade unit values to quantify per-unit trade costs is new, and builds on earlier literature that highlights the impact of per-unit trade costs on the pattern of trade. Hummels and Skiba (2004) show that the presence of per-unit trade costs encourages firms to export high-value-goods, which became known as the "Washington apples effect". Our work focuses on the importance of per-unit costs for understanding the pass-through of production shocks, and complements work by Chen and Juvenal (2016) who find that real exchange rate shocks in the presence of destination-country per-unit trade costs leads to more pricing-to-market for high quality goods, using evidence from international trade in wine.

Our gravity model of trade with inelastic supply contributes to a recent theoretical literature that relaxes the assumption of perfectly elastic supply (Vannoorenberghe (2012), Soderbery (2014), Soderbery (2015)).[3] We estimate substitution elasticities by directly exploiting data on observed production shocks, which departs from the standard approach in the international trade literature pioneered by Feenstra (1994) that infers unobserved supply shocks by combining trade unit value and trade volume data.[4]

Our study is also related to a growing literature on food-price volatility, particularly those studies that focus on transmission of productivity shocks via international trade. While many studies have measured the pass-though of world prices to domestic prices (Mundlak and Larson, 1992; Baffes and Gardner, 2003; Dawe, 2008; Ferrucci et al., 2012; Imai et al., 2008; Yang et al., 2015), few have considered the implications of trade costs on the magnitude of price pass-through.

The study most related to our work is Reimer et al. (2009), who adapt the Eaton and Kortum (2002) model of Ricardian trade to estimate the effect of higher yield volatility on trade and welfare. They find that increased yield volatility would lead to increased trade, and that the welfare losses from increased volatility are amplified by trade costs. Their study uses one year of cross-section data on trade and production to calibrate the model and then explores various counterfactual

[3] Inelastic supply has been posited as an explanation for understanding low cost-price pass-through in the literature, in addition to the mechanisms of price rigidities and vertical restraints (Bonnet et al., 2013).

[4] Feenstra's (1994) methodology has been used in several studies, including Broda and Weinstein (2006), Chen and Novy (2011), Imbs and Mejean (2015) and von Cramon-Taubadel et al. (2016).

scenarios to make inferences on the pattern of trade and production. In contrast, our study uses historical panel data on production to explore the effectiveness of trade to adapt to short term (year-to-year) production variations. Another related study by Reimer and Li (2010) estimate and simulate a Ricardian-based model of the world crop market. They find that trade in crops is significantly lower than what it would be in a frictionless world. They also find that distance limits the extent by which changes in one country are transmitted to others. Our finding that trade responds inelastically to short-term fluctuations in produced quantities is complementary to their findings.

Our work is complementary to studies that emphasize the role of production in other countries and trade as a response to production shocks. Lybbert et al. (2014), for example, analyze whether the staggered growing seasons for wheat and soybeans in the southern and northern hemispheres helps to deal with production shocks. In particular, since the growing seasons are displaced by about a half year between the northern and southern hemisphere, the farmers on one of the hemispheres have time to react to production shocks on the other hemisphere and partially smooth global production. Dingle et al. (2016) consider the role of trade costs in limiting the role trade as an efficient means of coping with productivity shocks. Using the phenomena of El Niño and La Niña they find that geographical correlation of shocks combined with trade costs increasing in geographical distance pose a problem for trade as a mechanism for coping with shocks. Jones and Olken (2010) investigate the effects of weather on exports, and find that in poor countries, higher temperature is associated with lower exports, especially for agriculture and light manufacturing. Costinot et al. (2016) find that adapting to future climate scenarios through changes in growing patterns is more important than changing the trade patterns. Our results reinforce the Costinot et al. (2016) prediction that international trade in agricultural products plays a relatively small role in adapting to future climate shocks.

Conceptual Framework

We set up a gravity model of trade based on CES utility with consumer preference assumptions that follows the seminal work of Anderson (1979). Consider a trade economy where there is a set $\bar{J}$ of different countries. The income in country j is Y_j. There is a set $\bar{G}$ of different categories of goods. Within each category of goods, different types of goods are distinguished by country of origin. Let the produced quantity of good $g \in \bar{G}$ in country $i \in \bar{J}$ be X_{gi}. Let c_{gij} denote consumption in country j of the good of type g produced in country i and let the price of this good, in country j, be p_{gij}. We assume a nested structure for the preferences over consumption of the different goods. In particular, the utility of the representative household in country j from consuming the basket $(c_{gij})_{g\in\bar{G},i\in\bar{J}}$ is

$$U\left((c_{gij})_{g,i}\right) = \prod_{g\in\bar{G}} C_{gj}^{\alpha_g},$$

where

$$\sum_{g\in\bar{G}} \alpha_g = 1 \tag{1}$$

and

$$C_{gj} \equiv \left[\sum_{i\in\bar{J}} c_{gij}^{\frac{\sigma-1}{\sigma}}\right]^{\frac{\sigma}{\sigma-1}}. \tag{2}$$

Hence, the utility-maximization problem facing the representative household in country j is

$$\max_{(c_{gij})_{g,i}} U\left((c_{gij})_{g,i}\right) \text{ s.t. } \sum_{i\in\bar{J},\, g\in\bar{G}} p_{gij}c_{gij} \leq Y_j. \tag{3}$$

Solving this problem yields that a the share of income spent on goods of type g is $\alpha_g Y_j$.[5] This implies that, as long as we disregard potential changes in overall income Y_j, we can treat the different types of goods (where the categories are denoted by g) separately and we drop the subscript g from the notation. The demand in country j for the good (of type g) from country i, c_{ij}, is given by

$$c_{ij} = \frac{\alpha Y_j p_{ij}^{-\sigma}}{P_j^{1-\sigma}}, \tag{4}$$

where

$$P_j \equiv \left[\sum_{i\in\bar{J}} p_{ij}^{1-\sigma}\right]^{\frac{1}{1-\sigma}} \tag{5}$$

is the price index in country j (for goods in the considered category g).

There are two types of trade costs: proportional (iceberg) τ_{ij} and per unit t_{ij}. The internal trade cost in country i is normalized by setting $t_{ii} = 0$ and $\tau_{ii} = 1$. The price in country j is then

$$p_{ij} = \tau_{ij}p_{ii} + t_{ij} \tag{6}$$

[5] See appendix A for the derivations.

and the price index (5) becomes

$$P_j = \left[\sum_{i \in \bar{J}} (\tau_{ij} p_{ii} + t_{ij})^{1-\sigma} \right]^{\frac{1}{1-\sigma}}. \tag{7}$$

We now turn to the supply-side of the model. Much of the international trade literature imposes constant marginal costs and markup pricing (Helpman and Krugman, 1985) based on equation (4). Although the constant marginal cost assumption is relevant in some contexts, we argue that it does not capture the nature of short-run supply for agricultural commodities. Our framework differs by assuming a perfectly inelastic supply function. As discussed above, this is intended to capture that short run fluctuations in yield give rise to exogenous variation of produced quantities. In the model, this is done by imposing market-clearing conditions.

If X_i is the produced amount of the good in country i then the market-clearing condition is $X_i = \sum_{j \in \bar{J}} \tau_{ij} c_{ij}$. Using (4) gives[6]

$$X_i = \sum_{j \in \bar{J}} \frac{\alpha Y_j \tau_{ij} p_{ij}^{-\sigma}}{P_j^{1-\sigma}}. \tag{8}$$

This yields a system with one equation per country producing the good in the prices p_{ii} (also one per country producing the good). The incomes, Y_j, and the trade costs τ_{ij} and t_{ij} are constant parameters. We are interested in seeing how changes in the produced quantities X_i affect the equilibrium. We can do this by assuming exogenous changes in the produced quantities and derive the endogenous responses of prices and traded quantities. In general this is difficult and we will first consider a special case.

Equilibrium Without Trade Costs

Without trade costs, $\tau_{ij} = 1$ and $t_{ij} = 0$ for all i, j, and the price index (7) is independent of j so that we can write $P_j = P$ for all i. Furthermore, the price of variety i is the same in all countries j so that $p_{ij} = p_{ii}$ for all j. The market-clearing

[6]Note that we here assume that iceberg trade costs entail physical losses of the shipped product (which is the conventional way of treating them). For the per-unit costs, we do not assume any physical losses. This simplifies the analysis but does not affect the results.

conditions (8) then imply

$$X_i = \sum_{j \in \bar{J}} \frac{\alpha Y_j p_{ii}^{-\sigma}}{P^{1-\sigma}} = \frac{\alpha p_{ii}^{-\sigma} \sum_{j \in \bar{J}} Y_j}{P^{1-\sigma}}. \tag{9}$$

Solving for the price p_{ii} relative to the price index P gives

$$\frac{p_{ii}}{P} = \left(\frac{\alpha \sum_{j \in \bar{J}} Y_j}{P X_i} \right)^{\frac{1}{\sigma}}. \tag{10}$$

Equation (9) can also be solved for $p_{ii}^{-\sigma}/P^{1-\sigma}$. Substituting the resulting expression in (4) yields a reduced-form expression for trade as a function of income and production:

$$c_{ij} = \frac{Y_j}{\sum_{j' \in \bar{J}} Y_{j'}} X_i. \tag{11}$$

The representative household in each country thus consumes its income share of the total endowment of each good.

On the left-hand sides of our regressions, we will be using either trade flows or prices at which trade is made. In order to eliminate the unobserved price index P we employ an odds ratio gravity specification, where we estimate the ratio of trade between country i and j and between reference exporter country k and j. Based on (10) and (11), the model without trade costs yields the following equations corresponding to such regression equations:

$$\log\left(\frac{c_{ij}}{c_{kj}}\right) = \log(X_i) - \log(X_k) \tag{12}$$

$$\log\left(\frac{p_{ij}}{p_{kj}}\right) = -\frac{1}{\sigma}\log(X_i) + \frac{1}{\sigma}\log(X_k). \tag{13}$$

The theory predicts that the coefficients on $\log(X_i)$ and $\log(X_k)$ will, in absolute terms, be equal to 1 in the quantity regressions and $1/\sigma$ in the price regressions. This should be compared to the predictions of a model with constant unit costs where the coefficients instead would have been σ and 1 respectively.

Adding Trade Costs

In this section we consider the case with trade costs (both proportional and per-unit trade costs). Solving analytically for the case with inelastic supply is then no longer possible. We can, however, derive some testable predictions. In particular, we can derive the expected relation between the coefficients in the quantity and price regressions and predictions that allow us to test for the presence of per-unit trade costs.

The demand function (4) gives

$$\frac{c_{ij}}{c_{kj}} = \left(\frac{p_{kj}}{p_{ij}}\right)^{\sigma}, \tag{14}$$

which implies that the relationship between the coefficients of the price and quantity regressions are the same as in (12) and (13).

The prices p_{kj} and p_{ij} are the prices paid in the importing country (including trade costs) and thus correspond to the CIF unit values in the data. We also have data on the FOB unit values. The differences between the two unit-value measures are the trade costs. In particular, the CIF price for the trade flow from country i to country j is p_{ij} while the FOB price is p_{ii}. Consider now the effect of a change in some production X on the prices. Using (6), the difference between the effect on the logs of the ratios of CIF and FOB prices is

$$\frac{p_{kj}}{p_{ij}}\frac{d}{dX}\frac{p_{ij}}{p_{kj}} = \frac{1}{p_{ij}}\frac{dp_{ij}}{dX} - \frac{1}{p_{kj}}\frac{dp_{kj}}{dX} = \frac{\tau_{ij}p_{ii}}{\tau_{ij}p_{ii}+t_{ij}}\frac{1}{p_{ii}}\frac{dp_{ii}}{dX} - \frac{\tau_{kj}p_{kk}}{\tau_{kj}p_{kk}+t_{kj}}\frac{1}{p_{kk}}\frac{dp_{kk}}{dX}. \tag{15}$$

If there are no per-unit trade costs ($t_{ij} = t_{kj} = 0$), then the effects on CIF and FOB prices should be the same (since the ratios in front of the derivatives are then equal to one). With strictly positive per-unit trade costs, we expect the effects on CIF unit values to be smaller than the effects on FOB prices (since the per-unit trade costs dampen the effects of price changes in percentage terms). The (relative) difference between the coefficients then measures the size of the per-unit trade costs.

In Appendix B we consider the case of three symmetric countries. We show that if trade costs are only proportional in nature then the predicted coefficient on changes in produced quantities for traded quantities is larger than one. However, if only per-unit trade costs are present then the predicted coefficient could be larger or smaller than one, with smaller than one the most likely case.

Data and Descriptive Statistics

For our empirical analysis we combine country-level data on food production and yields with data on bilateral trade flows. The country-level data on food production and yields is taken from the FAOSTAT database. Production data is reported in tonnes, while yield data is reported in tonnes per hectare. Figure 1 illustrates that the data exhibits significant variation in yield over time.

The aggregate bilateral trade unit-value data is taken from CEPII's Trade Unit Values (TUV) Database, which is available for the years 2000-2014 at the 6-digit HS product level. The database reports unit-value data in terms of "Free On Board" (FOB) and "Cost Insurance Freight" (CIF). FOB unit values reflect the price when the good leaves the exporting country, while CIF unit values reflect the price when it arrives at its destination. We use traded quantity data directly from the COMTRADE website for the same years, products and country pairs as the CEPII data. We match trade flows with exporter and importer production and yield data using FAOSTAT's concordance between its own commodity classification and 6-digit HS2007.

We are interested in studying primary agricultural production and thus focus our analysis on grains, vegetables, fruits and nuts. We do not use processed food production data in this analysis since the amount of processed food production is endogenous to trade in primary agricultural commodities. We also disregard animal-based commodities since it is difficult to interpret year-to-year variations in animal production in the same way as crop production. A complete list of products included in the analysis is provided in Appendix D. In addition, we remove trade flows where the exporting country does not produce the commodity according to FAOSTAT. There are many instances where country-pairs do not trade certain goods, which we do not include in the analysis. After all of these restrictions we are left with a maximum of 316,043 observations for the CIF unit value analysis covering 77 FAO products, 185 exporting countries and 175 importing countries for the years 2000 to 2014. Descriptive statistics are given in Table 1.

Empirical Specification

As mentioned above, we employ an odds ratio gravity specification and use, on the left-hand side, ratios of prices or quantities for trade between importing country j and exporting countries i and k, where k is a reference exporting country. Following Feenstra (1994), the reference exporter country is a country k from which country j imports the considered good in as many years as possible in the data. Normalizing the data in this way subsumes any unobserved importer-product-year fixed effects,

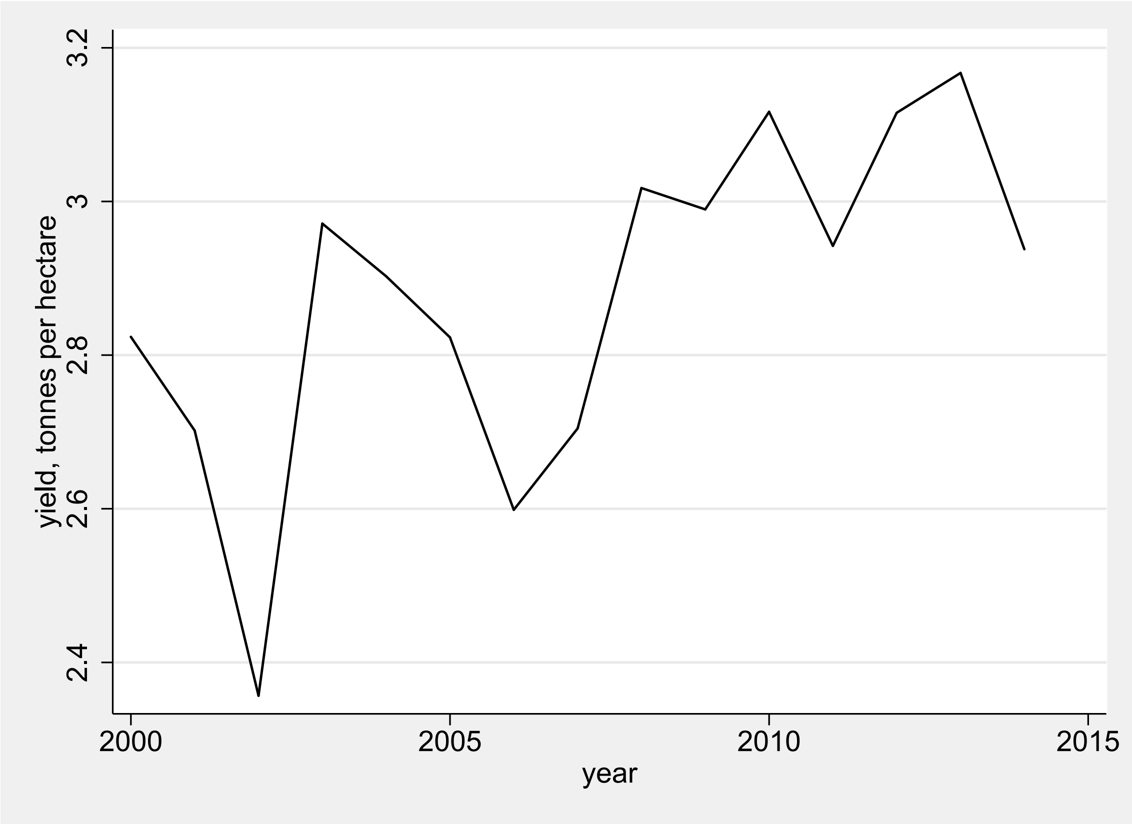

Figure 1: U.S. wheat yields, 2000-2014. Source: FAOSTAT

including the importer's price index.[7]

We first-difference both our regression equations along the time dimension since our production variables are not trend stationary.[8] First-differencing the data subsumes the panel (exporter-reference country-product) fixed effects and thus controls for all time-constant variation along the country-pair-product dimension, including bilateral distance. The regression analysis for traded quantities take the following form, following (12):

$$\begin{aligned}\Delta \ln\left(\frac{q_{ijgt}}{q_{kjgt}}\right) &= \beta_0 + \beta_1 \Delta \ln\left(X_{igt}\right) + \beta_2 \Delta \ln\left(X_{ig,t-1}\right) + \beta_3 \Delta \ln\left(X_{kgt}\right) \\ &\quad + \beta_4 \Delta \ln\left(X_{kg,t-1}\right) + \alpha_{it} + \alpha_{kt} + \epsilon_{ijgt},\end{aligned} \tag{16}$$

[7]An alternative specification would be use the importing country i as the reference country, but this approach is impeded by a lack of direct data on non-traded production in the importing country and its corresponding price. Using the importer as the reference country also imposes the restriction that only trade flows can be included if the importing country also produces the same good. Nonetheless, as a robustness check, we have repeated the regression analysis using the importing country as the reference. The results are consistent and reported in Appendix C.

[8]The Hadri LM test statistics for exporter production ($\Delta \ln(X_{igt})$) and importer production (X_{kgt}) are 49.29 and 51.25 respectively, which rejects the null hypothesis that all panels are stationary. We perform the test on 2923 time series for which there are no missing year observations.

Table 1: Descriptive statistics

Variable	Obs.	Mean	St. dev.	Min	Max
CIF unit values, USD: $p_{ijkt,CIF}$	316974	2240.6	3957.3	13.5	392350.5
Ref. country domestic price (USD): p_{kjkt}	316974	1930.6	3090.2	13.5	392350.5
Unit value ratio: $p_{ijkt,CIF}/p_{kjkt}$	316974	1.51	2.62	0.0061	138.3
Quantity traded, th tonnes: q_{ijkt}	199494	11.1	181.4	0	25009.8
Ref. country quantity traded th tonnes: q_{kjkt}	292003	12.8	135.1	0.000006	11190.5
Quantity ratio: q_{ijkt}/q_{kjkt}	283494	67.6	8410.9	0.00000002	4333544.5
Exporter production, million tonnes: X_{ikt}	316974	3.68	19.4	0.000001	361.1
Ref. country production, million tonnes: X_{ikt}	316974	4.54	21.0	0.000001	361.1
Exporter yield, tonnes/Ha	316974	20.2	96.6	0.0056	4651.5
Ref. country yield, tonnes/Ha	316974	20.2	102.9	0.013	4651.5
Exporter pop. (millions)	315218	141.7	310.2	0.049	1369.4
Ref. country pop. (millions)	316274	161.5	349.5	0.070	1369.4
Exporter GDP per capita (th USD)	315218	22.7	15.3	0.34	164.1
Ref. country GDP per capita (th USD)	316274	23.6	14.5	0.54	90.6

Based on observations from column (4) of Table 3.

where q_{ijgt}/q_{kjgt} is the quantity traded from exporter country i to importer country j of product g in year t, divided by the reference country k's trade. $\Delta\ln(X_{igt})$ and $\Delta\ln(X_{ig,t-1})$ are the percent changes in production in exporter country i of product g in year t and $t-1$ respectively, while $\Delta\ln(X_{kgt})$ and $\Delta\ln(X_{kg,t-1})$ are the percent changes in production in reference country k of product g in year t and $t-1$ respectively. Exporter-year and reference country-year fixed effects subsume the population and GDP per capita growth terms, and are denoted by α_{it} and α_{kt}. Country-year fixed effects control for any unobserved country-year variation that can explain trade flows or prices, including inflation. The combination of exporter-year and reference country-year fixed effects also controls for the effect of exchange rates on trade and prices and also controls for any changes in trade costs over time that are country-year-specific.

We include lagged production terms since many commodities are storable and experience long time lags due to transportation, implying that production the previous year can affect current trade patterns. This is especially important in the Northern Hemisphere where many crops are harvested in the fall and then exported the next calendar year. The combination of first-differencing and using lags requires that a country-pair must produce and trade a particular good for at least three years in a row in order to be included in the regression. We thus explore the product-country intensive margin of trade in this study.

First-differencing the trade unit value regression equation (13) and adding exporter-year and reference country-year fixed effects yields the regression equation for our analysis of unit values:

$$\Delta \ln \left(\frac{p_{ijgt}}{p_{kjgt}}\right) = \beta_0 + \beta_1 \Delta \ln (X_{igt}) + \beta_2 \Delta \ln (X_{ig,t-1}) + \beta_3 \Delta \ln (X_{kgt}) + \beta_4 \Delta \ln (X_{kg,t-1}) + \alpha_{it} + \alpha_{kt} + \upsilon_{ijgt}, \quad (17)$$

where p_{ijgt}/p_{kjgt} is the trade unit value from exporter country i to importer country j in product k in year t, divided by the reference country k's trade unit value.

The coefficients from these regressions reflect the impact of production shocks on changes in the ratio of import prices from countries i and k. As discussed in the conceptual framework, higher production in exporting country i should increase q_{ijgt}/q_{kjgt} and decrease p_{ijgt}/p_{kjgt}, while higher production in the reference country k should decrease q_{ijgt}/q_{kjgt} and increase p_{ijht}/p_{kjgt}.

Empirical Analysis

We present the regression results for traded quantities and trade unit values using bilateral trade data at the country level. Our variables of interest are current and lagged year-product-country-specific production.

First Stage

We instrument for contemporaneous and lagged production using yield data in order to remove any endogeneity of trade on seeded acreage. According to Roberts and Schlenker (2013), year-to-year variation in yields is mainly caused by short term fluctuations in growing conditions. Moreover, Choi and Helmberger (1993) find that yields are insensitive to commodity price changes. The first-stage results are reported in Table D.2. The results suggest that a one percent increase in yield leads to a 0.8 percent contemporaneous increase in production, with a very high t-statistic. Since our assumption is that it should be one, the first-stage coefficients are thus close to the assumed values.[9]

[9] An alternative to using yield as instrument would be to try to capture the underlying weather conditions that drive the yield variations. However, the aspects of weather outcomes that drive yields are very difficult to capture. Especially for many crops in many countries. What drives yields can often be very specific aspects of the weather. For instance, temperature minima or precipitation above or below some threshold. These outcomes often also matter particularly during specific growth phases of the crops. While much data on weather, what crops are grown where and crop calendars (planting and harvesting times) exists, it is not obvious that the data would allow for constructing strong instruments. We therefore treat yield as a sufficient statistic that captures the effects of weather, disease and any other factors that affect production.

Production Shocks and Traded Quantities

We first investigate the effect of production shocks on traded quantities in order to establish some stylized facts about trade volumes. We restrict the sample to trade flow observations for which we have corresponding data on CIF trade unit values. The results for traded quantities are presented in Table 2. In column (1) of Table 2 we present the results using first-differenced data but without any additional fixed effects. In each successive column we add more fixed effects until we arrive at our preferred specification. Year fixed effects are added in column (2), which controls for changes in any unobserved covariates that affect all countries and products, such as the global business cycle. In column (3) we add importer and exporter fixed effects, which controls for differential trends between countries caused by domestic factors such as the institutional environment that may impact traded quantities or unit values. In column (4) we include exporter-year and importer-year fixed effects, which control for any country-specific variation over time that affects traded quantities or unit values of all commodities. Country-level macroeconomic shocks would be captured by the exporter-year and importer-year fixed effects, as would any weather event that affected exports of all crops. The fixed effects used in column (4) thus allow us to exploit the variation in production between crops in the same country. The combination of first-differencing and country-year fixed effects follows the work of Baier and Bergstrand (2007) in a gravity equation context. We instrument production in the exporting countries using average yield in all specifications in order to ensure that production variation is not reflecting changes in planted acreage, which may be endogenous to commodity prices and trade flows. The individual F-statistics suggest that the yield instruments are strong.

The point estimates in Table 2 suggest that a one percent increase in exporter production leads to a 0.27 percent increase in relative trade quantities the same year and a further 0.19 percent increase the following year. For the reference country, the effect is 0.25 percent in the same year and 0.26 percent for the lag.

As a robustness check, we estimate the effect of production shocks on traded quantities by product group and country characteristics. In Table D.3 we present the subsample results separately for grains, vegetables and fruits. The point estimates suggest that production shocks in both exporter countries and reference countries affect trade in grains and fruits with high statistical significance, while the effects for vegetables are weakly significant. Comparing the results for grains and fruits we can see that if the effects over two years are summed they are of similar size but the time profiles are quite different. For grains, lagged production shocks matter as well, while in the case of fruits the contemporaneous shocks are much more important. This difference is consistent with grains being significantly more storable than fruits.[10] In Table D.4 we present the subsample results distinguishing between OECD and non-OECD member countries. Our main result holds regardless of whether the importing and/or exporting country is an OECD member.

Based on these quantity regressions, we can argue that our assumption of inelastic supply seems warranted. The model without trade costs, Equation (12), predicts coefficients on production equal to one. The alternative model assuming perfectly elastic supply would instead imply a coefficient of $\sigma > 1$. The fact that the coefficients in Table 2 are smaller than one (summed over the two years, they are about 0.5) suggests that the relationship between exporter production and traded quantities is thus relatively inelastic. Further inspection of the data suggests that total exports are small relative to domestic production, which would imply a trade elasticity in excess of unity. We illustrate the fact that exports are relatively small relative to production in Figure 2, which ranks exporter-product-year observations by export intensity, defined as the ratio of total exports to domestic production for each product and year. This illustrates that exports are small relative to domestic production in the majority of cases in the data. Given that exports tend to be small relative to domestic production, we would expect much higher trade elasticities in a frictionless world as small percentage changes in production would lead to large percentage changes in imports or exports in absolute terms. Based on the analysis in Appendix B, the estimated coefficients suggest the presence of large per-unit trade costs. In sum, these results suggest that the forces of trade costs or consumers' ability to substitute with other commodities or processed goods diminishes the role of international trade to smooth out the year-to-year volatility in production caused by weather and other factors.

Production Shocks and Unit Values

In our unit-value regressions, we first consider CIF unit values. In Equation (14), the prices on the right-hand side include trade costs and thus correspond to CIF unit values. Hence, using CIF unit values allows us to back out the elasticity of substitution parameter σ by comparing the coefficients of the quantity and price

[10] Note, however that carryover stocks are relatively small compared to production even for storable commodities. In the case of wheat, for example, world ending stocks were approximately one quarter of total world production during the 1997-2016 period. (USDA, 2017)

Table 2: Production shocks and import quantities

	(1)	(2)	(3)	(4)
Exporter production:	0.318***	0.319***	0.318***	0.273***
$\Delta \ln(X_{git})$	(0.0382)	(0.0385)	(0.0384)	(0.0386)
Lagged exporter production:	0.242***	0.239***	0.238***	0.190***
$\Delta \ln(X_{gi,t-1})$	(0.0317)	(0.0319)	(0.0320)	(0.0323)
Ref. country production:	-0.269***	-0.265***	-0.269***	-0.250***
$\Delta \ln(X_{gkt})$	(0.0332)	(0.0326)	(0.0328)	(0.0339)
Lagged ref. production:	-0.262***	-0.258***	-0.261***	-0.259***
$\Delta \ln(X_{gk,t-1})$	(0.0289)	(0.0284)	(0.0289)	(0.0272)
$\Delta \ln(pop_{i,t})$	-0.600	-0.551	-0.404	
	(0.531)	(0.491)	(0.492)	
$\Delta \ln(pop_{k,t})$	-0.358	-0.179	1.220	
	(0.406)	(0.413)	(0.834)	
$\Delta \ln(gdppc_{i,t})$	-0.226*	-0.176	-0.181	
	(0.122)	(0.115)	(0.116)	
$\Delta \ln(gdppc_{k,t})$	0.267***	0.354***	0.395***	
	(0.0961)	(0.0881)	(0.0971)	
$\Delta \ln(gdppc_{i,t-1})$	-0.122	-0.0958	-0.0937	
	(0.0875)	(0.0867)	(0.0853)	
$\Delta \ln(gdppc_{k,t-1})$	0.0323	0.0620	0.0854	
	(0.0995)	(0.0926)	(0.0923)	
Fixed effects:		year	year exporter ref. country	exp.*year ref.*year
Kleibergen-Paap rk LM stat	48	48	48	44
Kleibergen-Paap rk Wald F stat:	98	96	96	95
Observations	208,888	208,888	208,888	275,888
R-squared	0.003	0.003	0.006	0.040

Notes: Dependent variable is first-differenced ratio of log quantity exported from country i to country j relative to reference country k, using importer-reported values.

A constant term is included, but not reported, in all specifications

Underidentification is rejected by the Kleibergen-Paap rk LM statistic in all specifications.

Robust standard errors in parenthesis, clustered by exporter country. *** p<0.01, ** p<0.05, * p<0.1

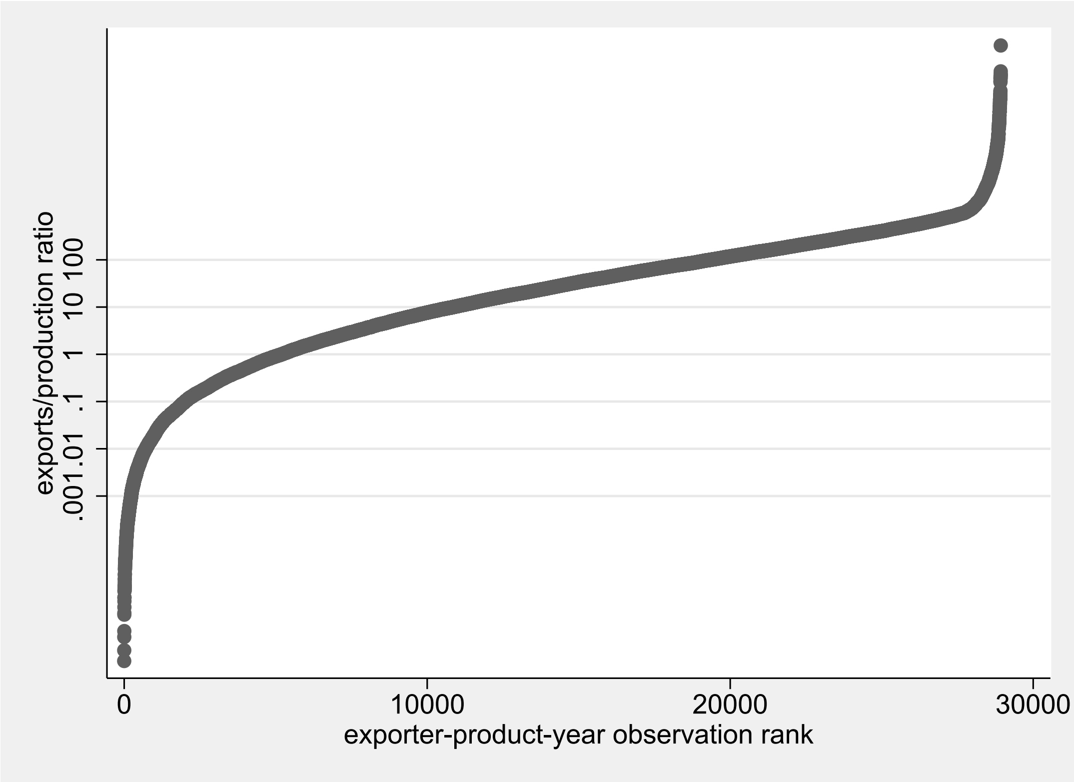

Figure 2: Distribution of export intensities. Notes: Based on observations from column (4) of Tables 3. Source: FAOSTAT

regressions. In the next section we consider the differences between coefficients from regressions using CIF and FOB unit values.

The results describing the impact of production shocks on CIF relative unit values are given in Table 3. The dependent variable is first-differenced log of ratios of unit values for exports from country i to country j to unit values of exports from country k to country j, using importer-reported (CIF) values.

The results indicate that changes in production among exporters influence unit values, with point estimates that are statistically significant at the 1 percent level for both contemporaneous and lagged production in the exporter country i. All production coefficients are statistically different from zero at the 1 percent level and have the expected signs. The point estimates in column (4) of Table 3 suggest that a one percent increase in production in country i decreases the ratio by 0.043 percent in the same year and by 0.030 percent for the lag. For the reference country, the effect is 0.027 percent in the same year and 0.019 for the lag. These point estimates are an order of magnitude smaller compared to the quantity regressions, which lends support to our theoretical model.

As a robustness check, we estimate the effect of production shocks on CIF unit

Table 3: Production shocks and (CIF) unit values

	(1)	(2)	(3)	(4)
Exporter production:	-0.0513***	-0.0497***	-0.0495***	-0.0430***
$\Delta\ln(X_{git})$	(0.00933)	(0.00904)	(0.00907)	(0.00843)
Lagged exporter production:	-0.0445***	-0.0437***	-0.0437***	-0.0298***
$\Delta\ln(X_{gi,t-1})$	(0.00790)	(0.00779)	(0.00787)	(0.00786)
Ref. country production:	0.0299***	0.0293***	0.0280***	0.0269***
$\Delta\ln(X_{gkt})$	(0.00797)	(0.00783)	(0.00799)	(0.00752)
Lagged ref. production:	0.0275***	0.0275***	0.0274***	0.0187***
$\Delta\ln(X_{gk,t-1})$	(0.00744)	(0.00738)	(0.00757)	(0.00679)
$\Delta\ln(pop_{i,t})$	0.0144	0.00537	0.0106	
	(0.121)	(0.118)	(0.131)	
$\Delta\ln(pop_{k,t})$	0.0765	0.0484	-0.669***	
	(0.132)	(0.132)	(0.229)	
$\Delta\ln(gdppc_{i,t})$	0.100***	0.0820**	0.0811**	
	(0.0330)	(0.0320)	(0.0323)	
$\Delta\ln(gdppc_{k,t})$	-0.0695**	-0.102***	-0.0846***	
	(0.0333)	(0.0287)	(0.0305)	
$\Delta\ln(gdppc_{i,t-1})$	-0.0237	-0.0181	-0.0183	
	(0.0242)	(0.0250)	(0.0255)	
$\Delta\ln(gdppc_{k,t-1})$	0.00383	0.0215	0.0438	
	(0.0287)	(0.0276)	(0.0288)	
Fixed effects:			year	exp.*year
		year	exporter	ref.*year
			ref. country	
Kleibergen-Paap rk LM stat	48	48	48	45
Kleibergen-Paap rk Wald F stat:	152	150	150	141
Observations	298,753	298,753	298,753	316,043
R-squared	0.001	0.001	0.002	0.036

Notes: Dependent variable is first-differenced ratio of log unit values for exports from country i to country j relative to reference country k, using importer-reported (CIF) values.

A constant term is included, but not reported, in all specifications

Robust standard errors in parenthesis, clustered by exporter country.

Underidentification is rejected by the Kleibergen-Paap rk LM statistic in all specifications.

*** p<0.01, ** p<0.05, * p<0.1

values by product group and country characteristics. The results are presented in the Appendix. In Table D.5 we present the subsample results for grains, vegetables and fruits. Exporter production negatively affects the relative unit values of grains and fruits, but not vegetables. As a further robustness check we present the subsample results distinguishing between OECD and non-OECD member countries in Table D.6. The point estimates suggest that the link between exporter production and relative unit values is insensitive to country characteristics, with the exception of an insignificant point estimate for the lag when the exporting nation is an OECD member in column 3.

Equation (14) from the theory implies that the elasticity of substitution can be calculated by taking the ratio between the effects in the CIF unit-value and quantity regressions. By summing the contemporaneous and lagged coefficients from Table 2 and Table 3 and computing the relevant ratios, we find that these imply an elasticity of substitution is about 10 in both cases. This is a high value, but seems reasonable given that we are computing the elasticity of substitution for relatively homogeneous agricultural commodities differentiated only by country of origin.

Estimating per-unit trade costs

We now exploit the differences in the estimates of effects of production changes on CIF and FOB unit values in order to compute the size of per-unit trade costs. As shown in Equation (15), if the effects were the same for both types of unit costs, this would indicate that per-unit trade costs were insignificant. Table 4 presents the estimates of the sensitivity of trade unit values to production using both CIF and FOB unit values. In the table we have restricted the sample to observations where both types of unit values are available in order to allow for comparison.

Let β_{CIF} and β_{FOB} denote the coefficients from the CIF end FOB regressions respectively. Based on Equation (15), the per-unit trade cost's share of import prices can be inferred from the estimates using the following formula:

$$\frac{\beta_{CIF}}{\beta_{FOB}} = \frac{\tau_{ij} p_{ii}}{\tau_{ij} p_{ii} + t_{ij}} \Rightarrow \frac{t_{ij}}{\tau_{ij} p_{ii} + t_{ij}} = 1 - \frac{\beta_{CIF}}{\beta_{FOB}}.$$

Comparing the columns of Table 4, we find that the per-unit trade costs in relation to the CIF unit value is between 14 and 29 percent.[11] This suggests that per-

[11] For the contemporaneous and lagged exporter point estimates we calculate 29 and 14 percent respectively.

Table 4: CIF versus FOB estimates

Dep. var:	(1) CIF	(2) FOB
Exporter production: $\Delta \ln(X_{git})$	-0.0505***	-0.0710***
	(0.0112)	(0.0107)
Lagged exporter production: $\Delta \ln(X_{gi,t-1})$	-0.0353***	-0.0411***
	(0.0109)	(0.0154)
Ref. country production: $\Delta \ln(X_{gkt})$	0.0226*	0.0742***
	(0.0126)	(0.0105)
Lagged ref. production: $\Delta \ln(X_{gk,t-1})$	0.0175*	0.0416***
	(0.0105)	(0.0108)
Kleibergen-Paap rk LM stat	34	34
Kleibergen-Paap rk Wald F stat:	99	99
Observations	142,768	142,768
R-squared	0.042	0.078

Notes: Dependent variable is first-differenced ratio of log unit values for exports from country i to country j relative to reference country k, using CIF trade unit values in column (1) and FOB trade unit values in column (2)

Robust standard errors in parenthesis, clustered by exporter country.

Exporter*year and ref. country*year fixed effects included in all specifications.

A constant term is included, but not reported, in all specifications

Underidentification is rejected by the Kleibergen-Paap rk LM statistic in all specifications. *** p<0.01, ** p<0.05, * p<0.1

unit trade costs here are quantitatively larger compared to manufacturing, where Irarrazabal et al. (2015) find that per-unit trade barriers are on average 14 percent of the median price. These results also underscores the importance of using CIF unit values to estimate the elasticity of substitution since demand depends on prices including trade costs.

Conclusion

The purpose of this study is to measure how agricultural trade responds to agricultural production shocks. We find that the unit values of trade flows vary systematically with production shocks using aggregate data on a large sample of countries and internationally traded agricultural commodities. Using an odds ratio gravity model we find that traded quantities and trade unit values respond to production shocks in exporting countries in a way that is consistent with product differentiation by country of origin. Traded quantities are more elastic than trade unit values, and the difference between the CIF and FOB regression estimates suggests the presence of large per-unit trade costs. These results support the predictions of our theory, which assumes perfectly inelastic supply and per-unit trade costs.

The fact that the elasticity of traded quantities with respect to production is below 1 suggests that trade is relatively insensitive to changes in exporter production. This insensitivity could be caused by trade costs that introduces frictions to the shock transmission process. Overall, the results of the article suggest that trade frictions or substitution with other goods diminishes the role of international trade as way of coping with production volatility.

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A Deriving the demand functions

The Lagrangian associated with maximization problem (3) is

$$\mathcal{L} = \prod_{g\in\bar{G}} C_{gj}^{\alpha_g} + \sum_g \mu_{gj} \left[\left[\sum_{i\in\bar{J}} c_{gij}^{\frac{\sigma-1}{\sigma}} \right]^{\frac{\sigma}{\sigma-1}} - C_{gj} \right] + \lambda_j \left[Y_j - \sum_{gi} p_{gij} c_{gij} \right].$$

The first-order conditions are

$$C_{gj} : \mu_{gj} = \alpha_g \frac{U_j}{C_{gj}} \tag{18}$$

$$c_{gij} : \lambda_j p_{gij} = \mu_{gj} \left(\frac{C_{gj}}{c_{gij}} \right)^{\frac{1}{\sigma}} = \{(18)\} = \alpha_g \frac{U_j}{C_{gj}} \left(\frac{C_{gj}}{c_{gij}} \right)^{\frac{1}{\sigma}}, \tag{19}$$

where $U_j = U\left(\{c_{gij}\}_{g,i}\right)$. Total spending on goods of type g is

$$\sum_{i\in\bar{J}} p_{gij} c_{gij} = \frac{1}{\lambda_j} \alpha_g \frac{U_j}{C_{gj}^{\frac{\sigma-1}{\sigma}}} \sum_{i\in\bar{J}} c_{gij}^{\frac{\sigma-1}{\sigma}} = \{(2)\} = \frac{1}{\lambda_j} \alpha_g U_j. \tag{20}$$

Total overall spending is

$$Y_i = \sum_{g\in\bar{G}} \sum_{i\in\bar{J}} p_{gij} c_{gij} = \frac{1}{\lambda_j} U_j \sum_{g\in\bar{G}} \alpha_g = \{(1)\} \frac{1}{\lambda_i} U_j. \tag{21}$$

Substituting this in (20) yields

$$\sum_{i\in\bar{J}} p_{gij} c_{gij} = \alpha_g Y_j. \tag{22}$$

Total spending on goods of type g is thus a given share α_g of total income Y_j. In our empirical analysis we will focus on variations of production of about 10 percent for one type of products. The overall effect of this variation on the income of the entire economy should be relatively small and we will assume that it is zero. Equation (22) then implies that we can treat different types of goods separately. In the continued analysis we drop the subscript g from the notation.

Combining (19) and (21) gives

$$p_{ij} = \frac{\alpha Y_j}{C_j^{\frac{\sigma-1}{\sigma}} c_{ij}^{\frac{1}{\sigma}}}.$$

This can also be written as

$$c_{ij} = C_j^{1-\sigma} \left(\frac{\alpha Y}{p_{ij}} \right)^{\sigma}. \tag{23}$$

Using (22), we obtain

$$\alpha Y_j = C_j^{1-\sigma} (\alpha Y_j)^{\sigma} \sum_{i \in \bar{J}} p_{ij}^{1-\sigma} \Rightarrow \alpha Y_j = C_j \left[\sum_{i \in \bar{J}} p_{ij}^{1-\sigma} \right]^{\frac{1}{1-\sigma}}.$$

The last factor is the price index defined in (5) and substitutting for this we arrive at

$$\alpha Y_j = P_j C_j.$$

Solving for C_j and substituting this in (23) delivers (4).

B The case of three symmetric countries

Consider an exogenous change Δ that affects production. Writing the market-clearing condition (8) in the form

$$X_i = \sum_{j \in \mathbb{J}} \tau_{ij} c_{ij} \tag{24}$$

and differentiating with respect to Δ gives

$$\frac{1}{X_i} \frac{dX_i}{d\Delta} = \sum_{j \in \mathbb{J}} \frac{\tau_{ij} c_{ij}}{X_i} \frac{1}{c_{ij}} \frac{dc_{ij}}{d\Delta}. \tag{25}$$

Differentiating (4) and (5) with respect to Δ yields

$$\frac{dc_{ij}}{d\Delta} = -c_{ij} \left[(1-\sigma) \frac{1}{P_j} \frac{dP_j}{d\Delta} + \sigma \frac{1}{p_{ij}} \frac{dp_{ij}}{d\Delta} \right] \text{ and } \frac{1}{P_j} \frac{dP_j}{d\Delta} = \sum_{i \in \mathbb{J}} \left(\frac{p_{ij}}{P_j} \right)^{1-\sigma} \frac{1}{p_{ij}} \frac{dp_{ij}}{d\Delta}.$$

Substituting these in (25) and rewriting we arrive at

$$\begin{aligned} \frac{1}{X_i} \frac{dX_i}{d\Delta} &= -\left[\sum_{j \in \mathbb{J}} \frac{\tau_{ij} c_{ij}}{X_i} \left((1-\sigma) \left(\frac{p_{ij}}{P_j} \right)^{1-\sigma} + \sigma \right) \right] \frac{1}{p_{ij}} \frac{dp_{ij}}{d\Delta} \\ &\quad - \sum_{i' \in \mathbb{J} \setminus \{i\}} \left[\sum_{j \in \mathbb{J}} (1-\sigma) \frac{\tau_{ij} c_{ij}}{X_i} \left(\frac{p_{i'j}}{P_j} \right)^{1-\sigma} \right] \frac{1}{p_{i'j}} \frac{dp_{i'j}}{d\Delta}. \end{aligned}$$

Using that Equation (6) implies

$$\frac{1}{p_{ij}}\frac{dp_{ij}}{d\Delta} = \frac{\tau_{ij}p_{ii}}{\tau_{ij}p_{ii}+t_{ij}}\frac{1}{p_{ii}}\frac{dp_{ii}}{d\Delta}$$

we get

$$\frac{1}{X_i}\frac{dX_i}{d\Delta} = -A_{ii}\frac{1}{p_{ii}}\frac{dp_{ii}}{d\Delta} - \sum_{i'\in\mathbb{J}\setminus\{i\}} A_{ii'}\frac{1}{p_{i'i'}}\frac{dp_{i'i'}}{d\Delta}, \tag{26}$$

where

$$A_{ii'} \equiv \begin{cases} \sum_{j\in\mathbb{J}} \frac{\tau_{ij}c_{ij}}{X_i}\left((1-\sigma)\left(\frac{\tau_{ij}p_{ii}+t_{ij}}{P_i}\right)^{1-\sigma}+\sigma\right)\frac{\tau_{ij}p_{ii}}{\tau_{ij}p_{ii}+t_{ij}} & \text{if } i'=i \\ \sum_{j\in\mathbb{J}}(1-\sigma)\frac{\tau_{ij}c_{ij}}{X_i}\left(\frac{\tau_{i'j}p_{i'i'}+t_{i'j}}{P_j}\right)^{1-\sigma}\frac{\tau_{i'j}p_{i'i'}}{\tau_{i'j}p_{i'i'}+t_{i'j}} & \text{if } i'\neq i \end{cases}. \tag{27}$$

Consider now the case where there are three countries. The first country represents the importer, the second country the exporter and the third represents the reference country. Since our regression equations consider the effects of production in the exporter and reference country, we here assume that production in country 1 is unaffected by Δ. Setting $\frac{1}{X_1}\frac{dX_1}{d\Delta} = 0$, Equation (26) for $i = 1, 2, 3$ gives

$$\begin{aligned} \frac{1}{X_2}\frac{dX_2}{d\Delta} - \frac{1}{X_3}\frac{dX_3}{d\Delta} &= \left[\frac{A_{21}-A_{31}}{A_{11}}A_{12}+A_{32}-A_{22}\right]\frac{1}{p_{22}}\frac{dp_{22}}{d\Delta} \\ &+ \left[\frac{A_{21}-A_{31}}{A_{11}}A_{13}+A_{33}-A_{23}\right]\frac{1}{p_{33}}\frac{dp_{33}}{d\Delta}. \end{aligned} \tag{28}$$

We will now evaluate these expressions for three symmetric countries and thus assume

$$X \equiv X_i \;\forall i,\; Y \equiv Y_j \;\forall j, \tau_{ij} = \begin{cases} 1 & \text{if } i=j \\ \tau & \text{if } i\neq j \end{cases} \text{ and } t_{ij} = \begin{cases} 0 & \text{if } i=j \\ t & \text{if } i\neq j \end{cases}.$$

This implies that all prices are the same, $p_{ii} = p$ for all i and the price index (5) becomes

$$P_i = P \equiv \left[p^{1-\sigma} + 2\left(\tau p + t\right)^{1-\sigma}\right]^{\frac{1}{1-\sigma}}.$$

Furthermore, combining (4) and (24) gives

$$c_{ij} = \begin{cases} \frac{p^{-\sigma}}{p^{-\sigma}+2\tau(\tau p+t)^{-\sigma}}X \equiv c & \text{if } i=j \\ \frac{(\tau p+t)^{-\sigma}}{p^{-\sigma}+2\tau(\tau p+t)^{-\sigma}}X \equiv \tilde{c} & \text{if } i\neq j \end{cases}.$$

All this combined implies that

$$A_{ii} = A \equiv \frac{c}{X}\left((1-\sigma)\left(\frac{p}{P}\right)^{1-\sigma}+\sigma\right)+2\frac{\tau\tilde{c}}{X}\left((1-\sigma)\left(\frac{\tau p+t}{P}\right)^{1-\sigma}+\sigma\right)\frac{\tau p}{\tau p+t}$$

$$A_{ii'} = \tilde{A} \equiv (1-\sigma)\left[\left(\frac{\tau\tilde{c}}{X}+\frac{c}{X}\right)\left(\frac{\tau p+t}{P}\right)^{1-\sigma}\frac{\tau p}{\tau p+t}+\frac{\tau\tilde{c}}{X}\left(\frac{p}{P}\right)^{1-\sigma}\right],$$

for all i and $i' \neq i$.

Substituting this in (28) yields

$$\frac{1}{p_{22}}\frac{dp_{22}}{d\Delta}-\frac{1}{p_{33}}\frac{dp_{33}}{d\Delta}=-\frac{1}{A-\tilde{A}}\left(\frac{1}{X_2}\frac{dX_2}{d\Delta}-\frac{1}{X_3}\frac{dX_3}{d\Delta}\right).$$

Combining this with (15) and (14) we arrive at

$$\frac{1}{c_{12}}\frac{dc_{12}}{d\Delta}-\frac{1}{c_{13}}\frac{dc_{13}}{d\Delta}=\sigma\frac{\frac{\tau p}{\tau p+t}}{A-\tilde{A}}\left(\frac{1}{X_2}\frac{dX_2}{d\Delta}-\frac{1}{X_3}\frac{dX_3}{d\Delta}\right).$$

We can compute the predicted coefficient for the quantity regressions as

$$\beta \equiv \sigma\frac{\frac{\tau p}{\tau p+t}}{A-\tilde{A}} = \frac{1}{1+\frac{t}{\tau p}\frac{p^{-\sigma}}{p^{-\sigma}+2\tau(\tau p+t)^{-\sigma}}+\frac{1-\sigma}{\sigma}\frac{(p^{-\sigma}-\tau(\tau p+t)^{-\sigma})^2}{(p^{-\sigma}+2\tau(\tau p+t)^{-\sigma})(p^{1-\sigma}+2(\tau p+t)^{1-\sigma})}}.$$

Without trade costs, we get $\beta = 1$ as expected. With only proportional trade costs, $t = 0$, we get

$$\beta = \frac{1}{1+\frac{1-\sigma}{\sigma}\frac{(p^{-\sigma}-\tau(\tau p)^{-\sigma})^2}{(p^{-\sigma}+2\tau(\tau p)^{-\sigma})(p^{1-\sigma}+2(\tau p)^{1-\sigma})}} > 1.$$

With only per-unit trade costs, $\tau = 1$, we get

$$\beta = \frac{1}{1+\frac{t}{p}\frac{p^{-\sigma}}{p^{-\sigma}+2(p+t)^{-\sigma}}+\frac{1-\sigma}{\sigma}\frac{(p^{-\sigma}-(p+t)^{-\sigma})^2}{(p^{-\sigma}+2(p+t)^{-\sigma})(p^{1-\sigma}+2(p+t)^{1-\sigma})}}.$$

While this can, in general, be smaller or larger than one, it can be shown numerically that, for relevant values of t it will typically be smaller than one.

C Normalization Using Importing Country

As a robustness check we present the results when we use the non-traded quantity and price of the importing country's domestically produced variety for normalization. Instead of dividing by a reference exporter, c_{ij}/c_{kj} and p_{ij}/p_{kj}, on the left-hand side we divide by the importing countries domestic consumption and domestic price, c_{ij}/c_{jj} and p_{ij}/p_{jj} respectively. We construct c_{jj} using domestic production minus total reported exports. For p_{jj} we use the domestic price of the good reported by FAOSTAT. This measure of quantity is potentially problematic since it relies on the data capturing all exports of the good. The price measure is also problematic since it is the average domestic price paid for the good in the importing country, which is not necessarily the domestically produced variety of the good. As instruments for production we now use yield in the importing and exporting countries. The results for the quantity regression is given in Table C.1. The results for the price regressions are given in Table C.2. We can see that the results for the exporting country are similar to those in the baseline regressions (in Table 2). The effect of production in the exporting country is larger than one. This suggests that the effect of domestic production in the importing country is larger than that of production in the exporting countries, which is could be driven by the fact that trade costs are lower for domestic trade compared to international trade. Similarly, looking at the effects in the price regressions, the point estimate for production in the exporting country is similar to those in the baseline regression (Table 3) while the effects of domestic production are larger. Again, this is consistent with the existance of larger trade costs for internationally traded goods compared to domestically traded goods. In fact, the effects of domestic production in Table C.2 are more similar to the effects on FOB prices in Table 4 which is what we should expect.

Table C.1: Traded quantities, importer reference country

	(1)	(2)	(3)	(4)
Exporter production: $\Delta \ln(X_{ikt})$	0.339***	0.341***	0.341***	0.293***
	(0.0468)	(0.0456)	(0.0459)	(0.0466)
Lagged exporter production:	0.167***	0.172***	0.169***	0.126***
$\Delta \ln(X_{ik,t-1})$	(0.0315)	(0.0325)	(0.0324)	(0.0340)
Importer production: $\Delta \ln(X_{jkt})$	-1.354***	-1.346***	-1.351***	-1.348***
	(0.0537)	(0.0537)	(0.0539)	(0.0538)
Lagged importer production:	-0.0611	-0.0526	-0.0579	-0.0540
$\Delta \ln(X_{jk,t-1})$	(0.0441)	(0.0439)	(0.0437)	(0.0382)
$\Delta \ln(pop_{i,t})$	-0.0570	-0.144	-1.785	
	(0.599)	(0.621)	(1.247)	
$\Delta \ln(pop_{j,t})$	-0.833	-0.915	-1.825*	
	(0.589)	(0.598)	(0.984)	
$\Delta \ln(gdppc_{i,t})$	-0.0420	-0.189	-0.241	
	(0.160)	(0.154)	(0.165)	
$\Delta \ln(gdppc_{j,t})$	1.114***	0.888***	0.894***	
	(0.191)	(0.188)	(0.235)	
$\Delta \ln(gdppc_{i,t-1})$	0.0186	-0.0547	-0.0874	
	(0.116)	(0.110)	(0.114)	
$\Delta \ln(gdppc_{j,t-1})$	-0.0469	-0.118	-0.200	
	(0.139)	(0.163)	(0.190)	
Fixed effects:			year	exp.*year
		year	exporter	imp.*year
			importer	
Kleibergen-Paap rk	40	40	39	38
Wald F stat:	67	57	53	56
Observations	108,172	108,172	108,172	114,785
R-squared	0.058	0.058	0.062	0.118

Notes: Dependent variable is first-differenced ratio of log quantity exported from country i to country j to logged domestic production (net of exports) in country j , using importer-reported values. A constant term is included, but not reported, in all specifications Robust standard errors in parenthesis, clustered by exporter country. *** p<0.01, ** p<0.05, * p<0.1

Table C.2: CIF unit values, importer reference country

	(1)	(2)	(3)	(4)
Exporter production: $\Delta \ln(X_{ikt})$	-0.0256*	-0.0444***	-0.0442***	-0.0488***
	(0.0148)	(0.0143)	(0.0144)	(0.0137)
Lagged exporter production:	-0.0177	-0.0286***	-0.0288***	-0.0276**
$\Delta \ln(X_{ik,t-1})$	(0.0110)	(0.0102)	(0.0104)	(0.0127)
Importer production: $\Delta \ln(X_{jkt})$	0.209***	0.194***	0.192***	0.176***
	(0.0226)	(0.0222)	(0.0218)	(0.0205)
Lagged importer production:	0.0516**	0.0451**	0.0429**	0.0230
$\Delta \ln(X_{jk,t-1})$	(0.0206)	(0.0200)	(0.0199)	(0.0202)
$\Delta \ln(pop_{i,t})$	0.192	0.297	0.808**	
	(0.176)	(0.192)	(0.393)	
$\Delta \ln(pop_{j,t})$	-1.145***	-1.045***	-0.142	
	(0.331)	(0.320)	(0.586)	
$\Delta \ln(gdppc_{i,t})$	-0.0696	0.0762**	0.126***	
	(0.0606)	(0.0385)	(0.0406)	
$\Delta \ln(gdppc_{j,t})$	-0.376***	-0.226**	-0.202	
	(0.108)	(0.115)	(0.125)	
$\Delta \ln(gdppc_{i,t-1})$	0.143**	0.00525	0.0637	
	(0.0678)	(0.0472)	(0.0512)	
$\Delta \ln(gdppc_{j,t-1})$	0.197***	0.0439	0.0602	
	(0.0745)	(0.0880)	(0.0974)	
Fixed effects:			year	exp.*year
		year	exporter	imp.*year
			importer	
Kleibergen-Paap rk	44	44	43	42
Wald F stat:	119	104	111	123
Observations	111,552	111,552	111,552	117,791
R-squared	0.007	0.012	0.017	0.086

Notes: Dependent variable is first-differenced ratio of log unit values for exports from country i to country j to logged domestic prices in country j , using importer-reported (CIF) values.

A constant term is included, but not reported, in all specifications

Robust standard errors in parenthesis, clustered by exporter country.

Underidentification is rejected by the Kleibergen-Paap rk LM statistic in all specifications.

*** p<0.01, ** p<0.05, * p<0.1

D Additional tables

Table D.1: List of FAOSTAT Commodities

Almonds	Eggplants	Peas, dry
Apples	Garlic	Peas, green
Apricots	Grapefruit and pomelo	Pineapples
Asparagus	Grapes	Pistachios
Avocados	Groundnuts, in shell	Plums
Bananas	Hazelnuts (Filberts)	Poppy seed
Barley	Hops	Potatoes
Beans, dry	Kiwi fruit	Rapeseed or colza seed
Brazil nuts	Leeks and other alliaceous vegetables	Raspberries
Broad beans, Green	Lemons and limes	Rice, milled
Broad beans, dry	Lentils, dry	Rye
Buckwheat	Lettuce and chicory	Sesame seed
Cabbages	Linseed	Sorghum
Canary seed	Maize	Soybeans
Carrot	Mangoes	Spinach
Cashew nuts	Mate	Strawberries
Cassava	Melons, Cantaloupes	Dates
Cauliflowers and broccoli	Millet	Sunflower seed
Cherries	Mushrooms	Tangerines, mandarins etc.
Chestnuts	Mustard seed	Tea
Chick-peas, dry	Oats	Tomatoes, fresh
Chillies and peppers (green)	Onions, shallots (green)	Triticale
Cocoa beans	Oranges	Walnuts
Coconuts	Papayas	Watermelons
Coffee green	Peaches and nectarines	Wheat
Cucumbers and gherkins	Pears	

[1] Based on observations from column (4) of Table 3.

Table D.2: First stage results

	(1)	(2)	(3)	(4)
Dep. var:	Exporter production $\Delta\ln(X_{igt})$	Lagged exporter production $\Delta\ln(X_{gi,t-1})$	Ref. Country production $\Delta\ln(X_{gkt})$	Lagged ref. production $\Delta\ln(X_{gk,t-1})$
Exporter yield:	0.804***	-0.0614***	0.00808	0.00364
$\Delta\ln(\psi_{git})$	(0.0286)	(0.0154)	(0.00600)	(0.00438)
Lagged exporter yield:	-0.0305**	0.787***	0.00356	0.00598
$\Delta\ln(\psi_{gi,t-1})$	(0.0129)	(0.0259)	(0.00355)	(0.00539)
Ref. country yield:	0.0306***	0.00884	0.878***	-0.0498***
$\Delta\ln(\psi_{gkt})$	(0.0107)	(0.00645)	(0.0130)	(0.00654)
Lagged ref. yield:	0.00380	0.0288***	-0.0537***	0.869***
$\Delta\ln(\psi_{gk,t-1})$	(0.00481)	(0.0106)	(0.00466)	(0.0139)
Observations	316,069	316,068	316,069	316,069
R-squared	0.472	0.471	0.534	0.531

Notes: Dependent variable reported at the top of each column.

Robust standard errors in parenthesis, clustered by exporter country.

Exporter*year and ref. country*year fixed effects included in all specifications.

A constant term is included, but not reported, in all specifications.

*** p<0.01, ** p<0.05, * p<0.1

Table D.3: Production shocks and import quantities by product group

	(1) Grains	(2) Vegetables	(3) Fruits
Exporter production: $\Delta \ln(X_{git})$	0.189***	0.134*	0.426***
	(0.0479)	(0.0809)	(0.0576)
Lagged exporter production:	0.264***	0.0725	0.0729
$\Delta \ln(X_{gi,t-1})$	(0.0555)	(0.0909)	(0.0487)
Importer production: $\Delta \ln(X_{gkt})$	-0.140***	0.00581	-0.519***
	(0.0423)	(0.0836)	(0.0499)
Lagged importer production:	-0.335***	-0.166**	-0.0978**
$\Delta \ln(X_{gk,t-1})$	(0.0586)	(0.0798)	(0.0401)
Kleibergen-Paap rk LM stat	21	31	37
Kleibergen-Paap rk Wald F stat:	75	17	83
Observations	83,234	61,395	120,835
R-squared	0.079	0.088	0.085

Notes: Dependent variable is first-differenced ratio of log quantity exported from country i to country j relative to reference country k, using importer-reported values.

Exporter*year and importer*year fixed effects included in all specifications.

A constant term is included, but not reported, in all specifications

Underidentification is rejected by the Kleibergen-Paap rk LM statistic in all specifications.

*** p<0.01, ** p<0.05, * p<0.1

Table D.4: Production shocks and import quantities by country

	(1) Both OECD	(2) Neither OECD	(3) Exporter OECD	(4) Importer OECD
Exporter production:	0.266***	0.314***	0.249***	0.225***
$\Delta \ln(X_{git})$	(0.0701)	(0.0554)	(0.0887)	(0.0532)
Lagged exp. production:	0.198***	0.111**	0.329***	0.153***
$\Delta \ln(X_{gi,t-1})$	(0.0610)	(0.0493)	(0.0628)	(0.0470)
Ref. country production:	-0.310***	-0.191***	-0.188**	-0.306***
$\Delta \ln(X_{gkt})$	(0.0492)	(0.0543)	(0.0830)	(0.0661)
Lagged ref. production:	-0.263***	-0.228***	-0.345***	-0.191***
$\Delta \ln(X_{gk,t-1})$	(0.0608)	(0.0556)	(0.0635)	(0.0597)
Kleibergen-Paap rk LM stat	12	42	11	38
Kleibergen-Paap rk Wald F stat:	59	18	141	188
Observations	59,389	95,302	83,117	70,535
R-squared	0.058	0.071	0.058	0.093

Notes: Dependent variable is first-differenced ratio of log quantity exported from country i to country j relative to reference country k, using importer-reported values.
Robust standard errors in parenthesis, clustered by exporter country.
Exporter*year and ref. country*year fixed effects included in all specifications.
A constant term is included, but not reported, in all specifications
Underidentification is rejected by the Kleibergen-Paap rk LM statistic in all specifications.
*** p<0.01, ** p<0.05, * p<0.1

Table D.5: Production shocks and CIF unit values by product group

	(1) Grains	(2) Vegetables	(3) Fruits
Exporter production: $\Delta \ln(X_{git})$	-0.0225	-0.0187	-0.0650***
	(0.0149)	(0.0270)	(0.0153)
Lagged exporter production:	-0.0553***	-0.0349	0.0137
$\Delta \ln(X_{gi,t-1})$	(0.0152)	(0.0250)	(0.0120)
Ref. country production: $\Delta \ln(X_{gkt})$	0.0137	-0.00827	0.0670***
	(0.0137)	(0.0347)	(0.0112)
Lagged ref. production:	0.0280**	0.00260	-0.0235*
$\Delta \ln(X_{gk,t-1})$	(0.0129)	(0.0309)	(0.0127)
Kleibergen-Paap rk LM stat	22	27	37
Kleibergen-Paap rk Wald F stat:	79	249	90
Observations	84,745	57,367	130,388
R-squared	0.071	0.089	0.074

Notes: Dependent variable is first-differenced ratio of log unit values for exports from country i to country j relative to reference country k, using importer-reported (CIF) values.

Robust standard errors in parenthesis, clustered by exporter country.

Exporter*year and ref. country*year fixed effects included in all specifications.

A constant term is included, but not reported, in all specifications

Underidentification is rejected by the Kleibergen-Paap rk LM statistic in all specifications.

*** p<0.01, ** p<0.05, * p<0.1

Table D.6: Production shocks and CIF unit values by country

	(1) Both OECD	(2) Neither OECD	(3) Exporter OECD	(4) Importer OECD
Exporter production:	-0.0646***	-0.0284**	-0.0396**	-0.0403**
$\Delta \ln(X_{git})$	(0.0160)	(0.0133)	(0.0158)	(0.0172)
Lagged exp. production:	-0.0439**	-0.0269**	-0.0143	-0.0305**
$\Delta \ln(X_{gi,t-1})$	(0.0176)	(0.0134)	(0.0165)	(0.0137)
Ref. country production:	0.0328*	0.0287	0.0179	0.0442**
$\Delta \ln(X_{gkt})$	(0.0181)	(0.0178)	(0.0138)	(0.0199)
Lagged ref. production:	0.0244	-0.000790	0.0138	0.0353**
$\Delta \ln(X_{gk,t-1})$	(0.0160)	(0.0128)	(0.0122)	(0.0173)
Kleibergen-Paap rk LM stat	12	43	10	41
Kleibergen-Paap rk Wald F stat:	44	49	63	126
Observations	58,986	102,881	79,097	75,004
R-squared	0.052	0.069	0.051	0.080

Notes: Dependent variable is first-differenced ratio of log unit values for exports from country i to country j relative to reference country k, using importer-reported (CIF) values.

Robust standard errors in parenthesis, clustered by exporter country.

Exporter*year and ref. country*year fixed effects included in all specifications.

A constant term is included, but not reported, in all specifications.

Underidentification is rejected by the Kleibergen-Paap rk LM statistic in all specifications.

*** p<0.01, ** p<0.05, * p<0.1