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# Distance and border effects on price transmission– a meta-analysis

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## Abstract

In a meta-analysis of spatial price transmission (PT) literature we aim to test for the presence of distance and border effects on price transmission. We use PT estimates for 1189 cereal market pairs extracted from 57 studies and seek to explain them by airline distance and existence of a border. The findings indicate distance and border effects on both price cointegration and price transmission. A border separating two markets reduces the probability of cointegration of price series by 23% compared with markets located in the same country. 1000 kilometers of distance reduces the probability of cointegration by 7%. The speed of price adjustment is on average 13% slower in international than in intra-national market pairs. 1000 kilometers of distance within a country on average yields 6-20% slower price adjustment. Distance effects become negligible and economically insignificant for international market pairs. Maize price pairs are less often cointegrated compared to rice prices and cointegration is most prevalent for barley. Price transmission is slowest in wheat markets. In peer reviewed studies cointegration is more prevalent and price transmission is faster. However the explanation need not be a publication bias but can also result from higher quality methodologies. Moreover, we identify a set of model specifications that significantly affect price transmission estimates. The study contributes to the literature by presenting a first meta-analysis of spatial PT literature and providing insights into distance and border effects on price transmission.

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**JEL:**

C32, L11, Q11, Q17

**Keywords:**

meta-analysis, cointegration, spatial price transmission, distance, borders, developing countries, agricultural trade, cereals, rice, maize, wheat

# 1. Introduction

After international food price spikes in 2007-2008 and 2010-2011 (FAO et al., 2011), the transmission of price changes to domestic markets has attracted a great deal of attention in agricultural economics (Conforti, 2004; Cudjoe et al., 2010; Greb et al., 2012; Minot, 2011). The field of price transmission (PT) analysis has been analyzing food price dynamics between markets or market levels for much longer. In the past three decades, a vast literature has concerned itself with the modelling of price changes and price equilibria. An recent online search with the keywords "price transmission" in the database AgEcon-Search[2] resulted in 403 hits. The abundance of empirical work has triggered a number of meta-analyses on vertical price transmission, i.e. between different market levels (Amikuzuno and Ogundari, 2013; Bakucs et al., 2013; Greb et al., 2012). So far, no one has attempted to apply a similar approach to spatial price transmission studies to draw some general conclusions on its determinants. In line with theoretical and empirical approaches on distance and border effects in the trade and price literature we seek to shed light on distance and border effects. This is the first study employing a meta-analysis of the empirical spatial price transmission literature to test whether geographic distance and borders have a systematic effect on the strength and speed of PT. The intuitive hypothesis that we wish to test is that PT is weaker and slower over longer distances and across borders.

Our meta-analysis is limited to studies on main staple cereals such as rice, wheat and maize for a number of reasons. Cereals are relatively homogeneous goods for which quality differences play a much smaller role in PT than that for, e.g. fresh fruit and vegetables. The latter two are also more perishable, while cereals can be traded widely and internationally in large volumes and over long distances. The three main staple products rice, maize and wheat account for the largest shares of food trade worldwide and are thus economically significant products. Particularly in developing countries, cereals account for a large share of agricultural value-added (Rashid, 2011). Overall this has resulted in a large literature on PT between cereal markets that are separated by a wide range of distances both with and without international borders. This provides a rich dataset with which to test the effect of distance and borders on the PT after controlling for other potential determinants.

We find evidence of statistically and economically significant distance and border effects on PT. These findings contribute to the literature by providing further insights into the influence of different components of trade costs, which are related to crossing a border or trading over longer distances, on price transmission. The rest of this paper is structured as follows. Section II describes the theory of price transmission and the theoretical link to distances and borders. Section III the methods that we use to generate and analyze our meta-dataset. Section IV presents the data and the estimation and in section V the results are shown. Section VI concludes.

# 2. Theory

The toolbox of price transmission analysis offers a number of useful analytical applications to improve our understanding of price behavior. Whether and how price changes are transmitted to markets in other locations, or to goods or up- or downstream in the marketing chain can help assess the functioning, efficiency and integration of agri-food markets. In developing countries, food prices play an important role for both producers and consumers. The agricultural sector typically accounts for a large share of employment and income in rural areas. Efficient price signals between markets influence production and trade decisions. For consumers food prices determine their access to food, which often accounts for a large share of the household budget. Consequently price shocks can have important welfare effects. For policy makers the subject of price transmission is highly relevant as it may be helpful for the assessment of political decisions such as policy reforms and market liberalization measures (Abdulai, 2006). Information about the transmission of prices can also help in

[2] http://ageconsearch.umn.edu, accessed on May 20, 2014

assessing the depth of regional integration or the functioning of trade agreements with regard to single commodities. The knowledge of market performance and efficiency is relevant in order to avoid inefficiencies and to assess the capacity of markets to absorb and buffer shocks. Physical barriers and large distances can result in high transport and transaction costs and hinder market integration. A lack of market integration can also result from political barriers in the context of trade policies and red tape and thus point to a need for corresponding policy reforms. The knowledge of price mechanisms reduces uncertainty for policy-makers and the risk of duplication of interventions in two markets (Goletti et al., 1995).

According to the law of one price, price differences for similar products in different locations are reduced by spatial arbitrage until they amount to no more than the transfer costs. Price transmission analysis helps to assess how closely prices in two markets are linked to each other in the long run. In addition, the nature and level of price reactions indicate market integration and "the extent to which markets function efficiently" (Rapsomanikis et al., 2006). Empirical PT analysis provides insights into the price dynamics between two markets: if and when prices are linked together in a long-run equilibrium, and whether and how quickly deviations or price shocks are transmitted so as to restore this equilibrium.

Many PT studies confirm that trade costs influence price transmission. The model that is currently most commonly used to study spatial price transmission, the threshold vector error correction model (TVECM, see Greb et al., 2013) includes a threshold parameter that is introduced to reflect the magnitude of trade costs between two markets. Nevertheless, only few PT studies have explicitly studied the effect of distance and borders on price transmission (Hernandez-Villafuerte, 2011; Ihle et al., 2011). Ihle et al. (2011), in a study of maize price transmission in Eastern Africa, find evidence for distance and border effects. They estimate the speed of price transmission for 85 pairs of markets and regress these speeds on the geographic distance between the markets and a set of border-specific dummy variables. While distance has a nonlinear and statistically significant effect on the speed of price transmission, border effects are heterogeneous and vary with differences in trade policies employed by the countries. Amikuzuno & Donkoh (2012) look at border effects in tomato trade between Ghana and Burkina-Faso and model trade flows with a regime-switching model that includes a cross-border regime and a domestic trade regime. They find that PT is more rapid in the domestic trade regime and attribute this to a negative border effect.

Some studies in the trade literature provide more theoretical and empirical research on distance and border effects. Empirical trade volume studies have found that borders and distance have a strong inhibiting effect on trade flows. Interprovincial trade within Canada or within the U.S. is more than twenty times higher than across the U.S.-Canadian border, controlling for distance and the size of the markets (provinces) (McCallum, 1995). Standard gravity models thus include distance as one of the explanatory variables for trade, together with other factors such as common language, history or economic development. The effects of distance and borders on price disparities between spatially separated markets also has been subject to empirical analyses. The seminal work of Engel and Rogers (1995) shows the effect of an international border and inter-market distance on differences in real consumer prices in cities on both sides along the U.S.-Canadian border. Their results indicate that a border separating two markets has the same effect as 2500 miles of distance between them. Aker et al. (2013) find borders and ethnicities play a major role for staple price disparities in Niger. Additional transaction costs associated with crossing a border increase price disparities by 17 to 26 percent.

We follow these findings and employ a trade cost approach (Anderson and Van Wincoop, 2004). It is empirically established that high trade cost are associated with lower trade volumes. Trade due to spatial arbitrage is an important driver of spatial price transmission. If trade volumes fall due to distance and borders, then we might expect spatial price transmission to become weaker as well, all other things being equal. This theoretical link leads to the research question whether high trade cost influence price transmission via reduced trade.

Trade cost are comprised of transaction costs, transport costs and policy barriers. Transport costs vary with the means of transport and include the costs of fuel/energy, losses due to the perishability of the traded product, transport time and in some settings bribery. Policy barriers include tariffs, custom procedures, compliance costs, price policies and currency effects. Transaction costs are influenced by the business infrastructure, the prevalence of red tape and risks, and marketing costs for advertising and retailing. With increasing distance transport costs rise on average. National borders lead to additional transaction costs and policy barriers for traders. Other factors being equal, both distance and borders result in less trade. Consequently, price transmission caused by physical trade is expected to become weaker over longer distances or across national borders.

# 3. Method

We would like to test whether distance and borders affect the strength and speed of spatial PT. One empirical strategy would be to estimate PT for a large sample of market pairs worldwide and regress coefficients that measure the strength and speed of this PT on variables that measure distance and the presence of a border. Greb et al. (2012) do this for 497 international and domestic price series. Ihle et al. (2011) follow the same approach for 77 maize markets in Eastern Africa. We employ a similar approach with the important difference that rather than estimating price transmission, we draw on the vast number of empirical cereal PT estimates in the literature and conduct a meta-analysis. This tool is popular in natural sciences and medicine to combine empirical evidence from different studies with small sample sizes, and it is becoming increasingly popular in (agricultural) economics (Hess & von Cramon-Taubadel, 2008). One advantage of meta-analysis in our setting is that it provides us with a much larger sample of observations, markets and countries than we could generate ourselves in a reasonable amount of time.

A second advantage of meta-analysis is closely linked to two challenges. In the literature, each modeler/author has presumably taken a very careful look at his/her data and model specification. This should result in more precise estimates and fewer misspecification errors (e.g. due to disregarded structural breaks and non-linearities) than would obtain if we applied a one-size-fits-all model to a large dataset (see Greb et al., 2012). However, this also poses a challenge because different studies are based on different methodological approaches that are not directly comparable. The empirical price transmission literature is partly driven by model improvements and modifications of existing approaches. To address the issue of methodological heterogeneity, we control for study- and model-specific effects. The second challenge is that the literature sample may be subject to publication bias. For example, papers that present unambiguous and statistically significant results may be favored by peer-review journals (Stanley, 2001, 2005). Alternatively, the literature may reflect what is referred to as the file-drawer effect. In the spatial PT setting, authors might only pursue market pairs that display evidence of PT, and disregard others, so that journals receive for review a disproportionate number of manuscripts that report evidence of PT.

A simple model based on the theoretical considerations above defines PT between two markets as a function of distance and borders, controlling for product-, model- and study specific effects:

$$Price\ transmission\ estimates = f\ (distance, border, crop, model, data\ ) \tag{1}$$

Two dependent variables proved viable in the coding process. The first is a binary variable that indicates whether the prices being analyzed are cointegrated, i.e. share an equilibrium in the long-run. Standard bivariate cointegration tests are implemented in every statistical software program, and most empirical studies report the results of these tests. The second dependent variable measures the speed of PT. This variable is taken from the workhorse model in bivariate price transmission analysis, which is the vector error correction model (VECM):

$$\Delta p_t = \alpha\beta p_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta p_{t-i} + \varepsilon \tag{2}$$

where $p_t$ is a bivariate vector of prices, $\beta' = (1,\ \beta_1)$ is the cointegrating vector such that $\beta' p_{t-1} = ect$ measures the equilibrium error or deviation from long-run equilibrium, $\alpha' = (\alpha_A,\ \alpha_B)$ is the vector of adjustment parameters that measure the speed with which such deviations are corrected, and $\Gamma_i$ describes autoregressive short-run dynamics. We define the second dependent variable as the aggregate speed of adjustment:

$$A_{AB} = -\alpha_A + \alpha_B \tag{3}$$

According to the standard specification of the VECM the $\alpha_A$ must be negative to correct deviations from the long-run equilibrium between the two prices. Therefore, subtracting it from $\alpha_B$, which is expected to be positive, yields the aggregate price adjustment in both markets, which is a measure of how much of any disequilibrium is corrected per period and, hence, the speed of PT.

A bit less straightforward is the construction of this variable when nonlinearities are introduced to the estimation. Usually this involves some sort of regime switch which allows the adjustment and / or other parameters to change over time. In exogenous break VECMs (Qiu, 2013; Thompson and Bohl, 1999) this switch is an exogenous time point or event, and in Markov switching VECMs (MSVECM) (Brümmer et al., 2009) this switch follows a stochastic process. In threshold VECMs (TVECM) (Goodwin and Piggott, 2001), the size of the deviation $\beta' p_{t-1} = ect$ triggers the regime switch. All of these models can be summarized as follows:

$$\Delta p_t = \alpha^{I_t} ect_{t-1} + \sum_{i=1}^{k-1} \Gamma_i^{(I_t)} \Delta p_{t-i} + \varepsilon_t \tag{4}$$

where $I_t \in \{1, \dots, M\}$ is an indicator variable for regimes 1 to $M$ at time $t$, $M \geq 2$, and $\alpha^{(I_t)}$ are the regime-dependent adjustment speeds in the individual regimes.[3] To measure the aggregate speed of

[3] The TVECM is a special case of regime-switching ECMs and estimated as:

$$\Delta p_t = \begin{cases} \alpha^{(u)} ect_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta p_{t-i} + \varepsilon_t \text{ if } ect_{t-1} < \varphi^u \\ \alpha^{(m)} ect_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta p_{t-i} + \varepsilon_t \text{ if } \varphi^u \leq ect_{t-1} \leq \varphi^l \\ \alpha^{(l)} ect_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta p_{t-i} + \varepsilon_t \text{ if } \varphi^u < ect_{t-1} \end{cases}$$

with upper and lower thresholds $\varphi^u$, $\varphi^l$ and the adjustment parameters in the upper, middle and lower regimes: $\alpha^u, \alpha^m, \alpha^l$. Each adjustment parameter is weighed by the relative size of its regime. For the three regime case this is the following equation:

$$A_{AB} = \frac{(\alpha_B^u - \alpha_A^u) * obs^u + (\alpha_B^m - \alpha_A^m) * obs^m + (\alpha_B^l - \alpha_A^l) * obs^l}{obs}$$

In a two regime TVECM, the two thresholds can also be symmetric with $\varphi^u = \varphi^l$, yielding one inner and one price adjustment in a model with multiple regimes, we weight the adjustment parameters from each regime by the number of observations in that regime:

$$A_{AB} = \frac{\sum_{i=1}^{M}(\alpha_B^i - \alpha_A^i) * obs^i}{obs} \tag{5}$$

In a number of studies, only one equation of the VECM is reported (e.g. Conforti, 2004b; Minot, 2011).[4] This usually because the authors assume that one of the two prices is exogenous (e.g. in the case of PT between a large exporter and a small importer) and thus only estimate the adjustment parameter that corresponds to the other price. for such studies, we follow the assumption that this market does practically all the adjustment and use only this first component ($-\alpha_A$) for the estimation while $\alpha_B$ is set zero.

The two dependent variables are regressed on a number of covariates including the variables of interest, distance and border (see Table 1). Distance is measured as the straight airline distance, and the border variable is a dummy that takes one if the two markets A and B are separated by an international border. We include a crop dummy in all estimated equations to capture any systematic differences between cereals (Table 1). A dummy variable indicates estimates from peer reviewed studies in order to explore a publication bias effect. Altogether, the basic meta-regression that we estimate is:

$$Y = \beta_0 + \beta_1 dist + \beta_2 border + \beta_3 crop + \beta_4 peer.reviewed + \beta_5 levels + \varepsilon \tag{6}$$

where Y is either a dummy which measures whether the two prices are cointegrated (LOGIT), or the aggregate speed of PT defined above (OLS). An interaction term for the distance and border variables is included in order to differentiate the effect of inland distance from that of international distance. This accounts for other dominating modes of transport in international trade (ship, airplane) than in domestic trade (cars, railroad).

When Y is the aggregate speed of PT, we test the influence of several additional explanatory dummy variables. First, it has been demonstrated that the frequency of the underlying data affects the size of the adjustment parameters in a VECM (Amikuzuno, 2010; von Cramon-Taubadel et al., 2003). We include dummy variables for different data frequencies to account for this factor. Second, a dummy captures the effect of the different statistical properties of the underlying price series in levels or in logarithmic form (*levels*). Third, a variable takes into account whether the cointegrating vector is restricted to [1, -1] or estimated (*beta.restrict*). Fourth, if in the estimation one of the two markets is assumed exogenous the variable *B.exogenous* equals one and zero otherwise. Fifth, the authors can include the possibility of price transmission changing over time into the modelling process in different ways. We expect the respective specifications to affect the results. One is a structural break in the long-run equilibrium (*beta.break*). Other specifications allow for a change in the speed of price adjustment in ECMs (*break.ecm*) or SEECMs (*break.seecm*). Breaks may also occur following markov-switching regime changes (*msvecm*). Threshold models with two or three regimes let the speed of adjustment speed vary depending on the size of the price change relative to estimated threshold values. We employ two dummies, *tvecm.2* and *tvecm.3*, to account for the expected heterogeneity in the estimates. The addition of these explanatory variables results in the following model:

$$\begin{aligned} A_{AB} = \ & \beta_0 + \beta_1 dist + \beta_2 border + \beta_3 dist * border + \beta_4 beta.restrict + \beta_5 crop \\ & + \beta_6 frequency + \beta_7 levels + \beta_8 msvecm + \beta_9 tvecm.2 + \beta_{10} tvecm.3 \\ & + \beta_{11} beta.break + \beta_{12} break.ecm + \beta_{13} break.seecm + \beta_{14}\varepsilon \end{aligned} \tag{7}$$

outer regime. The weighing is adapted accordingly.

[4] Phillips and Loretan (1991) introduced the term single equation error correction model (SEECM) for this model type.

These models are estimated in un-weighted form, and using a weighing scheme, in which each observation is weighted by the inverse number of observations taken from the same study. This gives each study in our meta-sample the same weight (WLS). This is potentially important because in particular two studies in our meta-sample (Hernandez-Villafuerte, 2011 with 273 observations, and Greb et al. 2012 with 497 observations – see Annex Table 2) account for a disproportionate share of the total of 1189 observations at our disposal.

# 4. Data and estimation

Price transmission literature is a wide field and it is a demanding task to extract those studies that are suitable for the intended meta regression. A first, explorative search was conducted on AgEcon-Search[5], where the keywords cointegration and price transmission produced 689 hits. Based on a systematic screening of all titles and abstracts we developed a search algorithm to find those studies that report PT estimates for spatially separated markets (see Annex). This was used for an exhaustive search in March 2013 on ISI Web of Knowledge which resulted in 962 studies. A weekly search alert added 38 studies to the sample. In addition, we used online searches of 62 journals of Agricultural Economics with the keywords price transmission, cointegration, ECM and error correction (see Annex). 11604 titles were screened and 163 studies were added to the sample. A similar search of the first 1000 hits on Google scholar (scholar.google.de) added another 27 studies. A list of references and several other sources added a further 49 studies. To assess the consistency of the study selection by the reviewers we performed Cohen's Kappa test (Cohen, 1960). Two independent reviewers screened 655 study titles retrieved on ISI Web of Science and the overlap of their study selection was compared with the test. The result indicates moderate agreement (0.570) which is sufficient for a systematic review.

After checking for duplicates and updated versions, we were left with 1648 non-duplicate studies. We screened the information in abstracts and retrieved from this sample all English-language studies of spatial grain price transmission. We further limited the sample to include only those that employ some sort of ECM, data from after 1980, and data of at least quarterly frequency. Conference posters were excluded in general. This selection procedure left us with 189 studies that were subjected to detailed screening. Of these, 57 studies proved suitable for coding with all required parameters reported.

[5] http://ageconsearch.umn.edu

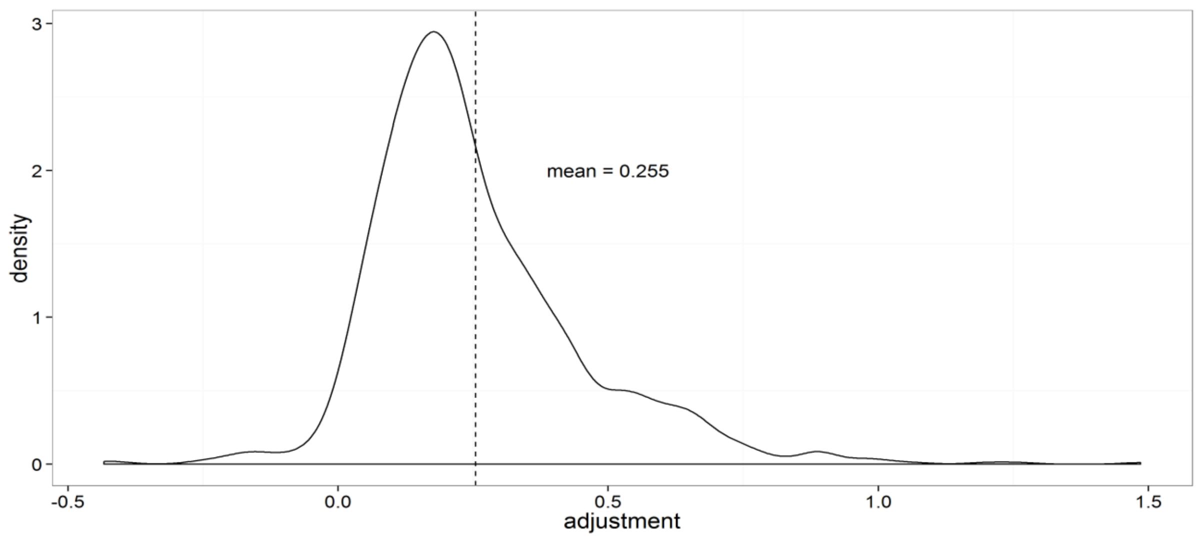
**Table 1: Variables and description with descriptive statistics**

| | Variable | Description | sum | mean | min | max |
|---|---|---|---|---|---|---|
| Y | cointegration | tested at 5% level of significance | 708 | 0.595 | 0 | 1 |
| | $A_{AB}$ | $\alpha_B$ - $\alpha_A$ | - | 0.255 | -0.434 | 1.486 |
| X | distance | airline distance in 1000 km | - | 5.785 | 0.019 | 19.713 |
| | border | = 1 if border separates price pairs | 761 | 0.640 | 0 | 1 |
| crop | rice | = 1 for rice (0 otherwise) | 641 | 0.539 | 0 | 1 |
| | maize | = 1 for maize | 385 | 0.324 | 0 | 1 |
| | wheat | = 1 for wheat | 133 | 0.112 | 0 | 1 |
| | soybeans | = 1 for soybeans | 20 | 0.017 | 0 | 1 |
| | sorghum | = 1 for sorghum | 6 | 0.005 | 0 | 1 |
| | teff | = 1 for teff | 1 | 0.001 | 0 | 1 |
| | barley | = 1 for barley | 3 | 0.002 | 0 | 1 |
| study | peer.reviewed | = 1 if article underwent peer review before published | 111 | 0.093 | 0 | 1 |
| data | levels | = 1 if series are in levels rather than logarithms | 37 | 0.031 | 0 | 1 |
| model | ecm | = 1 if error correction model | 833 | 0.701 | 0 | 1 |
| | B.exogenous | = 1 if one market is exogenous | 210 | 0.176 | 0 | 1 |
| | beta.break | = 1 for VECM with break in beta | 121 | 0.102 | 0 | 1 |
| | break.ecm | = 1 for VECM with break | 5 | 0.004 | 0 | 1 |
| | break.seecm | = 1 for break in model with one exogenous market | 6 | 0.005 | 0 | 1 |
| | markov | = 1 for MSVECM | 9 | 0.008 | 0 | 1 |
| | tvecm.2 | = 1 for 2 regime TVECM | 14 | 0.012 | 0 | 1 |
| | tvecm.3 | = 1 for 3 regime TVECM | 18 | 0.015 | 0 | 1 |
| | beta.restrict | = 1 if cointegrating vector is set unity | 35 | 0.029 | 0 | 1 |
| data | monthly | = 1 for monthly data | 1096 | 0.922 | 0 | 1 |
| | daily | = 1 for daily data | 28 | 0.024 | 0 | 1 |
| | every4days | = 1 for 4-daily data | 4 | 0.003 | 0 | 1 |
| | weekly | = 1 for weekly data | 58 | 0.049 | 0 | 1 |
| | quarterly | = 1 for quarterly data | 3 | 0.003 | 0 | 1 |
| study | study size (weights) | *av. number of observations per study | - | *20.86 | 1 | 497 |

*Note: N = 1189, each observation is one extracted price pair from a sample of 57 studies.*

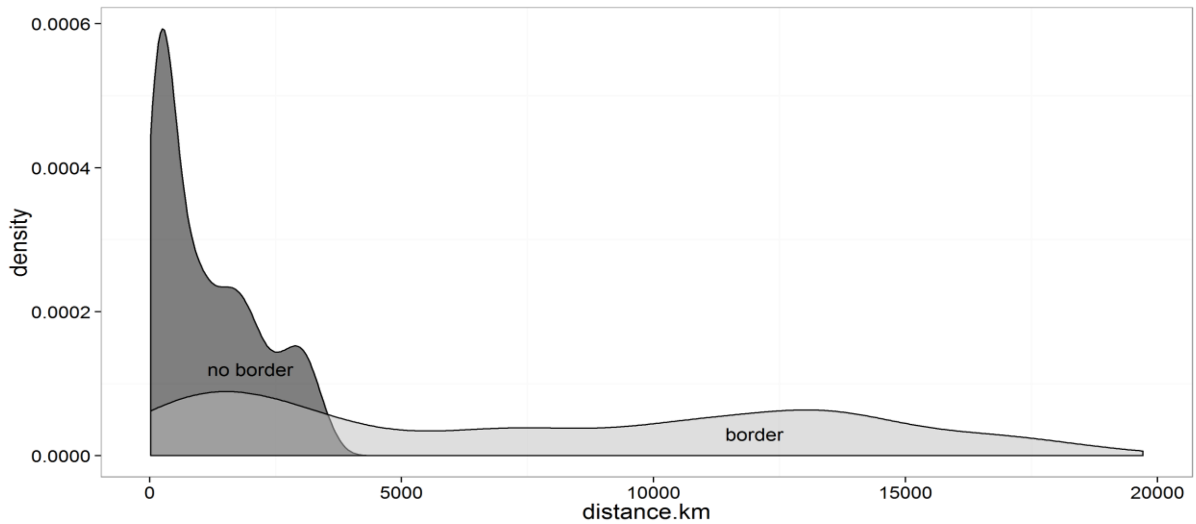
The extracted market pairs sum up to 1189 observations of which 167 come from nonlinear model specifications. Overall, 59.5% of all market pairs are cointegrated. If a price shock occurs, prices in both markets correct this deviation on average by 26% in each following period. The median is 0.213 with a right-skewed distribution (Figure 1).

**Figure 1: Density plot of speed of adjustment**




The mean of geographical distance is 5785 kilometers and 64% of the market pair are divided by a border. These two main explanatory variables are correlated, as illustrated in the density plot in Figure 2. If two markets are located in the same country, the distance between them never exceeds 5000 kilometers. Markets divided by a border are separated by as much as 19713 kilometers (Lima-Bangkok)

**Figure 2: Density plot of distance and borders**




More than half of the observations are taken from the studies of rice PT (53.9%), followed by maize with 32.4% and wheat with 11.2%. Soybeans, sorghum, teff and barley account for the remaining 2.5% of the sample. 70.1% of the 1189 PT estimates in the meta-dataset are produced using linear VECMs. 10.2% of the estimates are modelled with a structural break in the long run equilibrium and 0.9% with a structural break in the price adjustment coefficients. In 17.6% of the cases the authors assume that the second price exogenous and only estimate one equation. Only 3.5% of the observations are estimated with more complex nonlinear models such as the MSVECM or the TVECM. Most price series are of monthly frequency (92.2%), and almost studies employ prices in logarithms (96.9%). While 24 of the 57 studies are published in peer reviewed journals, these contribute only 9.3% of the observations in the sample. On average, each study provides 21 observations.

# 5. Results and discussion

Meta-regression results in Table 2 indicate that distance and borders have statistically significant and negative effects on both the likelihood of cointegration and the speed of price adjustment. The size of these effects differs according to model specification and weighting scheme.

**Table 2: Results from meta-regressions**

| Dependent variable → | | cointegration | | speed of adjustment | | | |
|---|---|---|---|---|---|---|---|
| **Type↓** | **Covariate↓** | **LOGIT** | ***(se)*** | **OLS** | ***(se)*** | **WLS** | ***(se)*** |
| | (Intercept) | 0.346*** | *(0.096)* | 0.368*** | *(0.023)* | 0.468*** | *(0.036)* |
| X | distance | -0.069* | *(0.040)* | -0.057*** | *(0.012)* | -0.199*** | *(0.033)* |
| | border | -0.225** | *(0.104)* | -0.020 | *(0.028)* | -0.127*** | *(0.029)* |
| | distance *border | 0.059 | *(0.040)* | 0.056*** | *(0.012)* | 0.186*** | *(0.033)* |
| crop | maize | -0.154*** | *(0.049)* | -0.012 | *(0.021)* | -0.001 | *(0.025)* |
| | wheat | -0.109* | *(0.056)* | -0.102*** | *(0.030)* | -0.035* | *(0.020)* |
| | soybeans | 0.289 | *(0.250)* | -0.086 | *(0.055)* | -0.057 | *(0.040)* |
| | sorghum | -0.170 | *(0.207)* | 0.052 | *(0.135)* | 0.178 | *(0.126)* |
| | teff | 0.453 | *(0.382)* | 0.030 | *(0.192)* | -0.049 | *(0.062)* |
| | barley | 0.461** | *(0.222)* | -0.002 | *(0.115)* | -0.001 | *(0.060)* |
| study | peer.reviewed | -0.117 | *(0.091)* | 0.078** | *(0.038)* | 0.050** | *(0.020)* |
| data | levels | 0.292 | *(0.181)* | 0.102** | *(0.045)* | -0.040 | *(0.026)* |
| frequency | daily | | | -0.229*** | *(0.060)* | -0.251*** | *(0.045)* |
| | every 4 days | | | -0.295** | *(0.130)* | -0.355*** | *(0.073)* |
| | weekly | | | 0.001 | *(0.037)* | -0.092*** | *(0.024)* |
| | quarterly | | | -0.133 | *(0.156)* | -0.266*** | *(0.066)* |
| model | msvecm | | | 0.092 | *(0.083)* | 0.200*** | *(0.056)* |
| | tvecm.2 | | | -0.040 | *(0.064)* | -0.073* | *(0.039)* |
| | tvecm.3 | | | -0.056 | *(0.081)* | -0.185*** | *(0.043)* |
| | beta.break | | | 0.050** | *(0.024)* | 0.211** | *(0.090)* |
| | break.ecm | | | -0.124 | *(0.099)* | -0.177*** | *(0.050)* |
| | break.seecm | | | -0.171 | *(0.114)* | -0.168*** | *(0.056)* |
| | beta.restrict | | | 0.012 | *(0.067)* | 0.117*** | *(0.030)* |
| | B.exogenous | | | -0.130*** | *(0.023)* | -0.110*** | *(0.018)* |
| | AIC | 1301.977 | | | | | |
| | BIC | 1360.774 | | | | | |
| | Log Likelihood | -638.989 | | | | | |
| | Deviance | 1277.977 | | | | | |
| | Num. obs. (df) | 992 (df=980) | | 705 (df=681) | | 705 (df=681) | |
| | (Pseudo-)$R^2$ | 0.120 | | 0.152 | | 0.275 | |
| | Adj. $R^2$ | | | 0.123 | | 0.251 | |

*Note: In the logistic model specification, average marginal effects rather than coefficient estimates and a Pseudo-R2 are reported. WLS model observations are weighted by inverse study size. Reported are standard errors in brackets and significance levels: ***p < 0.01, **p < 0.05, *p < 0.1.*

According to the LOGIT specification, the probability of cointegration between two prices falls by 22.5% if the markets in question are separated by an international border. The probability of cointegration between two prices within a country falls by 6.9% with each additional 1000 kilometer airline distance. The interaction term is not statistically significant at the 10% level, indicating that the distance effect is the same for two markets separated by an international border. Compared to the base category rice, maize and wheat prices are on average 15.4% and 10.9% less often cointegrated, while barley price pairs are 46.1% more likely cointegrated. Cointegration test results in peer-reviewed articles do not differ statistically significant from grey literature studies. Compared with rice, wheat market pairs are 10.2% less often cointegrated. Moreover, whether the price series were log-transformed or estimated in levels does not affect the cointegration test results.

A negative and significant effect of distance on the speed of price transmission is common to all linear model specifications. Each 1000 kilometers of inland distance reduce the speed of adjustment by 5.7 percent, international distance however only by 0.01%. A border reduces the speed on average by 2%, however the effect is not statistically significant. In peer-reviewed publications the speed of adjustment is 7.8% faster than in grey literature studies, pointing slightly to publication bias towards larger coefficients. Price series in levels produce 10.2% faster price transmission results than log-transformed series. This effect originates from the different statistical properties of series in levels and logs, e.g. with the logarithmic form exhibiting less variation. Using higher daily or 4-daily price data for the estimation, price adjustment is 23% and 30% faster respectively compared to monthly data frequency. This confirms findings of the impact of data aggregation on price transmission parameters (von Cramon-Taubadel et al., 2003). Only two model-type variables affect the results, namely a break in the long-run equilibrium increases the adjustment speed by 5% and in estimations with one exogenous market, the price adjustment is 13% slower.

In the WLS estimation the results indicate a much higher distance effect with 19.9% slower price adjustment per 1000 kilometers and 12.7% slower adjustment in presence of a border. The interaction term also decreases an international distance effect to only 1.3%. Transmission of wheat prices is on average 3.5% slower than transmission of rice prices, while adjustment speeds of other crops do not differ significantly from rice. Estimated speeds of adjustment in peer reviewed journals are 5% higher than in grey literature studies. This may be evidence of publication bias in journals in favor of studies with stronger results. Another possible explanation is that studies undergo a peer review that sorts out lower quality results and results in more accurate estimates of price transmission parameters. Model type variables explain the variance in results far better in the WLS than the unweighted specification. Adjustment speeds in MSVECMs are 20% higher and 7.3% and 18.5% slower for TVECMs with 2 and 3 regimes, respectively. A break that allows for different long run equilibria increases price transmission by 21.1%. Models with structural breaks in the adjustment process decrease the speed of adjustment by 17.7% for ECMs and 16.8% for ECMs with one endogenous market. Price transmission is 11.7% faster when the price transmission coefficient is restricted to 1 in the price transmission equation. If one market is exogenous, the adjustment speed is 11% slower. The coefficient of determination of the meta-regressions indicates that the covariates explain between 12% and 28% of the variation in the results.

# 6. Conclusion

We test whether the distance between two markets and whether they are separated by an international borders affects the strength and speed of price transmission between them. To do so we extract measures of the strength and speed of spatial cereal price transmission from the empirical PT literature, and regress these on the distance between markets and whether there is a border between them. The results of a number of meta-regression specifications confirm that distance and borders have significant effects on cointegration and the speed of price transmission. We find that distance effects are only economically significant for the speed of price transmission within a country. If two markets are separated by an international border, the distance effect is negligible. This confirms findings on international trade that transportation costs over longer distances are disproportionally low compared to those over shorter distances. Freight rates may thus not inhibit market integration over longer distances so that price transmission is almost unaffected by international distance. Furthermore, the properties of the data and the model choice also influence price transmission findings. When interpreting estimates from price transmission studies, these effects have to be taken into account. The policy implications of this study point at the relevance of borders and distance when focusing on market integration. Policy measures targeted at specific markets may not affect distant or foreign partner markets.

We have tested whether linear distance has an effect on the strength and speed of price transmission. The effect of distance might be nonlinear however, as different modes of transportation are used to cover different distances in cereal trade. Future research could look for evidence of such nonlinear effects, for example using semi-parametric techniques.

## 7. References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Greb F, Jamora N, Mengel C, et al. (2012) Price transmission from international to domestic markets. *GlobalFood Discussion Papers*, GlobalFood Discussion Papers, Göttingen, Germany, 15, 1–48.

Hernandez-Villafuerte K (2011) The relationship between spatial integration and geographical distance in Brazil. *Annual Meeting of the European Association of Agricultural Economics (EAAE), Zürich, Switzerland*, Zürich, 1–37.

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Phillips P and Loretan M (1991) Estimating long-run economic equilibria. *The Review of Economic Studies*, (58), 407–436.

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## 8. Appendix

**Table 1: List of peer-reviewed journals with number of keyword hits**

| Study no. | Name of journal (J) | access via | number of hits | | | |
|---|---|---|---|---|---|---|
| | | | price transmission | cointegration | ecm | error correction |
| 1 | Acta Oeconomica et Informatica | fem.uniag.sk | 63 | 6 | 29 | 47 |
| 2 | African J of Agricultural Research | academicjournal.org | 606 | 269 | 309 | 898 |
| 3 | Agribusiness: An International J | Wiley | 112 | 49 | 9 | 89 |
| 4 | Agricultural and Food Science | ojs.tsv.fi | 2 | 1 | 0 | 2 |
| 5 | Agricultural Economics | Science Direct | 62 | 33 | 14 | 88 |
| | | Wiley | 111 | 71 | 20 | 144 |
| 6 | Agricultural Economics Research Review | EBSCO | 2 | 2 | 0 | 0 |
| 7 | Agricultural Finance Review | Emerald | 0 | 1 | 0 | 1 |
| 8 | Agricultural Systems | Science Direct | 62 | 4 | 7 | 123 |
| 9 | Agriculture and Human Values | Springer | 37 | 0 | 1 | 15 |
| 10 | Agriculture, Ecosystems and Environment | Science Direct | 35 | 0 | 12 | 267 |
| 11 | Agroforestry Systems | Springer | 24 | 0 | 84 | 216 |
| 12 | American J of Agricultural Economics | EBSCO | 19 | 22 | 1 | 14 |
| 13 | Applied Economics Letters | EBSCO | 7 | 286 | 16 | 114 |
| | | Taylor & Francis | 118 | 700 | 167 | 549 |
| 14 | Asian Agricultural Research | scialert.net | 0 | 0 | 0 | 0 |
| 15 | Australian J of Agricultural and Resource Economics | Wiley | 52 | 22 | 4 | 69 |
| | | EBSCO | 2 | 6 | 0 | 2 |
| 16 | Bio-based and Applied Economics | fupress.net | 5 | 0 | 0 | 2 |

| Study no. | Name of journal (J) | access via | number of hits | | | |
|---|---|---|---|---|---|---|
| | | | price transmission | cointegration | ecm | error correction |
| 17 | British Food J | Emerald | 31 | 2 | 0 | 9 |
| 18 | Canadian J of Agricultural Economics | Wiley | 98 | 59 | 15 | 106 |
| 19 | Computers and Electronics in Agriculture | Science Direct | 65 | 0 | 6 | 329 |
| 20 | Economia | anpec.org.br/revista | 110 | 165 | 19 | 145 |
| 21 | European Review of Agricultural Economics | Oxford Journals | 1483 | 39 | 10 | 588 |
| 22 | Food Policy | Science Direct | 101 | 17 | 5 | 60 |
| 23 | Food Quality and Preference | Science Direct | 13 | 0 | 0 | 96 |
| 24 | J of Agricultural and Food Industrial Organization | De Gruyter | 16 | 7 | 2 | 11 |
| 25 | J of Agricultural Economics | Wiley | 109 | 51 | 16 | 141 |
| 26 | J of Agriculture and Rural Development in the Tropics and Subtropics | jarts.info | 1 | 1 | 0 | 0 |
| 27 | J of Applied Economics | EBSCO | 0 | 12 | 3 | 7 |
| | | IDEAS | 0 | 26 | 3 | 9 |
| 28 | J of Development Studies | EBSCO | 1 | 0 | 2 | 5 |
| | | Taylor & Francis | 107 | 42 | 29 | 152 |
| 29 | J of International Agricultural Trade & Development | EBSCO | 1 | 16 | 1 | 6 |
| | | Taylor & Francis | 44 | 57 | 15 | 46 |
| 30 | J of Policy Reform | EBSCO | 1 | 0 | 0 | 0 |
| 31 | J of Regional Analysis and Policy | jrap-J.org | 6 | 8 | 2 | 33 |
| 32 | J of Rural Development | krei.re.kr | 10 | 10 | 10 | 10 |
| 33 | J of Sustainable Agriculture / Agroecology and Sustainable Food Systems | Taylor & Francis | 11 | 0 | 1 | 1 |
| 34 | Marine Resource Economics | BioOne | 3 | 2 | 0 | 13 |

| Study no. | Name of journal (J) | access via | number of hits | | | |
|---|---|---|---|---|---|---|
| | | | price transmission | cointegration | ecm | error correction |
| 35 | Renewable Agriculture and Food Systems | Cambridge Js | 8 | 0 | 3 | 24 |
| 36 | Review of Agricultural Economics | EBSCO | 0 | 2 | 1 | 2 |
| 37 | Stata J | stata-J.com | 0 | 3 | 0 | 1 |
| 38 | Studies in Agricultural Economics | aki.gov.hu | 0 | 0 | 0 | 0 |
| 39 | Western J of Agricultural Economics | jstor.org | 0 | 0 | 0 | 0 |
| 40 | World Development | Science Direct | 368 | 82 | 19 | 356 |
| **Sum of hits** | | | **3906** | **2073** | **835** | **4790** |
| | **Journals included in AgEcon Search:** | | | | | |
| 41 | J of Agricultural and Applied Economics | | | | | |
| 42 | Agrekon | | | | | |
| 43 | Agricultural and Resource Economics Review | | | | | |
| 44 | Australian J of Agricultural Economics | | | | | |
| 45 | Brazilian J of Rural Economy and Sociology | | | | | |
| 46 | Current Agriculture, Food and Resource Issues | | | | | |
| 47 | International Food and Agribusiness Management Review | | | | | |
| 48 | J of Agribusiness | | | | | |
| 49 | J of Agricultural and Resource Economics / WJAE / WEF | | | | | |
| 50 | Scientific J on Agricultural Economics | | | | | |

**Table 2: Studies and number of observations in the meta-sample**

| Study | obs. |
|---|---|
| Acosta, A., 2012. Measuring spatial transmission of white maize prices between South Africa and Mozambique: An asymmetric error correction model approach. African Journal of Agricultural and Resource Economics, 7(1), pp.1–13. | 1 |
| Alam, M. & Begum, I., 2012. World and Bangladesh rice market integration: An application of threshold cointegration and threshold vector error correction model (TVECM). In Annual Conference of the Agricultural Economics Society (AES), Warwick, United Kingdom. University of Warwick, UK, pp. 1–19. | 1 |
| Alam, M. et al., 2012. Measuring Market Integration in the Presence of Threshold Effect: The Case of Bangladesh Rice Markets. In Agricultural & Applied Economics Association`s (AAEA) Conference, Seattle, Washington, USA. Washington, USA, pp. 1–27. | 10 |
| Alam, M.J. et al., 2012. The dynamic relationships between world and domestic prices of rice under the regime of agricultural trade liberalization in Bangladesh. Journal of the Asia Pacific Economy, 17(1), pp.113–126. | 1 |
| Araujo-Enciso, S., 2009. Evidence of non-linear price transmission between maize markets in Mexico and the US. In International Association of Agricultural Economists Conference, Beijing, China. Beijing, China, pp. 1–23. | 5 |
| Baek, J. & Koo, W.W., 2006. Price Dynamics in the North American Wheat Market. Agricultural and Resource Economics Review, 2(October 2003), pp.265–275. | 2 |
| Balcombe, K., Bailey, A. & Brooks, J., 2007. Threshold Effects in Price Transmission: The Case of Brazilian Wheat, Maize, and Soya Prices. American Journal of Agricultural Economics, 89(2), pp.308–323. | 5 |
| Baquedano, F.G., Liefert, W. & Shapouri, S., 2011. World market integration for export and food crops in developing countries: a case study for Mali and Nicaragua. Agricultural Economics, 42(5), pp.619–630. | 6 |
| Baulch, B. & Hansen, H., 2008. The spatial integration of paddy markets in Vietnam. Journal of Agricultural Economics, 59(2), pp.271–295. | 7 |
| Brosig, S. & Yahshilikov, Y., 2005. Interregional Integration of Wheat Markets in Kazakhstan. IAMO Discussion Paper Series, (88), pp.1–35. | 4 |
| Chapoto, A., 2012. The Political Economy of Food Price Policy: The Case of Zambia. UNU WIDER Working Paper, (100), pp.1–27. | 9 |
| Chirwa, E., 2001. Food pricing reforms and price transmission in Malawi: Implications for food policy and food security. University of Malawi Working Paper Series, (4), pp.1–34. | 14 |
| Conforti, P., 2004. Price transmission in selected agricultural markets. FAO Commodity and Trade Policy Research Working Paper Series, (7), pp.1–91. | 15 |
| Coxhead, I., Linh, V. & Tam, L., 2012. Global market shocks and poverty in Vietnam: the case of rice. Agricultural Economics, 43(5), pp.575–592. | 3 |
| Dawson, P. & Sanjua, A., 2006. Structural Breaks , the Export Enhancement Program and the Relationship between Canadian and US Hard Wheat Prices. Journal of Agricultural Economics, 57(1), pp.101–116. | 1 |
| Djuric, I., Götz, L. & Glauben, T., 2011. Effects of the governmental market interventions on the wheat market in Serbia during the food crisis 2007 / 2008. In Annual Meeting of the German Society of Economic and Social Sciences in Agriculture (GEWISOLA), Halle, Germany. Halle, Germany, pp. 1–14. | 1 |
| Dutoit, L., Hernandez-Villafuerte, K. & Urrutia, C., 2009. Price transmission in Latin American maize and rice markets. ECLAC Working Paper, pp.1–47. | 26 |

| Study | obs. |
|---|---|
| Fabiosa, J.F., 2000. Impact of GATT in the Functioning of Agricultural Markets: An Examination of Market Integration and Efficiency in the World Beef and Wheat Market under the pre-GATT and post-GATT Regimes. In Annual Meeting of the American Agricultural Economics Association (AAEA), Tampa, United States. Tampa, Florida, USA, pp. 1–17. | 1 |
| Fiamohe, R. et al., 2013. Price transmission analysis using threshold models: an application to local rice markets in Benin and Mali. Food Security, 5(3), pp.427–438. | 4 |
| Franken, J. et al., 2005. Market Integration: Case Studies of Structural Change. Agricultural and Resource Economics Review, 32(2), pp.163–172. | 6 |
| Getnet, K., 2007. Spatial Equilibrium of Wheat Markets in Ethiopia. African Development Review, 19(2), pp.281–303. | 1 |
| Getnet, K., Verbeke, W. & Viaene, J., 2005. Modeling spatial price transmission in the grain markets of Ethiopia with an application of ARDL approach to white teff. Agricultural Economics, 33(3, S), pp.491–502. | 1 |
| Ghoshray, A. & Ghosh, M., 2011. How Integrated is the Indian Wheat Market? Journal of Development Studies, 47(10), pp.1574–1594. | 8 |
| Ghoshray, A., 2008. Asymmetric Adjustment of Rice Export Prices : The Case of Thailand and Vietnam. International Journal of Applied Economics, 5(September), pp.80–91. | 1 |
| Ghoshray, A., 2007. An examination of the relationship between US and Canadian durum wheat prices. Canadian Journal of Agricultural Economics, 55, pp.49–62. | 1 |
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| Goodwin, B.K. & Piggott, N.E., 2001. Spatial Market Integration in the Presence of Threshold Effects. American Journal of Agricultural Economics, 83(2), pp.302–317. | 12 |
| Götz, L., Glauben, T. & Brümmer, B., 2013. Wheat export restrictions and domestic market effects in Russia and Ukraine during the food crisis. Food Policy, 38, pp.214–226. | 4 |
| Götz, L. et al., 2012. The Law of One Price under State-Dependent Policy Intervention: An Application to the Ukrainian Wheat Market. In Annual Meeting of the Agricultural & Applied Economics Association (AAEA), Seattle, United States. Seattle, Washington, USA, pp. 1–36. | 3 |
| Götz, L., Glauben, T. & Brümmer, B., 2010. Impacts of Export Controls on Wheat Markets During the Food Crisis 2007/2008 in Russia and Ukraine. In Annual Meeting of the Agricultural & Applied Economics Association (AAEA), Denver, United States. Denver, Colorado, USA, pp. 1–22. | 4 |
| Goychuk, K. & Meyers, W., 2011. Black Sea Wheat Market Integration with the International Wheat Markets: Some Evidence from Co-integration Analysis. In Annual Meeting of the Agricultural & Applied Economics Association (AAEA), Pittsburgh, United States. Pittsburgh, Pennsylvania, USA, pp. 1–16. | 10 |
| Greb, F. et al., 2012. Price transmission from international to domestic markets. GlobalFood Discussion Papers, 15, pp.1–48. | 497 |
| Greb, F. et al., 2013. The estimation of threshold models in price transmission analysis. American Journal of Agricultural Economics, 95(4), pp.900–916. | 12 |
| Grethe, H. et al., 2012. How do World Agricultural Commodity Price Spikes Affect the Income Distribution in Israel? Annual Meeting of the German Society of Economic and Social Sciences in Agriculture (GEWISOLA), Hohenheim, Germany, pp.1–13. | 2 |

| Study | obs. |
|---|---|
| Hai, T., 2003. Rice markets in the Mekong River Delta, Vietnam : A market integration analysis. CAS Discussion Paper Series, (40), pp.1–19. | 1 |
| Hernandez-Villafuerte, K., 2011. The relationship between spatial integration and geographical distance in Brazil. Annual Meeting of the European Association of Agricultural Economics (EAAE), Zürich, Switzerland, pp.1–37. | 273 |
| Huang, J., Yang, J. & Rozelle, S., 2013. The political economy of food pricing policy in China. UNU-WIDER Working Paper Series, (38), pp.1–29. | 4 |
| Ihle, R., von Cramon-Taubadel, S. & Zorya, S., 2011. Measuring the integration of staple food markets in Sub-Saharan Africa: Heterogeneous infrastructure and cross border trade in the East African community. CESifo Working Paper, 3413, 1–33. | 86 |
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| Minot, N., 2011. Transmission of World Food Price Changes to Markets in Sub-Saharan Africa. IFPRI Discussion Paper Series, 01059, pp.1–44. | 54 |
| Mohanty, S., Smith, D. & Peterson, W., 1996. Time Series Evidence of Relationships Between US and Canadian Wheat Prices. Iowa State University Working Paper Series, 154, pp.1–19. | 4 |
| Mohanty, S. & Langley, S., 2003. The Effects of Various Policy Regimes in the Integration of North American Grain Markets. Canadian Journal of Agricultural Economics, 51(1), pp.109–120. | 4 |
| Mutuc, M., Pan, S. & Hudson, D., 2011. Sino-US Price Transmission in Agricultural Commodities: How Important are Exchange Rate Movements? Annual Meeting of the Agricultural and Applied Economics Association (AAEA), Pittsburgh, United States, pp.1–26. | 3 |
| Nzuma, J., 2013. The political economy of food price policy: The case of Kenya. UNU-WIDER Working Paper Series, 26. | 5 |
| Pede, V. & McKenzie, A., 2008. Integration In Benin Maize Market: An Application Of Threshold Cointegration Analysis. Journal of International Agricultural Trade and Development, 5(1), pp.129–146. | 21 |
| Qiu, F., 2013. Dynamic impacts of export controls on price transmission. Annual Meeting of the Agricultural and Applied Economics Association (AAEA), Washington, United States, pp.1–13. | 3 |
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| Roche, M. & McQuinn, K., 2003. Grain price volatility in a small open economy. European Review of Agriculture Economics, 30(1), pp.77–98. | 2 |
| Sanogo, I. & Maliki Amadou, M., 2010. Rice market integration and food security in Nepal: The role of cross-border trade with India. Food Policy, 35(4), pp.312–322. | 1 |
| Sharma, R., 2003. The transmission of world price signals: the concept, issues and some evidence from Asian cereal markets. In In: Agricultural Trade and Poverty - Making Policy Analysis Count, OECD, Paris, France. pp. 141–160. | 24 |

| **Study** | **obs.** |
|---|---|
| Thompson, S., Sul, D. & Bohl, M., 2002. Spatial market efficiency and policy regime change: seemingly unrelated error correction model estimation. American Journal of Agricultural Economics, 84(4 (November)), pp.1042–1053. | 3 |
| Thompson, S. & Bohl, M., 1999. International wheat price transmission and CAP reform. Annual Meeting of the Agricultural & Applied Economics Association (AAEA), Nashville, United States, pp.1–20. | 4 |
| Vasciaveo, M., Rosa, F. & Weaver, R., 2013. Agricultural market integration : price transmission and policy intervention. Annual Conference of the Italian Association of Agricultural Economists (AIEAA), Parma, Italy, pp.1–27. | 3 |
| Yakhshilikov, Y. & Brosig, S., 2006. Spatial price transmission in Kazakh wheat markets. In Conference of the International Association of Agricultural Economists, Gold Coast, Australia. pp. 1–11. | 2 |
| **Sum of observations** | **1189** |