Proximity and price co-movement in West African rice markets

Carolin Kouyaté*
Department of Agricultural Economics and Rural Development, University of Göttingen, Germany. E-mail: c.kouyate@icloud.com

Stephan von Cramon-Taubadel
Department of Agricultural Economics and Rural Development, University of Göttingen, Germany. E-mail: scramon@gwdg.de

Ismael Fofana
International Food Policy Research Institute (IFPRI), West and Central Africa Office, Dakar, Senegal. E-mail: i.fofana@cgiar.org

* Corresponding author

Abstract

We have explored the link between proximity and price transmission in rice markets, where proximity is captured using variables for geographic, political and cultural distance. Linear and threshold cointegration was tested for 351 rice market pairs in six West African countries, with threshold specifications accounting for transaction costs. The influence of proximity on price transmission is estimated in a subsequent multinomial logistic regression. The results provide evidence of a robust and statistically significant link between the strength of price transmission and measures of proximity such as distance, international borders, contiguity and a common language. We conclude that proximity matters for market integration processes in West African rice markets.

Key words: West Africa; cointegration; rice; transaction costs; regional integration

1. Introduction

The relationship between market proximity and integration is widely acknowledged in the trade literature. The shorter the geographic, political and cultural distance between two locations, the more they trade and the more their prices adhere to the law of one price (LOP). However, the influence of proximity on price transmission has not yet been documented clearly in the literature. If proximity influences trade flows and price dispersion, then it presumably also influences price transmission. We therefore hypothesise that whether and how price signals are transmitted between markets depends on geographical, political and cultural proximity, and we propose an empirical test of this hypothesis based on an analysis of the determinants of rice market integration in West Africa.

West Africa is a food-deficit area and integration both within the region and with international markets can play an important role in cushioning shortages and food price shocks. The region is particularly vulnerable to international food price shocks due to its strong import dependence. During colonisation, France had fostered cheap rice imports from its former territory Indochina. Over time, West African consumers developed a preference for imported broken rice over local grains and tubers. Rice is the most important source of calories in the region.

Increasing local food supply is a central political goal in most West African countries, many of which have been severely affected by high food import prices in recent years. Rice is a key staple crop and plays an important role in national and regional agricultural development strategies. Regional rice consumption was estimated at 14 million tons of milled rice in 2012, of which 35 to 40% was imported. Almost all West African states have developed national rice development strategies (NRDS) alongside their sector-wide strategies. Furthermore, the Economic Commission of West African States (ECOWAS) actively supports NRDS under the regional programme for sustainable rice development in West Africa. All national and regional programmes aim to significantly increase the production of rice and achieve, in the medium term, rice self-sufficiency. Against this background, an improved understanding of price dynamics and market integration in the region can contribute to the formulation of welfare-enhancing policies.

The rest of this paper is structured as follows. In the following section we provide information on the West African rice market. In section 3 we review the literature and theory on the relationship between proximity, trade and price transmission. Sections 4 and 5 present our empirical approach and results respectively, and section 6 concludes.

2. The West African rice market

The West African rice market is segmented. Local and more nutritious varieties of rice are marketed alongside better packaged and standardised imported rice from international markets. Segmentation is also apparent in processing methods, with parboiled rice accounting for an important market share in some countries (e.g. Guinea, Mali and Nigeria), but not in all. West African rice markets are linked to international markets by trade. Three West African countries are among the top ten importers of rice in the world: Nigeria in 2nd place with 1.8 million tons; Côte d'Ivoire in 8th place with 1 million tons; and Senegal in 10th place with 0.8 million tons (ECOWAS *et al*. 2014). Increasing integration into international rice markets over the past decades has increased the region's exposure to global price shocks and growing agricultural price volatility. According to Fiamohe *et al*. (2015), transmission between international (Thai) and domestic rice prices in many West African markets is asymmetric, with increases in international prices being transmitted more rapidly than decreases. Senegal, where price transmission is symmetric, is an exception.

Intraregional trade flows, however, are limited (ECOWAS *et al*. 2014), as illustrated in Table 1. Indeed, imports from outside the region are about two hundred times higher than imports from other West African countries.

Table 1: Trade flow value of imported rice by importing country with trading partner located in or outside the West African region

Country	Regional trade: Imports from other West African countries (mill. USD)	Extra-regional trade: Imports from countries outside West Africa (mill. USD)	Share of regional trade in extra-regional trade (percent)
Benin	4.6	673.2	0.68
Burkina Faso	19.3	644.9	2.99
Côte d'Ivoire	0.2	4 021.4	0.00
Gambia, The	1.2	110.6	1.07
Ghana	15.0	2 299.1	0.65
Guinea	4.6	418.4	1.11
Mali	29.4	386.1	7.62
Mauritania	9.0	174.9	5.17
Niger	13.9	1 012.9	1.38
Nigeria	0.8	5 539.3	0.01
Senegal	0.2	3 490.4	0.01
Togo	2.1	104.8	1.97
Total	**100.3**	**18 875.9**	**0.53**

Note: The panel chart displays the reported trade value of imported rice in West African countries from 2002 to 2013 in US dollar (USD). The trade values are depicted separately by whether the trading partner is also located in or outside the West African region. The last column demonstrates the relative difference of intra-regional and extra-regional trade. We retrieved rice trade flow data on UN Comtrade (United Nations 2010), as classified under the Harmonized System Code 1006 (HS 2002).

Most local supply is consumed in the rural areas in which it was produced. Rural-urban trade flows are sometimes important in countries such as Guinea and Mali, where domestic production meets 80% of domestic needs. According to a recent study by CILSS, cross-border trade flows of paddy are observed from Benin to Nigeria; parboiled rice flows from Burkina Faso to Mali; and milled rice is traded among Mali, Guinea, Senegal and Mauritania (ECOWAS *et al*. 2014). Border policies for rice differ across countries in the region. Nigeria imposes customs duties of nearly 50%, while the eight members of the West African Economic and Monetary Union/Union Èconomique et Monétaire Ouest Africaine (UEMOA) impose 10%. These differences lead to trade diversions, with surplus imports in some countries being smuggled into neighbouring countries. An ECOWAS Common External Tariff is currently being negotiated (Diouf 2012) and is expected to facilitate the implementation of a more liberal rice trade policy in the region. However, there is still no agreement on a time schedule for implementing this tariff.

3. Theory and literature

Fackler and Goodwin (2001:978) define market integration as "the measure of the degree to which demand and supply shocks arising in one region are transmitted to another region". Price transmission analysis studies the dynamics of this transmission between markets. Price transmission can take place as a result of physical trade due to arbitrage, either directly or via third markets. It can also take place in the absence of physical trade as a result of information flows (Jensen 2007; Stephens *et al*. 2012). Over time, price transmission manifests itself as co-movement of prices in the respective markets. Since Ardeni (1989), this co-movement has overwhelmingly been modelled using cointegration techniques. In the case of price transmission between two geographically separated markets (spatial price transmission), so-called threshold vector error correction models (TVECMs) are commonly employed. The TVECM allows modellers to explicitly account for the costs of trade between two locations. In a TVECM, the transmission of supply and demand shocks between markets depends on the difference between the prices in these markets. If prices differ by more than the costs of trade between two markets, trade is triggered and shocks are transmitted; if prices differ by less than the costs of trade there is no incentive for trade and correspondingly no transmission of shocks. Hence, in a simple two-market setting, the TVECM combines a regime of price transmission, within which prices co-move as a result of trade flows, with a neutral regime within which prices move independently of one another (Greb *et al*. 2013). In multimarket settings, price transmission between any two markets can display more complex threshold structures. For example, even if the difference between two prices is not sufficient to trigger trade between the markets in question, these prices might nonetheless co-move as a result of trade or information flows with third markets. In such a setting, thresholds therefore will not necessarily separate regimes with and without price transmission, but rather regimes with different degrees and speeds of price transmission.

Whether and how prices in spatially separate markets co-move is thus closely related to the costs of trade and communication between these markets. Definitions of these costs and their effects on trade are discussed in the trade literature. Anderson and Van Wincoop (2004) classify trade costs as (i) transportation costs such as gas and tolls, (ii) trade barriers such as custom procedures and tariffs and (iii) transaction costs such as long-distance phone calls and translation. Each of these trade cost components can vary with proximity. Gravity models link trade volumes to market size and market proximity. Proximity has a geographic dimension, but cultural similarities such as a shared language can also imply proximity, as can political-economic factors such as joint membership in a regional trading agreement. Zannou (2010) finds that the volume of commodity trade between markets in West Africa falls with increasing distance between them and if they are separated by an international border. He also finds that a common official language and contiguity are positively correlated with trade volumes. Other findings suggest that price dynamics in East African maize markets also depend on borders and marketing costs, including for transport (World Bank 2009). In this strand of the literature, the explanation given is that trade costs increase with distance, thus reducing commercial activity and communication. This link between distance/borders, trade costs and trade volumes has been confirmed in many studies and settings.

The literature on price disparities and the LOP provides a related perspective on the link between prices and market proximity. A number of studies analyse the effects of distance and borders on deviations from the LOP. The seminal paper by Engel and Rogers (1995) finds that the border between Canada and the United States has the same effect on price disparities as 2 500 kilometres of distance. Similarly, Aker *et al*. (2013) find a statistical link between borders and price disparities between markets in Niger and Nigeria.

If proximity and borders affect trade flows and price differentials between markets, they will likely also affect the process of price transmission between these markets. However, only a few studies have explored this link to date. Hernandez-Villafuerte (2011) finds a significant negative effect of road distance on the long-run elasticity of price transmission between Brazilian rice markets. Similarly, in a meta-analysis of the spatial price transmission literature, Mengel and Von Cramon-Taubadel (2014) find that geographic distance and separation by an international border affect the likelihood of cointegration between the prices in two markets. According to their results, the likelihood of cointegration is 23% lower if an international border separates the markets in question. Furthermore, each additional 1 000 kilometres of distance between two markets within a country decreases the likelihood of cointegration by 7%.

In their study, Mengel and Von Cramon-Taubadel (2014) used meta-analysis to distil these estimates of border and distance effects from the extensive empirical literature on price transmission. However, meta-analysis is made difficult by the fact that price transmission studies use different econometric specifications, estimation approaches and types of data (frequency, spatial aggregation). Hence, the results of different studies are not always directly comparable. Furthermore, meta-analysis is susceptible to publication bias, and to the often incomplete documentation of methods and results in published studies. Hence, the aim of this paper was to complement the meta-analysis in Mengel and Von Cramon-Taubadel (2014) by generating own empirical estimates of distance and border effects on price transmission. To this end we tested for the presence of distance and border effects on the transmission of rice prices between markets in Western Africa.

4. Data and methods

We employed 27 monthly price series for imported rice in Benin, Mauritania, Niger, Chad, Senegal and Togo. We considered only prices for imported rice, because research has demonstrated that local and imported rice varieties are not close substitutes (Demont *et al*. 2013a; 2013b). Hence, if we included prices for local rice varieties in our analysis we would risk confounding border and distance effects with the influence of product heterogeneity and imperfect substitution on price transmission. We restricted the analysis to series with at least 100 observations and less than 10% missing values. The price series are taken from the Global Information and Early Warning System (GIEWS) of the Food and Agricultural Organization of the United Nations ([FAO] 2016) and the Famine Early Warning Systems Network database (FEWS) of the United States Department of Agriculture ([USDA] 2013). Most of the series start in the early 2000s and end in 2012 or 2013. The markets considered were retail markets. To ensure comparability, we converted all series to the West African Franc (XOF) per kilogram (kg).[1]

We linearly interpolated all missing values. After interpolation, each series included at least 105 and at most 222 observations (eight to 18 years of monthly observations). Geographically, the series cover three cities in Benin, two in Chad, one in Mauritania, five in Niger, four in Senegal and six in Togo. There is more than one price series for some markets. In such cases, all series were considered and numbered accordingly, e.g. Moussoro 1 and Moussoro 2. Several West African countries were not included in the estimation due to non-existing or incomplete data. The Augmented Dickey Fuller (ADF) test (Dickey & Fuller 1979) fails to reject the null hypothesis of a unit root in all of the price series in levels, but rejects this null hypothesis for all of the series in first differences (Table 2). The 27 price series were combined to form $[(27^2-27)/2] = 351$ price pairs for the subsequent analysis. The series start between May 1995 and October 2003 and end between April 2012 and November 2013. The price series are trimmed as necessary to the time period for which they overlap.

[1] Communauté française d'Afrique (CFA) franc is the name of two currencies used in West Africa and Central Africa. XOF is the West African currency and is used in Benin, Burkina Faso, Côte d'Ivoire, Guinea-Bissau, Mali, Niger, Senegal and Togo.

Table 2: Descriptive statistics and unit-root tests of price series

Series label	Country	Start	End	Number of missing values (share in percent)	ADF statistic (price level)	ADF statistic (price difference)	Mean (XOF/kg)	Coefficient of variation
Abomey	Benin (BEN)	Aug 95	Oct 13	15 (6.8)	0.75	-11.96	358.71	23.38
Cotonou		May 95	Oct 13	-	1.20	-13.10	388.11	22.10
Natitingou		May 95	Oct 13	8 (3.6)	0.89	-11.56	384.77	22.29
Moussoro 1	Chad (TCD)	Oct 03	Oct 13	-	0.10	-13.07	531.89	19.34
N'Djamena 1		Oct 03	Oct 13	-	-0.17	-11.07	463.24	15.66
Moussoro 2		Jan 02	Jun 13	-	-0.10	-12.25	526.53	19.86
N'Djamena 2		Jan 02	Jun 13	-	-0.06	-12.06	403.12	18.66
Nouakchott 1	Mauritania (MRT)	Oct 03	Oct 13	7 (5.8)	0.09	-7.16	404.14	16.64
Nouakchott 2		Apr 03	Jun 13	5 (4.1)	0.16	-9.53	376.61	12.76
Agadez 1	Niger (NER)	May 95	Apr 12	-	1.20	-11.58	353.74	20.54
Dosso		May 95	Apr 12	4 (2.0)	0.97	-11.31	339.44	18.67
Maradi		May 95	Apr 12	-	1.21	-11.06	348.82	25.08
Niamey 1		Jan 00	Apr 12	-	0.73	-9.19	349.89	20.06
Zinder 1		Jan 00	Apr 12	-	1.55	-9.68	368.23	19.66
Agadez 2		Jan 02	Jun 13	1 (0.7)	1.02	-8.13	390.27	19.52
Niamey 2		Oct 03	Jun 13	-	0.68	-9.01	381.87	16.77
Zinder 2		Oct 03	Jun 13	-	1.34	-9.14	408.44	17.58
Dakar	Senegal (SEN)	Oct 03	Jun 13	-	-0.09	-9.31	277.14	23.19
Kaolack		Oct 03	Jun 13	-	-0.07	-8.46	270.96	21.51
St. Louis		Oct 03	Jun 13	-	-0.13	-7.56	247.14	33.3
Ziguinchor		Oct 03	Jun 13	-	-0.04	-7.07	279.89	21.92
Amegnran	Togo (TGO)	Jan 01	Nov 13	-	0.44	-13.02	399.37	29.08
Anie		Jan 01	Nov 13	-	-0.01	-11.65	354.35	20.27
Cinkasse		Jan 01	Nov 13	-	0.47	-11.83	361.67	20.45
Kara		Jan 01	Nov 13	-	0.24	-11.77	366.24	17.66
Korbongou		Jan 01	Nov 13	3 (1.9%)	0.35	-12.23	347.62	20.11
Lomé		Jan 01	Nov 13	-	-0.05	-10.77	387.84	36.71

Note: The 5% critical augmented Dickey-Fuller (ADF) value for the null hypothesis of a unit root is -1.95.

The means of the rice price series vary between 247 and 532 West African Franc (XOF) per kilogram (kg), about 0.38 to 0.81 EUR. The coefficients of variation were measured between 12.76 and 36.71. In the first estimation step of our analysis we applied a series of tests for linear and threshold cointegration to each price pair (Table 3).[2] The testing procedure first distinguishes between no cointegration and linear cointegration using the Johansen (1988) test. If a pair fails to reject the null hypothesis of no cointegration, a test proposed by Seo (2006) is used to distinguish between no cointegration and threshold cointegration. Price pairs for which the Johansen test rejects the null hypothesis of no cointegration were subjected to tests proposed by Hansen and Seo (2002) and Larsen (2012) that distinguish between linear cointegration and different forms of threshold cointegration. This testing procedure follows Stigler (2013). The critical value for tests was chosen at the 5% level of significance. In all tests, the number of lags was selected according to the Bayesian information criterion. Based on this sequence of tests, each price pair could be allocated to one of three categories based on its cointegration status: i) no cointegration; ii) linear cointegration, and iii) threshold cointegration. The linear and nonlinear cointegration test procedure was run twice for each market pair so that each market was once the first component and once the second component in these four test equations. The results are not necessarily symmetrical, as test results can differ in finite samples

[2] As Gonzalo and Pitarakis (2006) point out, the term "threshold cointegration" is misleading in this context because we are referring to linearity vs. threshold effects in the error correction mechanism that maintains the long-run cointegrating relationship between two prices, and not linearity vs. threshold effects in this cointegrating relationship itself.

depending on normalisation (i.e. the order in which the two prices in a pair are entered into the test procedure). For the 26 price pairs for which the tests produced conflicting results we decided in favour of threshold over linear cointegration, and linear over no cointegration.

Table 3: Linear and nonlinear cointegration tests applied to the West African rice prices

Test	Null hypothesis	Alternative hypothesis
Johansen (1988)	no cointegration	linear cointegration
Seo (2006)	no cointegration	threshold cointegration
Hansen & Seo (2002)	linear cointegration	threshold cointegration (1 threshold, 2 regimes)
Larsen (2012)	linear cointegration	threshold cointegration (2 thresholds, 3 regimes)

Note: The tests were implemented with the statistical software R and the R-packages urca and tsDyn, and the test results were obtained at the 5% level of significance.

In the second step of our analysis we estimated a multinomial logistic regression in which the dependent variable was a qualitative indicator of a price pair's cointegration status as estimated above, and the independent variables were measures of proximity between the prices in questions. Hence, we estimated the following model:

$$CS_{ij} = f(Proximity_{ij}) \qquad (1)$$

where CS_{ij} is the cointegration status of the prices in markets i and j (no cointegration, linear cointegration or threshold cointegration), and $Proximity_{ij}$ is the vector of the covariates that measure various dimensions of the proximity between markets *i* and *j*.[3]

The specification of the covariates in $Proximity_{ij}$ is based on the theoretical considerations discussed above (section 3), and specifically the literature on the influence of distance and borders on trade volumes and price dispersion. We used both great circle and road distance in 1 000 km to measure the physical distance between two markets. Great circle distance is the shortest distance between two points, while road distance refers to the minimum distance using existing roads. These distances were obtained from an online distance calculator (www.distance.to) based on Google maps API (https://developers.google.com/maps). The covariates border, contiguity and language were obtained from Mayer and Zignago (2011), and are usually referred to as the CEPII dataset. A border dummy takes the value 0 for market pairs located in the same country and 1 otherwise, and for a subset of those cases for which border = 1, the contiguity dummy takes the value 1 if the countries in question share a common border. Keeping all other factors constant, we expected that more international borders between two markets would create more impediments to trade, information flows and price co-movement. As a measure of cultural proximity we included a dummy variable that equalled 1 if the countries in question shared a common official language. Note that this variable might also reflect the effect of a common colonial history.

Table 4 presents the descriptive statistics for the variables used in our estimation. A total of 35 (10.0%) of the price pairs in our sample are linearly cointegrated, and 122 (35%) are linked by threshold cointegration. The remaining 194 price pairs (55%) are not cointegrated.[4]

[3] An ordered logistic model would not be appropriate, as the three cointegration status categories are not ordered.
[4] More detailed results of the cointegration tests are provided further on.

Table 4: Summary statistics of market pair variables

	Full sample (N = 351)				Only pairs ≤ 1 000 km distance (N = 144)			
Variable	Mean	St. dev.	Min.	Max.	Mean	St. dev.	Min.	Max.
Linear cointegration	0.100	0.300	0	1	0.194	0.397	0	1
Threshold cointegration	0.348	0.477	0	1	0.451	0.499	0	1
1 000 km great circle distance	1.430	0.957	0.000	3.660	0.508	0.290	0.000	0.996
1 000 km road distance	1.930	1.290	0.000	4.920	0.703	0.427	0.000	1.700
Border (yes = 1, otherwise 0)	0.832	0.374	0	1	0.590	0.493	0	1
Contiguity (yes = 1, otherwise 0)	0.234	0.424	0	1	0.403	0.492	0	1
Same language (yes = 1, otherwise 0)	**0.758**	**0.429**	**0**	**1**	**0.535**	**0.501**	**0**	**1**

Note: Great circle distance and road distance are measured with the online distance calculator www.distance.to (German version). Contiguity and same official or ethnological language were obtained from Mayer and Zignago (2011).

The average great circle distance between market pairs is 1 430 km (maximum 3 660 km). Road distance between each two markets is on average 1 930 km and maximum distance is 4 920 km. A total of 83.2% of the market pairs are separated by an international border, i.e. are not located in the same country, while 23.4% of market pairs lie in contiguous countries. In 75.8% of the cases, the same language is spoken in both markets. To better illustrate the features of the dataset, we drew a subset with only market pairs that were at most 1 000 km distant from each other (Table 4, columns 6 to 9). Of these 144 observations, 19.4% were linearly cointegrated, almost twice as many as in the full dataset. At 45.1%, threshold cointegration was also relatively more frequent in the truncated data compared to the full sample. This already hints at our hypothesis, which postulated a more frequent price transmission (cointegration) the closer two markets are. The average great circle distance is 508 km (996 km maximum). Road distance is on average 703 km. The furthest road distance between two market pairs in the sample is 1 700 km. A total of 59% of the pairs are separated by a border, i.e. are not located in the same country. Most of them, viz. 40.3% of the sample, are located in contiguous countries. In 53.5% of the cases, the same language is spoken in both markets.

Table 5: Correlation matrix of dependent and independent variables

	Linear cointegration	**Threshold cointegration**	**Great circle distance**	**Road distance**	**Border**	**Contiguity**	**Same language**
Linear cointegration	1.000	-0.243	-0.249	-0.250	-0.206	0.086	-0.123
Threshold cointegration	-0.243	1.000	-0.181	-0.183	-0.120	0.177	0.008
Great circle distance	-0.249	-0.181	1.000	0.994	0.507	-0.372	0.409
Road distance	-0.250	-0.183	0.994	1.000	0.514	-0.360	0.409
Border	-0.206	-0.120	0.507	0.514	1.000	0.248	0.795
Contiguity	0.086	0.177	-0.372	-0.360	0.248	1.000	0.186
Same language	-0.123	0.008	0.409	0.409	0.795	0.186	1.000

Note: The correlation matrix shows correlation coefficients for each combination of covariates.

There is multicollinearity in this set of covariates (Table 5). The two measures of distance (great circle and road) are correlated (0.994). Moreover, market pairs that are separated by one or more borders will tend to be farther apart, leading to multicollinearity between the distance measures and the border dummy (0.507 and 0.514). Finally, the language variable was measured at the country level (Mayer & Zignago 2011) and consequently takes the same value (language = 1) for all domestic market pairs (border = 0). Consequently, the two variables are highly correlated. We estimated different versions of equation (1) with subsets of the covariates in Table 4 to isolate robust effects.

5. Results

Table 6 presents detailed results of the first-step cointegration tests. We can see that, in some countries such as Mauritania and Senegal, comparatively few markets are cointegrated with other markets. For markets in Chad, however, the prevalence of cointegration is higher. In Benin, all possible domestic market pairs display some form of cointegration, and the same is true in Senegal and in Chad. There is one domestic market pair in Mauritania, and the prices from these markets are not cointegrated. While no relationship between borders and cointegration is immediately apparent in Table 6, Figure 2 shows that, on the whole, linear and threshold cointegration are more prevalent between markets that are located close to one another (especially in the western part of the sample region).

Table 6: The prevalence of linear and threshold cointegration between prices by country

Country A	Country B	Linear cointegration	Threshold cointegration	No cointegration	Total
Benin (all pairs)		***9***	***29***	***37***	***75***
of which pairs with:	**Benin**	1	2	0	3
	Mauritania	0	0	6	6
	Niger	5	7	12	24
	Senegal	0	0	12	12
	Chad	0	12	0	12
	Togo	3	8	7	18
Mauritania (all pairs)		***0***	***1***	***23***	***24***
of which pairs with:	**Mauritania**	0	0	1	1
	Niger	0	0	11	11
	Senegal	0	0	4	4
	Chad	0	0	2	2
	Togo	0	1	5	6
Niger (all pairs)		***9***	***26***	***58***	***93***
of which pairs with:	**Mauritania**	0	0	5	5
	Niger	6	12	10	28
	Senegal	0	0	20	20
	Chad	2	6	2	10
	Togo	1	8	21	30
Senegal (all pairs)		***1***	***10***	***19***	***30***
of which pairs with:	**Mauritania**	0	0	4	4
	Niger	0	0	12	12
	Senegal	1	5	0	6
	Chad	0	5	3	8
Chad (all pairs)		***5***	***39***	***10***	***54***
of which pairs with:	**Mauritania**	0	2	4	6
	Niger	2	20	0	22
	Senegal	0	5	3	8
	Chad	2	4	0	6
	Togo	1	8	3	12
Togo (all pairs)		***11***	***17***	***47***	***75***
of which pairs with:	**Mauritania**	0	0	6	6
	Niger	4	4	10	18
	Senegal	2	1	21	24
	Chad	1	7	4	12
	Togo	4	5	6	15
All pairs		**35**	**122**	**194**	**351**

Note: The numbers refer to whether the two price series tested positively for linear cointegration or threshold cointegration, or whether the series are not cointegrated according to the testing procedure.

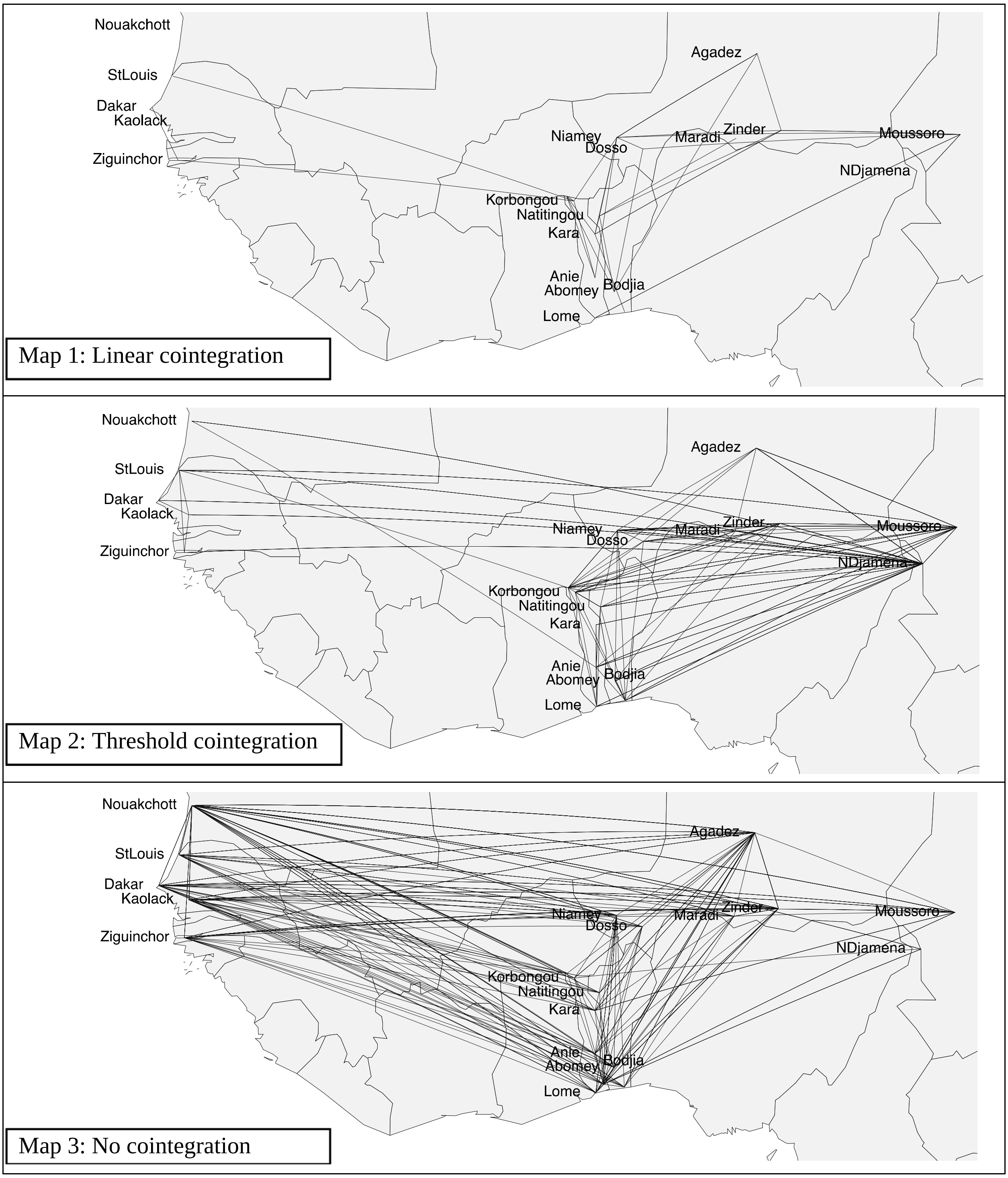

Figure 1: Graphical depiction of linear and nonlinear cointegration test results for geographically separated market pairs

Note: The maps present the estimated statistical relationships between rice prices in each two markets. In map 1, a line connects markets that are linearly cointegrated. Maps 2 and 3 display threshold cointegration and no cointegration respectively. Labels for Cinkasse, Amegnran and Cotonou are omitted for clarity.

Tables 7 and 8 present the results of the second step in our analysis, which was the estimation of equation (1). Table 7 presents results based on the entire sample of 351 price pairs, while Table 8 presents results based on the sub-sample of 292 market pairs that are separated by at least one international border. We repeated the analysis using this sub-sample because the variables for proximity (language) for one of the dummies always equal 1 for domestic market pairs[5] and we wished to control for any effect that this multicollinearity might have on our estimates. In addition, we included country dummies to control for effects that might have to do with country-specific properties. We report estimated marginal effects in both tables. Since the reference category for the dependent variable is "no cointegration", these marginal effects can be interpreted as proportional changes in the likelihood that a pair of prices will be linearly (or threshold) cointegrated due to a one-unit change in the corresponding covariate. Hence, a positive marginal effect indicates that the covariate in question increases the likelihood of one or the other form of cointegration, and thus of price transmission.

Table 7: Marginal effects according to multinomial logistic estimation with all market pairs (n= 351)

Model →	(1)	(1a)	(2)	(2a)	(3)	(4)
Linear cointegration: 1000 km great circle distance	-0.099*** (0.024)	-0.090*** (0.032)				
Threshold cointegration: 1000 km great circle distance	-0.074*** (0.026)	-0.114*** (0.035)				
Linear cointegration: 1000 km road distance			-0.074*** (0.018)	-0.071*** (0.025)		
Threshold cointegration: 1000 km road distance			-0.056*** (0.020)	-0.088*** (0.026)		
Linear cointegration: border					-0.165*** (0.057)	-0.215*** (0.068)
Threshold cointegration: border					-0.153** (0.071)	-0.169*** (0.076)
Linear cointegration: contiguity						0.118** (0.052)
Threshold cointegration: contiguity						0.209*** (0.065)
Country dummies	No	Yes	No	Yes	No	No
Pseudo-R^2	0.076	0.298	0.077	0.300	0.037	0.086
Log likelihood	-299.87	-227.92	-299.72	-227.27	-312.59	-296.89
LR chi²	49.55*** (df = 2)	193.44*** (df = 12)	49.58*** (df = 2)	194.75*** (df = 12)	24.11*** (df = 2)	55.51*** (df = 4)

Note: The reference category for the dependent variable is no cointegration. Standard errors of the estimated marginal effects are in brackets. *, ** and *** refer to significance at the 10%, 5% and 1% levels respectively. Detailed results for the country dummies are available from the authors.

Table 7 shows that distance has a negative effect on the likelihood of linear cointegration, and a similar but somewhat weaker effect on the likelihood of threshold error correction. The magnitudes of these effects are somewhat stronger for great circle than for road distance. Specifically, for each additional 1 000 km of great circle distance between two markets, the likelihood that the prices in these markets are linearly (threshold) cointegrated falls by 9.9% (7.6%) (model 1). The inclusion of country dummies changes these effects to 9.0% and 11.4% respectively (model 1a). Models 2 and 2a are similar to 1 and 1a but use road distance rather than great circle distance. Their explanatory power is slightly higher, which could be attributed to a more precise measure of proximity. Table 7 also shows that the likelihood of linear (threshold) cointegration falls by 16.5% (15.3%) if an international border separates the markets in question (model 3). The magnitude of this border effect increases to

[5] The bivariate covariates were obtained at the country level. By definition, therefore, these variables take the same value for two markets that are located in the same country.

21.5% (16.9%) if we control for contiguity of the countries, while contiguity itself increases the likelihood of linear (threshold) cointegration by 11.8% (20.9%) (model 4). Comparing the various estimates of border and of distance effects in Table 7 indicates that a border has approximately the same effect on the probability of cointegration as 1 500 to 3 000 km of distance, which is similar in magnitude to the border effect on price dispersion estimated by Engel and Rogers (1995).

The results in Table 8, which are based on the sub-sample of market pairs that are separated by international borders, confirm that the likelihood of cointegration falls with increasing distance between markets. The estimated effects of additional distance (road or great circle) on cointegration are similar in Tables 6 and 7 (compare models 1 and 1a with 5 and 5a, and 2 and 2a with 6 and 6a), which suggests that this effect is estimated robustly. A comparison of Table 8 (model 7) and Table 7 (model 4) indicates that the positive effect of contiguity on cointegration is also robustly estimated. However, including covariates for both distance and contiguity (Table 8, model 7a) leads to insignificant estimates of the effect of distance on threshold cointegration, and of contiguity on linear cointegration. This might be due to the multicollinearity between distance and contiguity, as markets separated by a common border will generally be closer to one another than markets separated by more than one border. Finally, a common language has significant and positive effects on the likelihood of linear (7.9%) and threshold (31.1%) cointegration (Table 8, model 8). These common language effects do not change when distance is included in the estimation (Table 8, model 8a). For robustness checks we tested for effects specific to landlocked countries and for coastal towns. The results did not change substantially and can be obtained from the authors.

Table 8: Marginal effects according to multinomial logistic estimation with the sub-sample of international market pairs (n = 292)

Model →	**(5)**	**(5a)**	**(6)**	**(6a)**	**(7)**	**(7a)**	**(8)**	**(8a)**
Linear cointegration: 1 000 km great circle distance	-0.072*** (0.023)	-0.074* (0.041)				-0.064** (0.027)		-0.074*** (0.024)
Threshold cointegration: 1 000 km great circle distance	-0.078*** (0.030)	-0.068 (0.042)				0.006 (0.038)		-0.076*** (0.029)
Linear cointegration: 1 000 km road distance			-0.055*** (0.018)	-0.061* (0.032)				
Threshold cointegration: 1 000 km road distance			-0.058*** (0.022)	-0.053* (0.031)				
Linear cointegration: contiguity					0.103** (0.041)	-0.023 (0.037)		
Threshold cointegration: contiguity					0.248*** (0.063)	0.242*** (0.082)		
Linear cointegration: same language							0.079*** (0.017)	0.079*** (0.016)
Threshold cointegration: same language							0.311*** (0.048)	0.311*** (0.047)
Country dummies	No	Yes	No	Yes	No	No	No	No
Pseudo-R^2	0.056	0.347	0.057	0.350	0.063	0.078	0.040	0.100
Log likelihood	-236.35	-163.62	-236.06	-162.87	-234.73	-230.84	-240.35	-225.51
LR chi^2 (2)	28.15*** (df = 2)	173.63*** (df = 12)	28.74*** (df = 2)	175.11*** (df = 12)	31.40*** (df = 2)	39.19*** (df = 4)	20.15*** (df = 2)	49.84*** (df = 4)

Note: The reference category for the dependent variable is no cointegration. Standard errors of the estimated marginal effects are in brackets. *, ** and *** refer to significance at the 10%, 5% and 1% levels respectively. Detailed results for the country dummies are available from the authors.

6. Discussion

The results presented here confirm that various dimensions of proximity affect not only trade volumes and price dispersion, but also price transmission. Geographic distance, borders, contiguity and a common language have systematic and expected effects on the likelihood of cointegration between prices in two spatially separate markets. The closer two markets are in terms of distance, the more likely it is that price changes are transmitted between them. If the markets are located in the same country or in contingent countries, the likelihood also increases, as well as when both markets share a common language. The size and direction of the effects are expected in accordance with previous research on trade volumes and price dispersion.

The results contribute to an improved understanding of the determinants of price and market integration in the region. Based on this, welfare-enhancing policies may address investments in infrastructure and measures to facilitate trade. In addition, policies addressing rice prices in the region may take into account their effect on neighbouring markets and, to a much lesser extent, on adjacent countries.

Several caveats should be mentioned. Correlation between covariates, for example between distance and the presence of a border, can lead to multicollinearity and make it difficult to disentangle partial effects. There is also a danger of selection bias if isolated, less-integrated markets are less likely to be included in international price datasets such as FAO-GIEWS (FAO 2016). If countries with better infrastructure and institutions are more likely to be included in such datasets, we might, for example, systematically underestimate the effect of distance on price transmission. For both of these reasons, extending the analysis to a broader set of countries therefore would be a worthwhile endeavour in the future.

Future research could study the effect of distance, not only on the likelihood of cointegration, but also on the speed of price adjustment. In addition, it might be possible to include covariates that capture different types of border policies between countries in future analyses of this nature. Ad valorem tariffs will have different effects on price co-movement than, for example, import quotas. However, to the extent that border policies have changed over time, especially in the recent turbulent years of high and volatile prices, defining variables that adequately capture border policies will likely prove difficult.

Acknowledgements

The authors gratefully acknowledge funding by the Research Training Group "GlobalFood – Transformation of Global Agri-Food Systems" and the German Research Foundation (DFG). The corresponding author thanks the International Food Policy Research Institute (IFPRI), Dakar, Senegal for providing facilities, and the IFPRI researchers for useful comments and feedback, in particular by Fousseini Traoré, Anatole Goundan and Tanguy Bernard.

References

Aker J, Klein M & Connell S, 2013. Borders, ethnicity and trade. Tufts University Working Paper Series: 1–72. Available at: http://sites.tufts.edu/jennyaker/files/2010/02/Niger_Border-March-2013.pdf (Accessed 1 February 2015).

Anderson J & Van Wincoop E, 2004. Trade costs. Journal of Economic Literature 42(3): 691–751.

Ardeni G, 1989. Does the law of one price really hold? American Journal of Agricultural Economics 71: 661–9.

Demont M, Rutsaert P, Ndour M, Verbeke W, Seck PA & Tollens E, 2013a. Experimental auctions, collective induction and choice shift: willingness-to-pay for rice quality in Senegal. European Review of Agricultural Economics 40: 1–26.
Demont M, Rutsaert P, Ndour M & Verbeke W, 2013b. Reversing urban bias in African rice markets: Evidence from Senegal. World Development 45: 63–74.
Dickey D & Fuller W, 1979. Distribution of the estimators for autoregressive time series with a unit root. Journal of the American Statistical Association 74(366): 427–31.
Diouf E, 2012. West Africa's new common external tariff and the individual WTO commitments of ECOWAS member states: No insurmountable incompatibilities. Bridges Africa 1(5): 7–9.
ECOWAS, UEMOA & NEPAD, 2014. Support for the implementation of the regional offensive for sustainable and sustained recovery of rice production in West Africa Programme. Available at http://www.hubrural.org/IMG/pdf/releve_de_conclusion_tasf_rice_cotonou_mars_2014_en_.pdf
Engel C & Rogers J, 1995. How wide is the border? International Finance Discussion Papers 498: 1–40.
Fackler PL & Goodwin BK, 2001. Spatial price analysis. In Gardner BL & Rausser GC (eds.), Handbook of Agricultural Economics, Volume 1, Part B. Amsterdam: Elsevier.
Fiamohe R, Alia D, Bamba I & Amovin Assagba E, 2015. Transmission of rice prices from Thailand into West African markets: The case of Benin, Mali, and Senegal. Journal of African Business 16(1-2): 128–43.
Food and Agricultural Organization of the United Nations (FAO), 2016. Global Information and Early Warning System (GIEWS database). Rome, FAO. Available at http://www.fao.org/giews/pricetool (Accessed 1 February 2015).
Gonzalo J & Pitarakis JY, 2006. Threshold effects in cointegrating relationships. Oxford Bulletin of Economics and Statistics 68: 813–33.
Greb F, Von Cramon-Taubadel S, Krivobokova T & Munk A, 2013. The estimation of threshold models in price transmission analysis. American Journal of Agricultural Economics 95(4): 900–16.
Hansen B & Seo B, 2002. Testing for two-regime threshold cointegration in vector error-correction models. Journal of Econometrics 110(2): 293–318.
Hernandez-Villafuerte K, 2011. The relationship between spatial integration and geographical distance in Brazil. Paper read at the Annual Meeting of the European Association of Agricultural Economics (EAAE), 30 August–2 September, Zürich, Switzerland.
Jensen R, 2007. The digital provide: Information (technology), market performance, and welfare in the South Indian fisheries sector. The Quarterly Journal of Economics CXXII(3): 879–924.
Johansen S, 1988. Statistical analysis of cointegration vectors. Journal of Economic Dynamics and Control 12: 231–54.
Larsen B, 2012. A threshold cointegration analysis of Norwegian interest rates. Master's thesis in Statistics, University of Tromsø, Norway.
Mayer T & Zignago S, 2011. Notes on CEPII's distances measures: The GeoDist database. CEPII Working Paper Series 2011-25, CEPII, Paris.
Mengel C & Von Cramon-Taubadel S, 2014. Distance and border effects on price transmission – a meta-analysis. GlobalFood Discussion Papers 37: 1–23.
Seo M, 2006. Bootstrap testing for the null of no cointegration in a threshold vector error correction model. Journal of Econometrics 134: 129–50.
Stephens E, Mabaya E, Von Cramon-Taubadel S & Barrett CB, 2012. Spatial price adjustment with and without trade. Oxford Bulletin of Economics and Statistics 74(3): 453–69.
Stigler M, 2013. Threshold cointegration: Overview and implementation in R. Available at cran.r-project.org/web/packages/tsDyn/vignettes/ThCointOverview.pdf
United Nations, 2010. UN Comtrade database. New York, United Nations. Available at http://comtrade.un.org (Accessed 1 February 2015).

United States Department of Agriculture (USDA), 2013. Famine Early Warning System (FEWS database). Washington, USA.

World Bank, 2009. Eastern Africa: A study of the regional maize market and marketing costs. Report No. 49831. Agriculture and Rural Development Unit, Washington DC.

Zannou A, 2010. Determinants of intra-ECOWAS trade flows. African Journal of Business Management 4: 678–86.