

Investigating Integration and Exchange Rate Pass-Through in World Maize Markets Using Post-LASSO Inference

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Abstract

This paper investigates the extent of market integration and exchange rate pass-through, as well as market factors that may be associated with deviations from perfect market integration and pass-through. To address the shortcomings of existing models of spatial market integration, we employ procedures outlined in [Yan \(2023\)](#) for inference and model selection, utilizing the debiased (or desparsified) LASSO method for high-dimensional threshold regression. Our results support the integration of global corn markets, especially when accounting for the existence of thresholds. We identify significant relationships among several variables representing domestic and world economic conditions.

Keywords: Law of One Price, Threshold Regression Model, Exchange Rate Pass-Through

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1 Introduction

Efficient markets are expected to eliminate any potential for riskless profits through arbitrage and trade, a result known as the "Law of One Price" (LOP). Economic arbitrage relies on the principle that prices of related goods should not arbitrarily differ from one another over the long run. The general implication here is that prices for homogeneous products at different geographic locations in otherwise freely functioning markets should differ by no more than transport and transactions costs. However, the existence of transactions costs can introduce a threshold effect, where deviations in prices above a certain threshold are necessary to trigger price movements. In recent years, studies analyzing this phenomenon have focused on developing nonlinear models that can better capture the effects of unobservable transaction costs in spatial price linkages. The motivation behind using such models is to better understand the dynamics of market integration and the role of transaction costs in the presence of regime changes. The use of nonlinear models has been largely driven by the application of threshold modeling techniques. These models are based on the idea that transaction costs and other barriers to spatial trade may lead to regime switching, with alternative regimes representing the trade and no-trade equilibria. This idea has been operationalized through various econometric techniques and model specifications.

Threshold autoregression (TAR) models have indeed had a significant impact on the analysis of asymmetric price transmission in agricultural economics. These models have been developed to capture the nonlinear dynamics of market integration and account for the effects of unobserved transaction costs that can affect spatial price linkages. A common approach to threshold modeling often involves an autoregressive model of the price differential. The study conducted by [Goodwin and Piggott \(2001\)](#) examined corn prices at local markets by combining a threshold structure with an error-correction model. [Goodwin et al. \(1990\)](#) noted that delivery lags that extend beyond a single time period may imply arbitrage conditions that involve noncontemporaneous price linkages. Based on this idea, [Lence et al. \(2018\)](#) examined the performance of the threshold cointegration approach, specifically Band-TVECM, in analyzing price transmission in an explicit context where trade decisions are made based on the expectation of final prices because trade takes time. In addition to

the threshold model, [Goodwin et al. \(2021\)](#) applied generalized additive models to empirical considerations of price transmission and spatial market integration.

The extensive literature addressing price transmission and market integration has largely focused on simple comparisons of prices in geographically distinct markets. In rare cases when the consideration applies to international market comparisons, exchange rates may also be considered. It is, however, possible that a number of other variables may be relevant to market linkages. These variables may largely proxy for unobservable transactions costs or other reasons for price differences, such as government policies and product heterogeneity. The puzzle confronting the analyst is which of these variables makes valid and useful proxies for the unobservable reasons as to why prices may differ. We attempt to address this puzzle by considering data-driven methods for selecting the optimal set of variables relevant to an understanding of price linkages in international maize markets. Specifically, we apply a high dimensional threshold model to examine the effect of exchange rates and market factors on price linkages among spatially distinct world maize markets. Such an application is a natural methodological extension of existing empirical studies on spatial market integration models.

Although exchange-rate pass-through, i.e. the degree to which exchange rate movements are reflected in prices has long been a question of interest in international economics, there is limited literature that examines exchange-rate pass-through in global agricultural commodity markets. One study by [Chambers and Just \(1981\)](#) uses an econometric model of the wheat, corn, and soybean markets to investigate the dynamic effects of exchange rate fluctuations on U.S. commodity markets. The study finds that exchange rate fluctuations have a significant real impact on agricultural markets, particularly on the volume of exports and the relative split between exports and domestic use of these commodities. The econometric model developed in the study shows that agricultural prices are sensitive to movements in the exchange rate, with short-run adjustments being more dramatic than longer-run adjustments. [Varangis and Duncan \(1993\)](#) employed a system of equations, including one for the Japanese export price and another for the US producer price, to estimate the impact of changes in the yen/dollar exchange rate and other factors on steel prices. It is worth noting that this trade is unidirectional, as Japanese steel is consistently imported into the USA. The study demonstrates that such exchange rate fluctuations

are not fully passed through to steel prices. The pricing decisions of US steel producers are primarily influenced by changes in their production costs and the US index of industrial production, rather than fluctuations in the yen/dollar exchange rate.

International trade in basic commodities is generally invoiced in US dollar terms. At first glance, this may seem to imply that exchange rates are irrelevant to market linkages. However, assuming that the commodities are valued in local currencies after being imported suggests that exchange rates may still be relevant to price linkages. We discuss this point in greater detail below.

[Barrett and Li \(2002\)](#) examine actual trade flows as a factor for assessing spatial market integration. They note that empirical tests should differentiate between the notions of spatial market integration and a competitive market equilibrium. The latter concept refers to market conditions where no trade occurs because arbitrage conditions do not provide opportunities for profitable trading. The authors highlight that prices in two segmented markets might react to exogenous factors like inflation or climatic conditions without representing a spatial equilibrium in markets. A recent overview from the World Bank [Rebello \(2020\)](#) addresses the factors influencing spatial market integration. The overview mentions the cooperation among policy-makers on matters such as trade and investment policies, migration, transportation infrastructure, macroeconomic policy, natural resource policy, and others related to "shared sovereignty." Furthermore, the overview highlights the critical role of regional integration in policy reforms, contributing significantly to overall peace and security.

The integration of world markets for grains and oilseeds has been of interest for many years. In recent years, the global maize market has been dominated by major exporters such as the United States, Argentina, and Ukraine, which have consistently ranked among the top maize producers and exporters worldwide. The US, the largest producer, alone accounts for over one-third of global maize exports. Argentina and Ukraine collectively account for over one-fourth of global maize exports. The dominance of these countries in the global maize market is representative of the market and makes them candidates for studying price transmission and market integration. They play a crucial role in global maize prices and influencing maize markets worldwide. Likewise, the extent to which distortions arise due to incomplete pass-through of exchange rate shocks has been an important indicator of the overall functions of markets.

In addition to prices and exchange rates, other market factors can be conceptually related to market linkages, such as aggregate economic indicators like industrial production, trade policies, and exogenous shocks, such as the recent pandemic, interest rates, and nominal inflation rates in each market. These factors may be associated with deviations from perfect market integration, as they can affect the costs of transportation, communication, and transactions between markets, as well as the demand and supply conditions in each market. Understanding the effects of these market factors on price linkages is essential for policymakers and market participants to make informed decisions about trade, investment, and risk management.

LASSO (least absolute shrinkage and selection operator) is a regression technique that uses shrinkage methods for variable selection. LASSO employs L1 regularization and shrinkage techniques to penalize the model based on the absolute value of parameter estimates. It is a valid approach for identifying an optimal model specification by selecting the variables that contribute the most to explaining a regression-type relationship. Although LASSO models have been widely used in economics studies, the shrinkage bias introduced due to the penalization in the LASSO loss function can affect the properly scaled limiting distribution of the LASSO estimator. Therefore, to conduct valid statistical inference, we need to remove this bias. This paper uses the debaised LASSO method for high dimensional threshold regression, recently developed by [Yan \(2023\)](#) to model the nonlinearity in the spatial price integration models. The fact is that existing literature on price transmission and exchange rate pass-through has developed from simple regression models to nonlinear specifications that allow differential impacts on price linkages. These differential effects are often identified using smooth or discrete threshold models.

2 Econometrics Models of Spatial Market Integration

Spatial market integration in agricultural product markets has been extensively studied in the literature. Consider a commodity traded in common currency in two regional or international markets represented by location indices j and k . The individual market prices are denoted by P^j and P^k , respectively. The arbitrage condition

of perfect market integration reflects the equation $P_t^j/P_t^k = 1$, abstracting from trade and transportation costs. This condition has been adjusted to account for the wedge between prices due to transaction or transportation costs, which may differ significantly in regional markets. The general representation for this adjusted arbitrage condition is $1/(1 - \kappa) \leq P_t^j/P_t^k \leq 1 - \kappa$, where κ represents the proportional loss in commodity value due to transaction or transportation costs ($0 < \kappa < 1$). The greater the distance between locations j and k , the closer κ is to one. It should be noted that many factors may be relevant to price differences across markets. Most existing studies have only considered simple price relationships. An important distinction exists between transportation and transactions costs, which include transport costs as well as other factors that contribute to price differences. These factors could include variables associated with economic and trade policies, product characteristics, and risk.

Many spatial economic models utilize the iceberg trade cost proposed by [Samuelson \(1954\)](#), which assumes that part of the produced output representing the material costs of transportation melts away during transportation. That is, after taking natural logarithms and denoting $p_t^j = \ln P_t^j$, the inequality is often presented as

$$|p_t^j - p_t^k| \leq |\ln(1 - \kappa)|. \quad (2.1)$$

The inequality (2.1) is generally considered to reflect two distinct states of the market. The first state corresponds to a condition where there is no profitable trading, with $|p_t^j - p_t^k| \leq |\ln(1 - \kappa)|$. Under conditions of trade or profitable arbitrage opportunities, the condition holds as $|p_t^j - p_t^k| > |\ln(1 - \kappa)|$. The speed at which the market adjusts to such deviations from the arbitrage equilibrium is often used as a measure of the degree of market integration. Typically, these discrete arbitrage and no-arbitrage conditions are represented using threshold models, where the threshold represents an empirical measure of the transaction cost, $|\ln(1 - \kappa)|$. Bidirectional trade models may allow for different thresholds depending on which market price is higher.

Over time, log price differentials within the band limits are expected to follow a unit root process. Conversely, log price differences outside the band are expected to be mean-reverting, which suggests the existence of a transactions cost band, as

assumed in the literature.

A wide literature has examined spatial market integration in world markets for agricultural commodities. Likewise, a large related literature has examined how shocks to exchange rates affect domestic and export prices, a phenomenon known as ‘pass-through’. If a shock to exchange rates is fully reflected in adjustments to prices, the shock is considered to have been fully passed through. Most empirical studies of market integration and exchange rate pass-through assume a linear relationship, as represented by

$$p_t^j = \alpha + \beta p_t^k + \gamma_{2,0} \pi_t^{jk} + \varepsilon_t, \quad (2.2)$$

where p_t^j is the price in market j in time period t and π_t^{jk} is the exchange rate between currencies in markets j and k , all in logarithmic terms.

Perfect integration is implied if $\alpha = 0$ and $\beta = 1$. In cases where prices are invoiced in different currencies, perfect integration also requires perfect exchange rate pass-through, which is implied if $\gamma_{2,0} = 1$. If prices are invoiced in a common currency, as is often the case when trade is conducted in US dollar terms, the exchange rate is 1 and thus the logarithmic value of zero eliminates the exchange rate effect¹. However, exchange rate distortions may still affect price linkages, which is implied if $\gamma_{2,0} \neq 0$, even if prices are quoted in a common currency. This could occur if imported goods are moved into internal markets in which a different currency applies.

It is also essential to consider other market factors associated with deviations from perfect integration. To this end, we consider an alternative version of equation (2.2) that is expressed as:

$$p_t^j - p_t^k = \gamma_{2,0} \pi_t^{jk} + \gamma_3 Z_t^{jk} + \varepsilon_t, \quad (2.3)$$

where Z_t^{jk} is a set of factors that may be conceptually relevant to price linkage, γ_3 is a vector of parameters corresponding to Z_t^{jk} . These factors include exogenous shocks such as exchange rates, interest rates, unemployment rates, and nominal inflation rates in each of the markets. We do not know which, if any of these factors, is likely to be relevant to price linkages. These factors largely proxy for unobservable factors that may be related to price relationships, such as local policies, product

¹If we define π_t^{jk} as the exchange rate of 1 unit of currency in market j to the currency in market k , a value of $\gamma_{2,0} = -1$ represents perfect exchange rate pass-through.

heterogeneity, and unobservable transactions costs.

To further analyze spatial price linkages, we evaluate deviations from a price parity condition, considering threshold effects of price differentials and isolated shocks in spatially distinct markets. In addition to the conventional specification, we propose an extension to this framework of spatial market integration that includes two regimes. One regime represents a case of no trade, while another represents conditions of profitable trade and arbitrage. The regime switch depends on a forcing variable, usually a lagged price differential, expressed as:

$$\begin{aligned} \Delta(p_t^j - p_t^k) = & \gamma_1(p_{t-1}^k - p_{t-1}^j) + \gamma_{2,0}\Delta\pi_t^{jk} + \gamma_3\Delta Z_t^{jk} \\ & + \mathbf{1}\{|p_{t-1}^j - p_{t-1}^k| \geq c\}(\delta_1(p_{t-1}^j - p_{t-1}^k) + \delta_2\Delta\pi_t^{jk} + \delta_3\Delta Z_t^{jk}) + \varepsilon_t, \end{aligned} \quad (2.4)$$

The parameters γ_1 and δ_1 reflect the degree of market integration. In particular, γ_1 and δ_1 represent the degree of 'error correction' characterizing departures from price parity, which are reflected in large values of $p_{t-1}^j - p_{t-1}^k$. The threshold parameter c represents the amount of proportional transaction costs that a price differential must exceed to cross the threshold and trigger the "trade" regime adjustments. We allow $\delta_1, \delta_{2,0}$ and δ_3 to nonzero according to whether $|p_{t-1}^j - p_{t-1}^k|$ is within (i.e., $|p_{t-1}^j - p_{t-1}^k| < c$) or outside (i.e., $|p_{t-1}^j - p_{t-1}^k| \geq c$) of a symmetric band

Differencing is employed in this study to measure short-run relationships between variables. The first-difference model is utilized to avoid nonstationary variables, allowing a focus on immediate changes between variables. Differencing captures short-run dynamics, while the error correction process reflects longer-run relationships.

To assess the potential presence of transaction costs and other factors affecting price relationships, we consider a multivariate threshold distributed lag model that includes the price differential, exchange rate, and exogenous shocks as well as their

lagged (past period) values, as follows:

$$\begin{aligned}
\Delta(p_t^j - p_t^k) &= \gamma_1(p_{t-1}^j - p_{t-1}^k) + \sum_{l=0}^L \gamma_{2l} \Delta \pi_{t-l}^{jk} + \sum_{l=0}^L \gamma_{3l} \Delta z_{t-l}^{jk} \\
&\quad + \mathbf{1}\{|p_{t-1}^j - p_{t-1}^k| \geq c\} \left[\delta_1(p_{t-1}^j - p_{t-1}^k) + \sum_{l=0}^L \delta_{2l} \Delta \pi_{t-l}^{jk} + \sum_{l=0}^L \delta_{3l} \Delta z_{t-l}^{jk} \right] + \varepsilon_t \\
t &= \{1, \dots, T\},
\end{aligned} \tag{2.5}$$

where L represents the maximum possible lag, which may increase with the sample size, slowly growing to infinity. We assume that the maximal lag order L is known. A distributed lag model ([Almon \(1965\)](#)) is utilized to reveal both short- and long-run dynamic effects between explanatory variables and response variables. Additionally, we employ LASSO, a flexible and supervised learning method. When dealing with time-lagged relationships, selecting the appropriate lag length is crucial in time series modeling. Typically, a well-defined lag length is chosen, and all lags up to that period are included in the model. However, in contexts like ours, where we investigate the dynamic relationship between price linkages, exchange rates, and market factors in agricultural commodities, the delivery time from one market to another spans several weeks to months. Consequently, not all lags are considered equally important in capturing price linkages in response to market shocks. In such scenarios, a distributed lag model (DLM) with lag selection, facilitated by LASSO, proves to be more suitable. LASSO's ability to determine distributed lags through a data-driven search enables a more precise representation of dynamic relationships in agricultural commodity markets. This framework offers a richer evaluation of price dynamics and patterns of adjustment.

Economic agents adjust their expectations of price differentials based on the level of transaction costs that pertain to previous periods. If the price differential exceeds certain thresholds, agents anticipate profitable gains from arbitrage and trade. The specified model offers the advantage of capturing simultaneous relationships between exchange rates and other relevant variables. Linear modeling techniques may not accurately capture the nonlinearities present in the model. Therefore, it is essential to investigate the impact of transaction costs on the market's response to an exchange

rate shock or other market shocks nonlinearly. The existence of different levels of transaction costs can influence how price differentials respond to exchange rates or other shocks, as it determines the presence or absence of arbitrage opportunities. Indeed, a limitation of most existing threshold models of spatial price linkages lies in the typical assumption that transactions costs are constant (in levels or proportional terms). We allow transactions costs, which are inherently unobservable, to vary according to many conceptually relevant economic variables. The proposed model recognizes that the movements in the exchange rate can adjust how markets respond to changes, leading to different regimes based on transaction costs. By considering the effects of transaction costs, we can gain a more comprehensive understanding of the dynamics of the exchange rate pass-through mechanism and the effect of market factors.

The lag coefficients γ_s for $s = 1, \dots, L$ represent the lag distribution and define the pattern of how $\Delta\pi_{t-s}$ or Δz_{t-s} affects $\Delta(p_t^1 - p_t^2)$ over time. The dynamic marginal effect of $\Delta\pi_t$ at the s -th lag is $\frac{\partial\Delta(p_t^j - p_t^k)}{\partial\Delta\pi_{t-s}} = \gamma_{1s}$. The dynamic marginal effect of $\Delta\pi_{t-s}^{jk}$ on $\Delta(p_t^j - p_t^k)$ at the s -th lag is given by $\frac{\partial\Delta(p_t^j - p_t^k)}{\partial\Delta\pi_{t-s}^{jk}} = \gamma_{1s}$. The dynamic marginal effect is essentially an effect of a temporary change in $\Delta\pi_{t-s}^{jk}$ on $\Delta(p_t^j - p_t^k)$, whereas the long-run cumulative effect $\sum_{s=1}^L \gamma_{1s}$ measures how much $\Delta(p_t^j - p_t^k)$ will be changed in response to a permanent change in $\Delta\pi$ when both $\Delta\pi_t$ and $\Delta(p_t^j - p_t^k)$ are stationary. The same derivation can be applied to any element of the vector Δz_{t-s}^{jk} . In the context of the threshold regression model considered here, γ_{1s} and γ_{2s} represent the effect regardless of the status of the forcing variable Q_t , termed the structural effect. On the other hand, δ_{1s}^{jk} and δ_{2s}^{jk} represent the effect when $Q_t > c$, referred to as the threshold effect.

Although a range of economic variables may be conceptually relevant to price linkages, the exact choice of variables and the resulting model specification is unclear. Transactions costs, local policies, and other economic phenomena may affect price linkages between international markets as well as between import and internal markets. Thus, we utilize data-driven methods to select an optimal specification from among a set of potentially relevant variables.

To obtain a specification that incorporates a broad range of variables in (2.5), we utilize a novel approach to inference and model selection: the debiasedLASSO (least absolute shrinkage and selection operator) method for high-dimensional threshold

regression, which was recently developed by Yan (2023). This method allows us to fit the threshold regression models using the threshold LASSO estimator of Lee et al. (2016) in conjunction with the work of van de Geer et al. (2014). Compared to other estimators, this approach can construct asymptotically valid confidence bands for a low-dimensional subset of a high-dimensional parameter vector. Understanding the significance of the estimators can provide insights into the changes in transaction costs and threshold effects over time. However, standard approaches to inference are not applicable to such models.

To simplify, let

$$\alpha = (\gamma_1, \gamma_{20} \cdots, \gamma_{2L}, \gamma_{30} \cdots, \gamma_{3L}, \delta_1, \delta_{20} \cdots, \delta_{2L}, \delta_{30} \cdots, \delta_{3L})'$$

be slope parameter vector, The dimension of α is $2 + 2(1 + p)(L + 1)$, where p is number of other exogenous shocks. Let \mathbf{X} be a $T \times [2 + (1 + p)(L + 1)]$ matrix of all regressors. To provide a more precise description of our estimation procedures, we propose a three-step estimation approach for the model. The three-step procedure can be outlined as follows:

Step 1.

For each $c \in \mathbb{C}$, $\hat{\alpha}(c)$ is defined as

$$\hat{\alpha}(c) := \operatorname{argmin}_{\alpha} \left\{ T^{-1} \sum_{t=1}^T (\Delta(p_t^j - p_t^k) - [X_t', X_t' \mathbf{1}\{Q_t \geq c\}]' \alpha)^2 + \lambda \|\mathbf{D}(c)\alpha\|_1 \right\}, \quad (2.6)$$

where we can rewrite the ℓ_1 penalty as

$$\lambda \|\mathbf{D}(c)\alpha\|_1 = \lambda \sum_{j=1}^{2+(1+p)(L+1)} [\|X^{(j)}\|_n |\alpha^{(j)}| + \|X^{(j)}(\tau)\|_n |\alpha^{(1+(1+p)(L+1)+j)}|],$$

to adjust the penalty differently for each coefficient, depending on the scale normalizing factor. The tuning parameter λ can be selected using either the Akaike Information Criterion (AIC) or the Schwarz Information Criterion (SBC). As the AIC tends to produce less sparse solutions overall, and the SBC applies a stronger penalty on the degrees of freedom, it is more conservative in variable selection compared to

AIC. Therefore, we use AIC for tuning the parameter λ .

Define \hat{c} as the estimate of c_0 such that:

$$\hat{c} := \operatorname{argmin}_{c \in \mathbb{C} \cup \mathbb{R}} \left\{ T^{-1} \sum_{t=1}^T (\Delta(p_t^1 - p_t^2) - [X_t', X_t' \mathbf{1}\{Q_t \geq c\}])' \hat{\alpha}(c) \right\}^2 + \lambda \|\hat{\alpha}(c)\|_1. \quad (2.7)$$

Following [Yan \(2023\)](#), we next turn to variable selection utilizing thresholding. We follow sharp threshold detection techniques provided by [Callot et al. \(2017\)](#) to finding out whether there is a threshold or not, that is, whether

$$(\delta_1, \delta_{20} \cdots, \delta_{2L}, \delta_{30} \cdots, \delta_{3L})'$$

is nonzero or not.

Step 2. We define the thresholded LASSO estimator as

$$\tilde{\delta}_{(j)}(\hat{c}) = \begin{cases} \hat{\delta}_{(j)}(\hat{c}), & \text{if } |\hat{\delta}_{(j)}(\hat{c})| \geq H, \\ 0, & \text{if } |\hat{\delta}_{(j)}(\hat{c})| < H. \end{cases} \quad (2.8)$$

where H is the threshold determining whether a coefficient should be classified as zero or nonzero and $\hat{\delta}_{(j)}(\hat{c})$ are elements of the LASSO estimator defined by (2.6) and (2.7) jointly. In particular, we shall see that choosing $H = 2D\lambda$ yields consistent model selection. The thresholding parameter D can be selected using AIC or SBC through grid search. This ensures that parameters smaller (in absolute value) than $\widehat{D}\widehat{\lambda}$ are set to zero by the thresholded LASSO.

The thresholded LASSO in (2.8) can achieve threshold selection consistency. The consistency of the thresholded LASSO estimator implies that if the underlying true model is nonlinear, then the thresholded LASSO estimator will correctly estimate any of the non-zero parameters, including $(\delta_1, \delta_{20}, \cdots, \delta_{2L}, \delta_{30}, \cdots, \delta_{3L})$, indicating the presence of a nonlinear relationship between the variables. This is in contrast to the conventional ‘self-exciting’ threshold autoregressive (SETAR) model, where nonlinear tests such as Hansen’s modification of standard Chow-type tests, [Tsay \(1989\)](#) linearity test, or neural network tests of linearity are utilized to detect nonlinearity. Therefore, if we misspecify a linear model and use the LASSO method for the threshold model

described here, we may estimate all threshold effects as zero for a sufficiently large sample size. To put it another way, if our estimates of $(\delta_1, \delta_{20}, \dots, \delta_{2L}, \delta_{30}, \dots, \delta_{3L})$ after steps 1 and 2 have at least one non-zero, it indicates that the probability of the model being linear approaches 0. Performing least squares after a sharp threshold detection technique yields inference that is asymptotically equivalent to inference based on least squares with only the relevant variables. However, such inference is of a pointwise nature.

The bias introduced by shrinkage through penalization in the LASSO loss function poses a barrier to developing the limiting distribution of the LASSO estimator. Therefore, to enable statistical inference, an estimation strategy must be employed to eliminate this bias. However, when modeling threshold regression with a rich set of variables, a challenge emerges. Threshold models entail splitting the sample based on a continuously-distributed variable. With a rich set of regressors, there's a risk that the number of observations in any split sample may be less than the number of variables, leading to a reduced-rank sample covariance matrix. Standard approaches are inadequate in such a situation. To debias (desparsify) our LASSO estimator, an approximate inverse of a certain singular sample covariance matrix is needed, as discussed by [van de Geer et al. \(2014\)](#). For a more in-depth exploration and extensions in the case of the LASSO applied to the high-dimensional threshold regression model, detailed information can be found in [Yan \(2023\)](#). However, we do not delve further into these extensions here.

Step 3

Finally, we can obtain debiased LASSO estimates for the threshold model, which is given by:

$$\hat{a}(\hat{c}) = \hat{\alpha}(\hat{c}) + \hat{\Theta}(\hat{c})\mathbf{X}'(\hat{c})(\Delta(p^j - p^k) - \mathbf{X}(\hat{c})\hat{\alpha}(\hat{c}))/n, \quad (2.9)$$

where

$$\hat{\Theta}(\hat{c}) = \begin{bmatrix} \hat{\mathbf{B}}(\hat{c}) & -\hat{\mathbf{B}}(\hat{c}) \\ -\hat{\mathbf{B}}(\hat{c}) & \hat{\mathbf{A}}(\hat{c}) + \hat{\mathbf{B}}(\hat{c}) \end{bmatrix}, \quad (2.10)$$

and $\hat{\mathbf{B}}(\hat{c})$ and $\hat{\mathbf{A}}(\hat{c})$ are the inverse or approximate (if the sample covariance matrix is singular) inverse of the split sample covariance matrices.

3 Empirical Application

The empirical analyses in our study focus on international corn markets, specifically three major exporting markets: the US, Argentina, and Ukraine. Despite its widespread consumption and spatial dispersion, corn production is typically concentrated in specific regions. These three markets collectively accounted for 66.5% of the world corn trade by volume in the 2021/2022 marketing year. Given the intricate spatial dynamics of the corn market, analyzing spatial linkages is crucial for an understanding of the underlying market dynamics and overall performance and behavior.

We collected monthly maize prices and other relevant data from multiple sources which are discussed below. As noted above, the main dependent variable of interest in this study is the maize price in international markets. We collected the yellow corn export prices of the US, Ukraine, and Argentina. Price data for the main three export markets were obtained from the FAO Food Price Monitoring and Analysis (FPMA) Tool, reporting prices in US dollars per metric ton.

Our dataset spans from April 2002 to December 2022, providing 243 monthly observations for each series. However, due to data availability constraints, market factors data for Ukraine is only accessible from April 2002 to February 2022, comprising 239 observations. Similarly, market factors data for Argentina is available from July 2003 to December 2022, encompassing 234 observations. To address missing values, we applied spline interpolation during the relevant periods where prices were missing.²

We collected exchange rates for Ukraine (United States dollar(USD) to Ukrainian Hryvnia(UAH)) and Argentina (USD to Argentine peso(ARS)). Additionally, we collected the Baltic Exchange Dry Index, measuring the cost of shipping dry goods like maize worldwide. To capture US market factors, we sourced data from the Federal Reserve Economic Data (FRED), including unemployment rates, the consumer price index, the industrial production index, interest rates, and gasoline prices. For US corn stocks data, we utilized quarterly information from the US Feed Grains Yearbook, converting it into monthly data for analysis³ Market factors for Ukraine, such as

²Cubic spline interpolation was employed to proxy missing price data within continuous periods. 20 observations are missing from April 2002 to February 2022 for the Ukrainian Maize export price.

³To align the data frequencies for our econometric analysis, cubic spline interpolation was applied

unemployment rates, the consumer price index, and the industrial production index, along with those for Argentina, such as unemployment rates, consumer price index, and inflation rate, were sourced from the National Summary Data Pages (NSDPs)⁴

The basic unit of analysis used throughout is the natural logarithm of the price ratio, denoted as $p_t^j - p_t^k (= \ln(P_t^j/P_t^k))$, where i and j indicate locations (i.e., $j, k = 1, 2, 3$ denote the US, Ukraine, and Argentina respectively), and t is a time index such that $t = 1, \dots, T$. The international price data are shown in logarithmic form in Figure 1b.

Figure 2 presents a graphical representation of logarithmic pairs of prices plotted against each other, offering insights into the relationship between price levels and price differentials. Deviations from the 45-degree line in each plot reveal distinct basis patterns, where one price tends to be higher or lower than the other. These patterns likely reveal the influence of transaction costs associated with regionally distinct market trades. While these countries are exporters only, in the market integration framework, maize flows between the three markets can occur in any direction, depending on potentially profitable arbitrage opportunities. Our observations from the figures indicate that situations where the price of Ukrainian maize surpasses the prices of US maize and Argentina maize occur more frequently.

To examine the characteristics of time series prices and identify the most appropriate model for evaluating spatial price linkages, we conducted augmented Dickey-Fuller tests for each pair of price differentials. The results of the Augmented Dickey-Fuller (ADF) tests for the stationarity of the price differentials are presented in Table 6 in

to convert the quarterly US corn beginning stock data into the same frequency as all other monthly variables. US corn beginning stock data is from 2001 Q2 (Dec-Feb) to 2021 Q3 (Mar-May), totaling 82 observations and converted to 246 monthly data. The seasonality of corn stocks was respected in our interpolation methodology.

⁴To align the data frequencies for our econometric analysis, cubic spline interpolation was applied to convert the quarterly Argentina unemployment rate and Ukraine employment rate into the same frequency as all other monthly variables. The data for Ukraine unemployment ranges from 2022 Q1 to 2021 Q4 (standard calendar quarters), totaling 80 observations and converted to 240-month observations. The data for Argentinian unemployment spans from 2002 Q4 to 2022 Q4, comprising 81 observations and converted to 243-month observations. Given that the variables, including the consumer price index of Ukraine, industrial production index of Ukraine, and industrial production index of Argentina, are segmented into multiple partitions over the selected period, and each partition is calculated using different units in the data sources, we employ cubic spline interpolation to estimate the data for the months where unit changes occur. Note that these variables are typically not volatile on a month-to-month basis, making spline interpolation a reasonable approach to converting the data to a monthly basis.

the appendix, which indicates that the null hypothesis of nonstationarity of the price differentials is strongly rejected in every case. This is as expected since a nonstationary differential would imply that prices can drift arbitrarily far apart.

Transmission elasticities ($\frac{\partial P_t^j}{\partial P_t^k}$) close to one provide support for market integration, with 1.0 corresponding to perfect market integration. Additionally, we performed ADF tests on the first differences of the logarithms of variables (all variables are logarithmic except for the unemployment rates of three countries and the inflation rate of Argentina), exchange rates, and other exogenous shocks. The results, presented in Table 7 in the appendix, indicate that all these variables significantly differ from nonstationary series. Our Augmented Dickey-Fuller (ADF) test on the first differences of all variables strongly rejects the null hypothesis of nonstationarity. Therefore, we can confidently implement Equation (2.5) for estimating the model with the available data.

Before consideration of two-regime switching models, we consider a suite of tests intended to detect departures from linearity in conventional time-series models. A range of (non-) linearity tests were conducted for the price data. We applied a standard Self-Exciting Threshold AutoRegressive (SETAR) model, as formulated by Goodwin and Piggott (2001), to prices in spatially distinct markets for each of the market pairs. The specification is given by:

$$\Delta(p_t^j - p_t^k) = \gamma_1(p_{t-1}^j - p_{t-1}^k) + \mathbf{1}\{|p_{t-1}^j - p_{t-1}^k| \geq c\} [\delta_1(p_{t-1}^j - p_{t-1}^k)] + \varepsilon_t \quad (3.1)$$

where c is a threshold parameter, and $\gamma_1 + \delta_1$ is the parameter for trade regime. Each of the linearity tests was applied to the collection of prices. Tests on pairs of prices were conducted on the differential between logarithmic prices. Nonlinearity testing results are contained in Table 1.⁵ The tests for all international market pairs reject linearity in at least one of the alternative linearity tests at a 10% significance level. Thus, these tests robustly reject linearity among the price linkages, prompting the exploration of alternative, flexible specifications capable of accommodating nonlinearities. Threshold models are a likely candidate for a nonlinear representation of the price relationships.

⁵Hansen's modification of standard Chow-type tests of the bootstrapping results presented in this paper utilized 1000 replications.

The question remains as to the most appropriate specification of the alternative models of price parity. We have suggested that, despite the fact that prices are all quoted in US dollar terms, exchange rates may nevertheless play a role in international price linkages. Specifically, if import prices pertain to an intermediary step in trade between internal markets, where different currencies may exist, and international markets, exchange rates may still be relevant to the price linkages. If exchange rates are found to exert a statistically significant effect on price linkages, exchange rate over- or under-shooting may exist. Further, it is unclear as to whether additional variables may be relevant to price linkages. Markets are separated by unobservable transactions costs, which may in turn be influenced by other economic variables. Hence, we utilize LASSO methods to select an optimal specification.

We estimate four versions of price transmission models which are successively more detailed. The first is a simple linear autoregressive model of prices alone. The second model introduces exchange rates into the models. A third model allows for nonlinear relationships through application of a threshold autoregressive model. Finally, we estimate a fourth model that includes other covariates, as selected by the debiased LASSO methods described above. These additional covariates are intended to capture other residual reasons why simple price linkages depart from equilibrium parity conditions.

We initially estimate price relationships for the three pairs of market prices using a standard autoregressive model of the form:

$$\Delta(p_t^j - p_t^k) = \gamma_1(p_{t-1}^j - p_{t-1}^k), \quad (3.2)$$

where $(j, k) = \{(1, 2), (1, 3), (2, 3)\}$, with the indices representing the US, Ukraine, and Argentina as 1, 2, 3 respectively. γ_0 and γ_1 are parameters reflecting the degree of market integration. In particular, we expect a small but negative value of γ_1 , so the price differential $p_t^j - p_t^k$ converges to 0 at the rate of $1 + \gamma_1 < 1$. A value of γ_1 closer to zero implies a slower adjustment to shocks. This model has been used extensively to evaluate price transmission and parity conditions.

Moreover, we then extend our analysis to include the exchange rate, considering

the following specification:

$$\Delta(p_t^1 - p_t^2) = \gamma_1(p_{t-1}^1 - p_{t-1}^2) + \gamma_2\Delta\pi_t^{12}, \quad (3.3)$$

where π_t^{12} is the exchange rate between countries one and two. If $\gamma_{2,0}$ is significantly different from zero, imperfect exchange rate pass-through is implied. It may seem odd to evaluate exchange rate effects when prices are quoted in a single currency, but price distortions caused by exchange rate shocks are possible, even in such cases.

Besides model (3.2) and (3.3), estimations are conducted based on model (3.1). In addition, we use exchange rates as covariates and estimate threshold models of the form:

$$\begin{aligned} \Delta(p_t^j - p_t^k) = & \gamma_1(p_{t-1}^j - p_{t-1}^k) + \gamma_2\Delta\pi_t^{jk} \\ & + \mathbf{1}\{|p_{t-1}^j - p_{t-1}^k| \geq c\} \left[\delta_1(p_{t-1}^j - p_{t-1}^k) + \delta_2\Delta\pi_t^{jk} \right] + \varepsilon_t. \end{aligned} \quad (3.4)$$

The 1nd to 4th columns in Table 2 present estimates of two standard autoregressive price parity models, as denoted by equations (3.2) and (3.3), and estimates of two threshold autoregressive price parity models, referenced by equations (3.1) and (3.4). The 2nd to 5th columns in Table 3 and Table 4 denote the same models for the other two pairs of market prices. When threshold behavior is neglected in the linear model, in every case, models incorporating exchange rates suggest adjustments in response to deviations from equilibrium that are at least as fast as the models when exchange rates are ignored. However, when considering the threshold models, the estimations for the “no-trade” regime are positive but close to zero in every case except for US/Ukraine with the exchange rate. This is consistent with expectations since, in our conceptual framework, markets are not linked in the no-trade regime.

The theoretical expectation is for these estimates to be negatively approaching zero. Nevertheless, except for the model of Argentina/Ukraine with the exchange rate, we obtain negative estimates for δ_1 , indicating much faster adjustments in response to deviations from equilibrium in the “trade” regime than the “no-trade” regime. This is implied if the estimates of structural effects γ_1 are negative. It’s interesting to note that when considering the exchange rate effect in the models, the exchange rates exhibit undershooting in all cases except for the trade regimes of US/Argentina and Argentina/Ukraine.

As mentioned earlier, our estimation procedures for threshold regression offer the advantage of variable selection and threshold detection, eliminating the need for conventional nonlinear tests commonly used in threshold models. In our study, the covariates included in all market pairs are the exchange rate, the Baltic Exchange Dry Index, unemployment rates, and industrial production indexes for each market. Additionally, for pairs including the US, we utilize US interest rates, US Corn Stock, and US gas prices as control variables. Due to data availability, consumer price indexes are collected for the US and Ukraine, while inflation rates are used for Argentina.

In this context, the symmetric lagged price differential $|p_{t-1}^j - p_{t-1}^k|$ transforms into Q_t , representing the quantile of $|p_{t-1}^j - p_{t-1}^k|$ in selected samples. The estimation of thresholds is conducted using a grid search. An assumption is made that all $|p_{t-1}^j - p_{t-1}^k|$ values are distinct. This is a convenient condition, ensuring that the transformation into quantiles is a one-to-one function without any loss of generality. This assumption holds under the assumption of continuous distribution for $|p_{t-1}^j - p_{t-1}^k|$.

The threshold estimates offer insights into transaction costs. Simultaneously, the quantile estimates (refer to the same tables) illuminate whether, during the selected periods, monthly observations more frequently align with trade regimes characterized by lower quantile estimates. In the scenarios of US/Ukraine and Argentina/Ukraine, price differentials within the bands occur more frequently, as quantile estimates of the threshold parameters exceed 0.5. However, for US/Argentina, arbitrage activities are triggered more frequently, leading to the “trade” regime. When examining the magnitude of the price differential estimates, the width of the band representing “no trade”, as implied by the thresholds, is widest for the Argentina/Ukraine markets and narrowest for the US/Argentina markets.

To make it comparable with our baseline model (the linear model that includes only the error correction term) in the first column, we set the first element in the scaling diagonal matrix $\mathbf{D}(c)$ in (2.6) to 0. This ensures that the LASSO estimation will always select the parameter γ_1 . The last (5th) column in each of the tables (Table 2, Table 3, and Table 4) presents estimates of the threshold model using debiased LASSO, as based on equation (2.5). In all three cases, the estimates of adjustment in response to deviations from equilibrium (γ_1) are negative and close to zero, indicating a significant convergence to market equilibrium. The nonlinear impact of the degree

of error correction manifests exclusively in the Argentina/Ukraine scenario, although the estimates are statistically insignificant. In contrast, in the other two cases, both nonlinear error correction and exchange rate terms are not selected.

Tables 5 display the signs of statistically significant estimates of market factors at the 10% significance level, excluding the error correction rate estimates (γ_1 and δ_1) and exchange rate effect estimates ($\gamma_{2,0}$ and $\delta_{2,0}$), which correspond to models represented by the last column in each table (Table 2, Table 3, and Table 4). It is important to note that debiased LASSO estimates are considered statistically insignificant when the slope estimates by equation (2.5) are zero. Thus, only non-zero estimates by equation (2.5) are presented. In the case of US/Ukraine, we observed that the Baltic Freight Index is not selected in the “no-trade” regime but is selected in the “trade” regime, suggesting that the Baltic Freight Index exerts a nonlinear effect, particularly in periods characterized by trade activity. In the case of Argentina and Ukraine, the 4th lag of Ukraine’s CPI and the 4th lag of Ukraine’s Industrial Production Index are selected only in the “trade” regime, indicating that these two variables produce a nonlinear effect, especially when trade opportunities exist. Furthermore, the Baltic Freight Index is selected in both the “no-trade” and “trade” regimes, but its magnitude in the “trade” regime is larger than in the “no-trade” regime, indicating a more pronounced effect when trade opportunities are present. However, there is insufficient evidence to suggest similar patterns for the US/Argentina pair.

Combining the findings from Tables 2, 3, 4, and 5, we can conclude that threshold effects are present in the US/Ukraine and Argentina/Ukraine cases. These results suggest that the market pairs of US/Ukraine and Argentina/Ukraine exhibit significant nonlinear adjustments influenced by specific market factors. However, no nonlinear effects are observed in the case of US/Argentina. This observation is consistent with the nonlinear tests presented in Table 1, nonlinearity is rarely and randomly significant in certain tests.

The standard threshold model assumes a fixed threshold, a potentially limiting assumption. It is reasonable to consider that relationships may evolve, signaling structural changes in the underlying economic dynamics. To explore this possibility, we introduce partitions that reflect changes in market environments. The data is segmented into two periods corresponding to two significant economic shocks: the 2014 Crimean crisis (e.g., [Korovkin and Makarin \(2023\)](#)) and the global financial/economic

crisis of 2008-09 (e.g., [Liefert et al. \(2021\)](#)). Specifically, the breakpoints for these events are defined as February 2014 and October 2008, respectively, in our monthly dataset. Introducing a break in the dataset corresponding to October 2008 for the global financial/economic crisis and February 2014 for the 2014 Crimean crisis reveals that, in most cases, the selected optimal lagged forcing variables differ across the entire period, pre-breakpoint, and post-breakpoint. Profitable arbitrage opportunities are more frequent only in the post-February 2014 period for the US/Ukraine markets, while in every other case, such opportunities are fewer. When examining the magnitude of the price differential estimates, post-break threshold bands are narrower in all comparisons except for US/Argentina pre/post-October 2008. Subsequently, we remove the shrinkage bias introduced by the penalization in Equation (2.6) using Equation (2.9) for post-selection statistical inference. Our estimation setup considers a richer examination of price linkage among global maize markets. The fundamental framework of the threshold model illustrates that if any of the estimates of the slope coefficients (exchange rate pass-through or exogenous shock) are regime-specific, the effects of certain lagged exchange rate or exogenous shock on price differentials (which could be lagged variables) between two distinct markets differs depending on the magnitude of a certain forcing variable representing unobserved transaction costs. Estimates of non-zero differences between the two regimes imply nonlinear relationships. The slope coefficient directly corresponds to elasticity, measuring the responsiveness of the dependent variable (the price linkages in time t) to changes in the explanatory factors (lagged exchange rate between the two markets or any market factor). A straightforward way to illustrate the effects of exchange rates, market factors, or exogenous shocks on potential deviations from price parity is by analyzing the coefficient estimates obtained from our estimations. All lagged variables are allowed to have a dynamic linear effect or a dynamic nonlinear effect depending on the existence of a regime switch (threshold).

Tables 8, 9, and 10 offer a comprehensive summary of the estimates for the degree of error correction and exchange rate effects based on (2.5) using the LASSO method⁶ These tables offer valuable insights into the adjustments and effects in each market

⁶If γ_1 , $\gamma_{2,0}$, δ_1 , or $\delta_{2,0}$ is not selected by the LASSO step, their estimates and standard errors are left blank in the table. The detailed debiased LASSO estimates for all variables are not included here due to space constraints but are presented in the Appendix.

pair, illuminating the interdependence between different markets. In almost every case, spanning entire periods and structural breaks, the estimates of the degree of “error correction” adjustments approach negative zero, except for the Post-October 2008 period in the US/Ukraine markets. The threshold models suggest adjustments in response to deviations from equilibrium in the “trade” regime that are at least as fast as those in the “no-trade” regime, except for the Post-October 2008 and Post-February 2014 periods in the US/Ukraine markets, and the Post-October 2008 period in the US/Argentina market, as well as the Pre-February 2014 period in the Argentina/Ukraine market. The threshold models indicate that exchange rates exhibit perfect pass-through to markets during the Post-October 2008, Pre-February 2014, and throughout the entire period in the US/Argentina pair across all 11 models.

In every case, some other market factors and their past period values are selected and tested as statistically significant.

4 Summary and Concluding Remarks

We develop a model of price parity in spatially distinct international export markets for maize to investigate the degree of “error correction”, the exchange rate pass-through, and other market factor effects. The models are developed within the framework of high-dimensional threshold models. We consider such nonlinear models, that has developed an increasingly rich set of factors in models of spatial market integration, as extensions to existing literature. The debiased LASSO estimation procedures are used to specify the models.

In summary, our findings consistently indicate faster adjustments in response to deviations from equilibrium in conditions of profitable trade and arbitrage compared to the case of no trade. The markets exhibit strong linkages in most cases, with confirmed nonlinear adjustments. Aligned with existing research, there is insufficient evidence to reject perfect pass-through of exchange rate fluctuations. However, the process of price convergence to market equilibrium is influenced by changes in some market factors, such as transportation cost, market unemployment rate, inflation, and the index of industrial production. These differences signify more substantial disequilibrium conditions, thereby presenting larger arbitrage opportunities.

Nonlinearity test				
	US/ Ukraine		US/Argentina	
	Test Statistics	p-value	Test Statistics	p-value
Teraesvirta's neural network test χ^2	2.596	0.273	2.541	0.281
White neural network test χ^2	2.898	0.235	6.587	0.037
Keenan's one-degree test for nonlinearity F-test	1.639	0.202	0.115	0.735
Tsay's Test for nonlinearity F-test	1.360	0.099	1.125	0.340
Likelihood ratio test for threshold nonlinearity χ^2	22.777	0.077	10.476	0.132
(SETAR) models: Linear AR versus 1 threshold TAR F-test	5.882	0.515	8.795	0.202
	Argentina/Ukraine			
	Test Statistics	p-value		
Teraesvirta's neural network test χ^2	3.005	0.223		
White neural network test χ^2	3.529	0.171		
Keenan's one-degree test for nonlinearity F-test	2.181	0.141		
Tsay's Test for nonlinearity F-test	2.166	0.048		
Likelihood ratio test for threshold nonlinearity χ^2	10.573	0.207		
(SETAR) models: Linear AR versus 1 threshold TAR F-test	6.216	0.479		

Table 1: Nonlinearity Specification Testing Results

Parameter	Linear without exchange rate	Linear with exchange rate	Threshold without exchange rate	Threshold with exchange rate	Threshold using debiased LASSO
degree of “error correction” γ_1	-0.114 *** (0.030)	-0.114 *** (0.029)	0.061 (0.069)	0.071 (0.069)	-0.074** (0.031)
exchange rate effect $\gamma_{2,0}$		0.193* (0.100)		0.188 * (0.107)	0.114 (0.084)
degree of “error correction” δ_1			-0.214** (0.076)	-0.228** (0.076)	
exchange rate effect $\delta_{2,0}$				0.188 (0.273)	
threshold estimate			0.130	0.130	0.124
threshold quantile			0.73	0.73	0.71
Observations	233	233	233	233	233
R ²	0.060	0.075	0.091	0.112	...
Adjusted R ²	0.056	0.067	0.084	0.097	...
F-statistic	14.92	9.42	11.63	7.24	...
AIC	-5.480	-5.487	-5.505	-5.511	-5.552

Note:

*p<0.1; **p<0.05; ***p<0.01

Table 2: Model Estimates: US/ Ukraine

Parameter	Linear without exchange rate	Linear with exchange rate	Threshold without exchange rate	Threshold with exchange rate	Threshold using debiased LASSO
degree of “error correction” γ_1	-0.134 *** (0.033)	-0.134 *** (0.033)	0.264 (0.219)	0.263 (0.220)	-0.177*** (0.030)
exchange rate effect $\gamma_{2,0}$		0.016 (0.062)		0.072 (0.116)	
degree of “error correction” δ_1			-0.406* (0.221)	-0.406* (0.222)	
exchange rate effect $\delta_{2,0}$				-0.079 (0.138)	
threshold estimate			0.033	0.033	0.069
threshold quantile			0.36	0.36	0.65
Observations	228	228	228	228	228
R ²	0.067	0.067	0.081	0.082	...
Adjusted R ²	0.063	0.059	0.073	0.066	...
F-statistic	16.31	8.15	9.92	5.03	...
AIC	-6.600	-6.592	-6.606	-6.600	-6.651

Note:

*p<0.1; **p<0.05; ***p<0.01

Table 3: Model Estimates: US/ Argentina

Parameter	Linear without exchange rate	Linear with exchange rate	Threshold without exchange rate	Threshold with exchange rate	Threshold using debiased LASSO
degree of “error correction” γ_1	-0.116 *** (0.031)	-0.126 *** (0.031)	0.126 (0.175)	0.095 (0.076)	-0.329** (0.156)
exchange rate effect $\gamma_{2,0}$		0.193 ** (0.082)		0.204** (0.092)	0.181** (0.076)
degree of “error correction” δ_1			-0.250 (0.128)	0.106 (0.142)	-0.117 (0.166)
exchange rate effect $\delta_{2,0}$				-0.285 (0.215)	
threshold estimate			0.063	0.119	0.133
threshold quantile			0.51	0.72	0.77
Observations	218	218	218	218	218
R ²	0.060	0.094	0.068	0.047	...
Adjusted R ²	0.055	0.085	0.060	0.029	...
F-statistic	13.76	11.1	7.90	2.65	...
AIC	-5.462	-5.479	-5.462	-5.422	-5.716

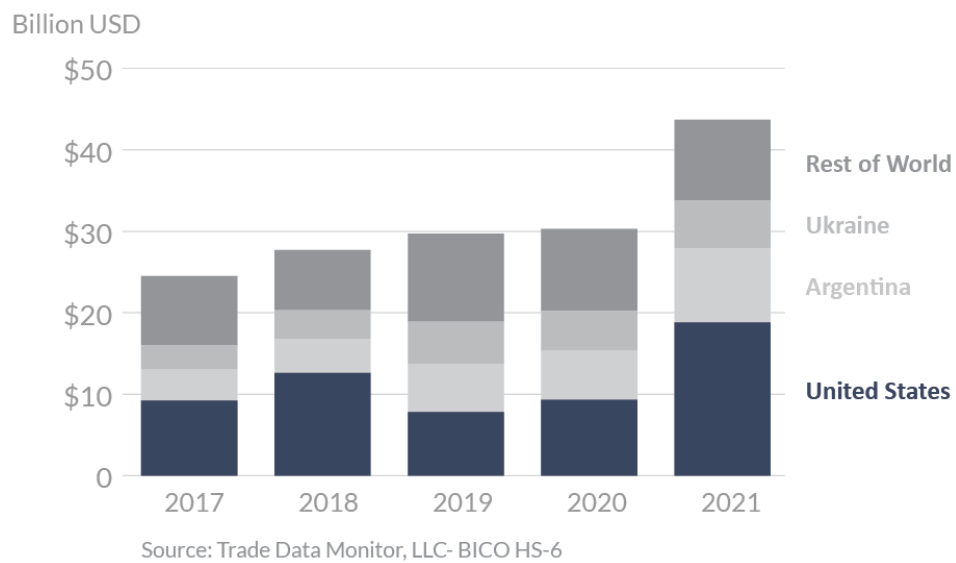
Note:

*p<0.1; **p<0.05; ***p<0.01

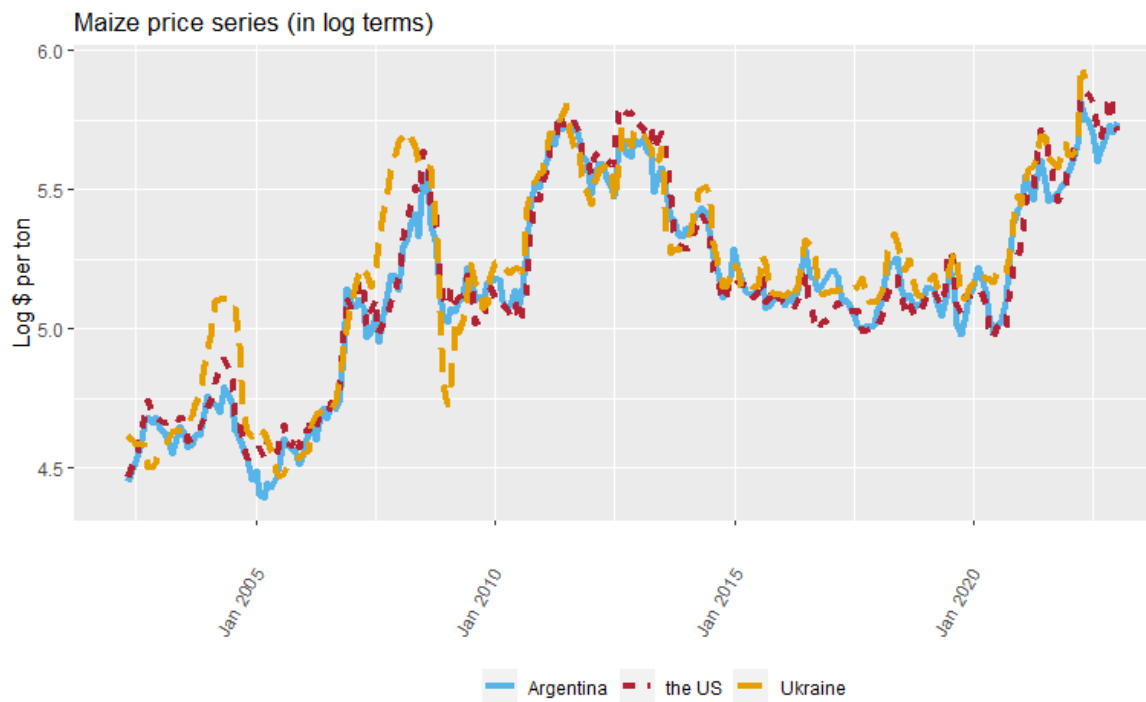
Table 4: Model Estimates: Ukraine/ Argentina

Model	US/Ukraine	US/Argentina	Argentina/Ukraine
lag 6 USD UAH Exchange	–		
lag 5 Baltic Freigh	–		–
lag 1 US CPI	–		
trade Baltic Freigh	–		–
lag 4 US Corn Stock		+	
lag 5 Argentina Unemployment Rate		–	
Baltic Freigh			–
Argentina Monthly Inflation			+
Ukraine INDPRO			+
lag 4 Peso UAH Exchange			+
lag 1 Baltic Freigh			–
lag 2 Baltic Freigh			–
lag 3 Baltic Freigh			–
lag 4 Argentina Monthly Inflation			+
lag 5 Argentina Monthly Inflation			–
lag 6 Argentina Monthly Inflation			–
lag 4 Ukraine CPI			–
lag 1 Ukraine INDPRO			+
lag 6 Ukraine INDPRO			–
trade lag 4 Ukraine CPI			+
trade lag 4 Ukraine INDPRO			–

Table 5: Signs of Market Control Variables Estimates at 10% significant Level



(a)



(b)

Figure 1: (a) Global Corn Exports by Country and Marketing Year, Source: [U.S. Department of Agriculture, Foreign Agricultural Service \(2022\)](#). (b) Maize Retail Price Series (in log terms) by Country

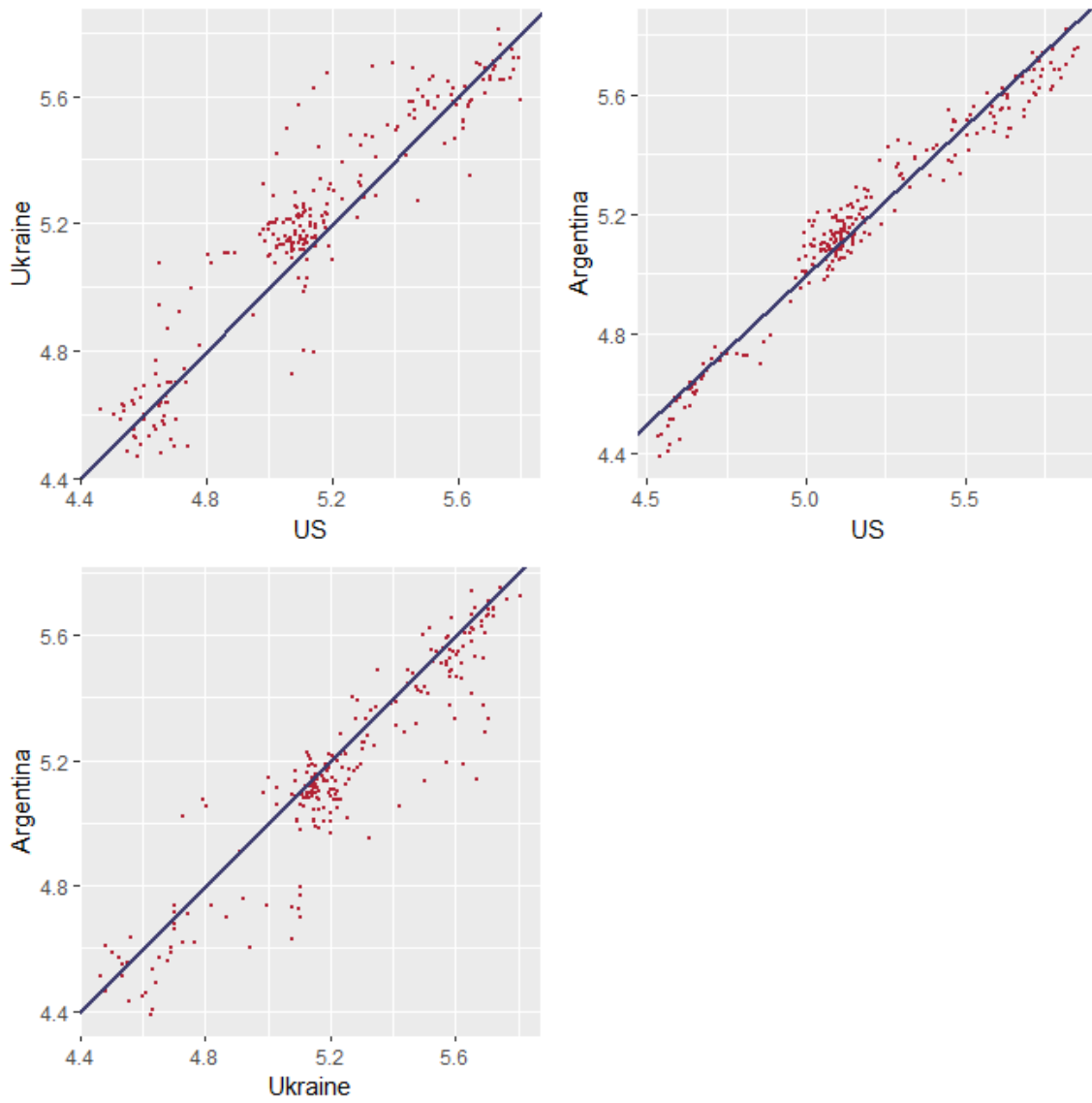


Figure 2: Maize Market Prices Pairs (in logarithms)

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5 Appendix

Augmented Dickey-Fuller test Results	
Variable	ADF
Unit Root	
ln_US_Ukraine_diff	-4.446
ln_US_Argentina_diff	-3.860
ln_Ukraine_Argentina_diff	-4.877

Alternative hypothesis: stationary	Lag order = 6
significance level	Critical value
1%	-3.96
5%	-3.41
10%	-3.12

*The critical values are interpolated from Table 4.2 of [Banerjee et al. \(1993\)](#).

Table 6: Augmented Dickey-Fuller Test Results of Price Differentials

Augmented Dickey-Fuller Test	
Variable (1st diff)	Dickey-Fuller
ln_US_Ukraine_diff	-5.403
ln_US_Argentina_diff	-8.379
ln_Ukraine_Argentina_diff	-6.283
ln_USD_UAH_Exchange	-5.207
ln_USD_Peso_Exchange	-5.222
ln_Peso_UAH_Exchange	-4.650
ln_Baltic_Freight	-8.028
US Unemployment Rate	-7.146
Ukraine Unemployment Rate	-6.261
Argentina Unemployment Rate	-5.148
ln_US Industrial Production Index	-5.794
ln_Ukraine Industrial Production Index	-6.467
ln_Argentina Industrial Production Index	-7.630
ln_US Consumer price index	-6.301
ln_Ukraine Consumer price index	-4.290
Argentina Monthly Inflation	-6.374
US Interest Rate	-3.055
ln_US Monthly Gas Price	-7.502
ln_US Corn Stock	-11.561
Alternative hypothesis: stationary	Lag order = 6
significance level	Critical value
1%	-3.96
5%	-3.41
10%	-3.12

*The critical values are interpolated from Table 4.2 of [Banerjee et al. \(1993\)](#).

Table 7: Augmented Dickey-Fuller Test Results of First Difference of Time Series

Parameter	Entire Period	Pre-October 2008	Post-October 2008	Pre-February 2014	Post-February 2014
degree of “error correction” γ_1	-0.074** (0.031)	-0.090 *** (0.001)	-0.295*** (0.000)	-0.258*** (0.020)	-0.165*** (0.001)
exchange rate effect $\gamma_{2,0}$	0.114 (0.084)	1.604 (0.553)	0.794 ** (0.326)	0.462 (0.293)	0.243*** (0.011)
degree of “error correction” δ_1			0.204 *** (0.003)		
exchange rate effect $\delta_{2,0}$			0.492 *** (0.017)		
other variables omitted
threshold estimate	0.124	0.099	0.068	0.097	0.062
threshold quantile	0.71	0.55	0.45	0.58	0.42
Observations	233	73	160	137	96

Note:

*p<0.1; **p<0.05; ***p<0.01

Table 8: Estimates of Threshold Model Using Debiased LASSO: US/ Ukraine

Parameter	Entire Period	Pre-October 2008	Post-October 2008
degree of “error correction” γ_1	-0.177*** (0.030)	-0.340*** (0.001)	0.164 (0.251)
exchange rate effect $\gamma_{2,0}$		-1.108*** (0.423)	0.050 (0.061)
degree of “error correction” δ_1			-0.085 (0.219)
exchange rate effect $\delta_{2,0}$		0.593*** (0.176)	0.479** (0.212)

threshold estimate	0.069	0.062	0.046
threshold quantile	0.65	0.56	0.50
Observations	228	58	170

Note:

*p<0.1; **p<0.05; ***p<0.01

Table 9: Estimates of Threshold Model Using Debiased LASSO: US/Argentina

Parameter	Entire Period	Pre-February 2014	Post-February 2014
degree of “error correction” γ_1	-0.329 ** (0.156)	-0.144*** (0.006)	-0.213*** (0.011)
exchange rate effect $\gamma_{2,0}$	0.181** (0.076)	1.639 *** (0.087)	0.159*** (0.014)
degree of “error correction” δ_1	-0.117 (0.166)	0.248*** (0.096)	
exchange rate effect $\delta_{2,0}$		-3.026*** (0.873)	0.183*** (0.026)

threshold estimate	0.133	0.098	0.093
threshold quantile	0.77	0.58	0.70
optimal threshold time delay	3	4	4
Observations	218	122	96

Note:

*p<0.1; **p<0.05; ***p<0.01

Table 10: Estimates of Threshold Model Using Debiased LASSO: Argentina/Ukraine