LONG RUN DYNAMICS OF WORLD FOOD, CRUDE OIL PRICES AND MACROECONOMIC VARIABLES: A COINTEGRATION VAR ANALYSIS

José M. Fernández

Department of Economics
University of Bristol
8 Woodland Road
Bristol BS8 1TN
United Kingdom

Long Run Dynamics of World Food, Crude Oil Prices and Macroeconomic Variables: A Cointegration VAR Analysis

José M. Fernández

University of Bristol, Department of Economics, 8 Woodland Road, Bristol BS8 1TN, UK.

Abstract

This study examines the long-run relationship between the real world price of maize, soybeans and sugar with the real world price of crude oil and a series of macroeconomic variables using a cointegration analysis from January 1982 until December 2012. The main empirical results support a strong causal relationship between maize and soybeans with crude oil, the real interest rate and the real U.S. exchange rate. Concretely, we show that real world crude oil prices are cointegrated with real world prices of maize and soybeans for the entire sample period and that real oil prices have a one-to-one relationship with these commodities. In other words, a one-percent increase in the price of real crude oil is associated with a one-percent increase in the price of maize and soybeans. Moreover, we find that permanent shocks to crude oil prices are transmitted to both maize and soybeans by a factor of 0.67 in both cases. In addition, our results show that despite the instability associated with the period between 2007/08, the long-run relationship between crude oil and these agricultural commodities has remained stable during the entire sample period. Finally, our results also support that although the real interest rate and the U.S. exchange rate are cointegrated with these commodities, it is only permanent shocks to real crude oil prices that have a permanent effect on these commodity price behavior.

Keywords:
Cointegration, crude oil, agricultural food commodities. JEL Classification: O13 C01 C32

Email address: Jose.Fernandez@bristol.ac.uk (José M. Fernández)

1. Introduction and Overview

The recent global price surge in energy and commodity markets have motivated the interest of economists to establish the underlining forces driving this relationship. Particularly, researchers have concentrated their efforts in determining the driving factors and channels through which energy and commodity price cycles are affected. In an effort to explain these dynamics, a significant number of studies such as Kilian (2008a); Spatafora and Tytell (2009); Carter et al. (2011); Wright (2011); Céspedes and Velasco (2012); Jacks (2013) have provided theoretical and empirical evidence for some of the reasons behind the most recent energy and commodity price fluctuations. Broadly, these can be categorized by demand and supply-side factors.

There are four demand-side factors that have been greatly studied in order to explain past and recent agricultural commodity and energy price increases. One of the most prevalent demand factors discussed in the literature is the increasing wealth of the BRIC (Brazil, Russia, India and China) and other developing nations within the past decade. This wealth effect is thought to have triggered greater global demand for energy and agricultural (food) products and together with relatively flat production was responsible for the 2007/08 price surge (Josling et al., 2010; Cairns and Meilke, 2012). Thus, as a number of authors claim, the last energy and commodity price shock as a result of primarily the unprecedented demand from developing markets as a result of this wealth effect (Abbott et al., 2009; Hamilton, 2009)[1]. By analyzing the dynamics of a number of demand-side factors before the 2007/08 commodity price crisis, Hamilton (2009) offers important contributions to understanding the causes leading to this episode. The author concludes that an important contributor to high crude oil prices (during this period) was the greater energy demand as a consequence of increasing wealth from developing nations (particularly that of China)[2]. Similarly, Kilian and Hicks (2013) provide empirical evidence supporting the view that strong (and unexpected) growth in emerging economies is able to explain the increase in the real world price of crude oil during commodity and energy shock of the mid 2000s. Consequently, the literature suggests that increasing income, and in particular from developing nations, is a crucial element in determining the price dynamics leading to the recent oil price crisis.

Greater demand from developing nations, as a consequence of their income effect, is argued to be responsible not only for the price increases in the global energy sector, but also in the agricultural commodity markets. For instance, Cooke and Robles (2009) indicate the purchasing power gains from developing nations, as an important element in increasing local demand for meat which in turn increases the demand for livestock, which competes with feedstock commodities such as maize and soybeans (See also, Zhang and Law (2010); Henderson (2011) for more details on equivalent arguments). Contrary to these arguments, Headey and Fan (2008) argue that demand from emerging economies particularly from China and India cannot

---

[1] In addition, there are also external factors such as an stagnate world oil production in previous years and historical low stock-to-use ratios.

[2] According to Leung (2010), between 2003 and 2007, China's oil demand alone, was responsible for approximately 37.1% of the increase in world oil consumption.

alone explain the increase in agricultural food commodity prices during this period. More precisely, Headey and Fan state that China during the years leading to the commodity price crisis (2000-2007) imported approximately 20% less wheat than the preceding period while India's imports of wheat and corn remained insignificant relatively to the world's imports. However, Headey and Fan (2008) conclude that at least in the crude oil and oilseeds market, increasing demand from emerging economies might explain some of the variation in global prices during this period. Consequently, if there is a causal relationship between crude oil and food commodity prices, there exists a possibility that some of the feedback in the oil market being transmitted to the agricultural food commodities. Therefore, global real economic activity appears as an important determinant of world agricultural commodity prices and should be considered in the analysis of price dynamics.

Another important demand-driven component affecting price dynamics in agricultural markets, is the increasing production of biofuels using agricultural commodities. The rise of biofuels production fostered by government subsidies and tax incentives, has expanded the supply for these products towards the biofuel production and away from food production. The biofuel revolution has significantly hoarded away from the food chain large quantities of grains and vegetable oils, as their primary inputs and compete with food crops for resources, such as land and water (Chen and Khanna, 2013)[3]. For instance, Mitchell (2008) cites the increasing production of biofuels in both the U.S. and EU markets as the crucial trigger for the 2007/08 food price crisis. Also, Tyner (2010) points out the strong correlation during 2008/09 between crude oil and ethanol with corn prices (0.95 and 0.84 respectively) as evidence of the significant shift in the fuel to food price dynamics. More recently a study conducted by Condon et al. (2013) estimates that in the U.S. "each additional billion gallons of ethanol causes a 5-10 percent increase in corn [nominal] prices." Nevertheless, while a number of authors have argued that the introduction of ethanol feedstocks has had a significant impact on food commodity prices, recent studies have highlighted that this impact is much less significant than economic growth (Zilberman et al., 2013).

A third determinant in the energy and commodity price dynamics is derived from the U.S. macroeconomic policies influenced by its monetary policy, exchange rate and inflation. Abbott et al. (2009); Headey and Fan (2008) singled out the weak U.S. exchange rate in the years leading to the peak of the 2008 commodity price as one of the main drivers of energy and agricultural commodity prices. Since all energy and agricultural commodities are traded in U.S. dollars as it depreciates (e.g. expansionary monetary policy) with respect to major trading partners, then commodity prices rise since now the international purchasing power has increased (the opposite effect occurs if the dollar appreciates) (Phillip and Friederich, 2013). On the other hand, Bernanke et al. (1997) and Barsky and Kilian (2001) have argued that loose monetary policy is responsible for the energy and commodity price shocks of the late 1970s and that anti-inflationary policies in the subsequent decade

---

[3]See also Colin Carter and Smith (2012) and Roberts and Schlenker (2013) for additional discussions in this subject.

drove prices down. The argument is very intuitive as there exists a negative trade-off between holding cash and storable commodities as interest rates increase. For example, low interest rates (i.e. expansionary monetary policy) today result in higher prices by increasing the demand for commodities as a consequence of the lower costs of borrowing, which act as a positive incentive to purchase and stockpile storable commodities with the intention of having positive returns in the near future (Frankel and Rose, 2010). On the other hand, real exchange rates have been shown to share a long-run equilibrium with a number of commodity prices in addition to being the long-run adjustment force for a number of these commodity prices (Cashin et al., 2004). Therefore, it is evident that from both empirical and theoretical perspectives, macroeconomic variables affect both the energy and agricultural commodity markets and should be included in the analysis.

The last, but not least, factor affecting these markets is increasing flows of capital or financialization of future contracts in the energy and agricultural commodity markets, have also thought to have contributed to the price increases in the last decade. The financialization of commodities has been argued to allow room for speculative trading behavior and consequently to be responsible for the large volatility swings during the last commodity price boom. Tang and Xiong (2010) have outlined that "index-related instruments" increased from $15 billion in 2003 to approximately $200 billion by the middle of 2008. In the same study, the authors conclude that as a result of the financialization process of commodities, demand and supply are not the sole determinants of individual commodity prices, but to a great extent their price is determined by the entire financial sector. These findings are also corroborated in a recent empirical work by Pen and Sevi (2013). There, the authors summarize their findings by stating that non-market fundamentals explain approximately 60% of the price variations. However, Kilian and Murphy (2013) find no evidence that speculation is responsible for the 2008 crude oil price peak; instead, substantial evidence is provided showing that market fundamentals were responsible for the price surge. And more recently, Hamilton and Wu (2014a) corroborates the previous study by not finding evidence that speculative positions of investors were able to drive future agricultural commodity prices. Similarly, Hamilton and Wu (2014b) show that before 2005, investors who consistently took long positions on crude oil futures contracts received on average positive returns on their investments; however, after this period, the authors find substantial evidence supporting the contrary position to the previous sample period. In summary, although early evidence in the literature supports non-market fundamentals theory as an explanation for increasing price shocks in the commodity market, recent work indicates even if there are instances where speculation is present, in the long-run the main drivers of commodity prices are economics fundamentals and more precisely demand-side factors.

Finally, there are also supply-side factors that have contributed to the commodity price surge in the past years. Although some of the supply shocks are associated with long-run effects on price variability (e.g. low growth in agricultural production as a result of lack of investment on R&D), a significant source of price shocks in the energy and agricultural commodity markets are associated with short-run dis-

ruptions. For the most part, supply shocks arise from social and political unrest, drastic changes in weather patterns as well as higher inputs costs of fertilizers and transportation[4]. This last factor, is an important aspect to consider, since price increases in crude oil affects agricultural commodities as an input, but also as the oil price increases so does the price of ethanol and other biofuels and therefore more grains and vegetable oils are redirected from food to energy consumption (Mutuc et al., 2010). Nevertheless, the literature provides significant empirical evidence showing that supply-side shocks are associated with having very low and temporary effects, relative to demand-side shocks, on both energy and commodity long-run price dynamics (Headey and Fan, 2008; Kilian, 2008b, 2009; Kilian and Murphy, 2012). Consequently, this study focuses its attention on the demand-side factors since according to the empirical literature in the long-run, supply-side factors have almost negligible effect on past energy and agricultural commodity price booms.

We can summarize the recent shocks to oil and agricultural commodity price by a series of demand-side factors: (1) Increasing wealth in developing economies; (2) biofuel production; (3) financialization of commodities, which allegedly result in market speculations; (4) macroeconomic cycles. Even though oil and agricultural commodity prices share mechanisms through which prices can be disturbed, their long-run causal effects are still not settled in the literature. Therefore, this study will focus on determining the causal relationship between crude oil and a set of individual agricultural commodities considering a number of macroeconomic variables. In essence, this study hopes to contribute to the literature in offering an understanding of the extent to which world real crude oil prices have a long-term effect on agricultural commodities as well as considering macroeconomic channels through which these might also be affected. Thus, offering an understanding of the long-term channels through which maize, soybeans and sugar global prices are affected.

From the analysis, is evident it is important to understand the channels through which agricultural commodity prices can be affected by crude oil and macroeconomic factors. In addition, and equally important it is necessary to establish the extent which fluctuations in the economy are propagated to the agricultural commodity markets. Therefore, it is essential to evaluate the strength of the relationship between agricultural commodity prices and the economic forces driving the price cycles.

2. Literature Review

In the wake of the energy and commodity price surge of 2007-08, economists became concerned with understanding world commodity price fluctuations. In particular, researchers paid close attention to two distinct areas of research. The first was concerned with the impact of rising biofuel production on food commodity prices (e.g. maize, soybeans and sugar). This area of research arose in order to understand soaring global food prices during this period from adopting biofuel production

---

[4]In addition to these, there are also factors affecting the world food commodity markets associated with population growth and dietary changes (Chakravorty et al., 2012).

policies. As a result, this body of literature concentrated on studying the impact of local subsidies fostering biofuels industries to offset increasing global energy prices in major grain and food markets such as Brazil, the U.S. and Europe[5]. The second body of empirical studies has been aimed at understand the causal relationship between global food commodity markets and world crude oil prices. The latter, is the main focus of this paper. We are mainly interested in understanding the dynamics affecting global food price cycles through macroeconomic and crude oil price fluctuations.

The literature on the causal effects from crude oil to commodity prices is vast. Within the past decade there has been a substantial number of studies concentrating on the impact of shocks in the crude oil market on food commodity prices. Yet, very few studies have emphasized the importance of understanding the fundamental economic forces underlining this relationship. The great majority of these studies have focused their attention on the co-movements across individual commodities linked with crude oil prices. These studies have used time series analysis and in particular Cointegration and Vector Error Correction Models (VECM) in order to estimate both short and long-run dynamics among these markets. Nevertheless, and despite the large body of research on this topic, the literature is far from in agreement and results vary widely from finding no evidence of long-run relationships at all, to strong and positive long-run relationships.

One of the first empirical works to address the long-run relationship between crude oil and commodity prices is by Chaudhuri (2001). The authors use a bivariate Johansen cointegration approach to model the effects of real oil price shocks on several primary commodities, using monthly data from 1973 to 1996. They conclude that all commodities analyzed (including maize and sugar), have a long-run relationship with real oil prices. Interestingly, the authors cite macroeconomic effects (e.g. low levels of interests rate and increasing economic activity from developing economies) as explanatory variables in the rise of commodity prices during this period, but do not include any of these variables in their analysis. One of the first studies specifically concerned with the relationship between crude oil and food commodity markets, is that of Campiche et al. (2007). Here, the authors main objective is to establish a direct link between world crude oil and soaring food prices. In this respect, Campiche et al. concentrate their efforts on the co-movement between major food commodities (corn, sorghum, sugar, soybeans, soybean oil and palm oil) and nominal crude oil prices by using a bivariate Cointegration Vector Autoregressive (CVAR) model and VECM with weekly data from 2003 to 2007. The authors' conclusions vary in terms of success and are unable to determine any long-run relationship (i.e. co-movement) among any individual commodity and crude oil for the 2003-05 period. Yet, they find a cointegrating relationship between corn and soybean prices with crude oil during the 2005-2007 period; thus providing some evidence of a strengthening relationship between energy and commodity markets during this period. On the other hand, Baffes (2007) examines the passthrough ef-

---

[5]Zilberman et al. (2013) provides a succinct review of this body of literature.

fect from real crude oil price shocks on several commodity indices using yearly data from 1960 to 2005. The authors also use a bivariate cointegration approach in order to estimate the long-run price elasticity of real crude oil prices and determine that the passthrough effect is 16% for all commodities and 18% for the food commodity during the entire sample. More importantly, the authors recognize the simplicity of their model and thus the inability to capture the complex relationship of commodity price dynamics. Similarly, in a recent study by Ciaian and Kancs (2011b) using a bivariate CVAR/VECM using weekly data from 1994 to 2008, the authors analyze the long-run relationship between nine agricultural commodities and crude oil prices. In order to account for structural breaks in the series, Ciaian and Kancs split the sample into three distinct periods of four years each from 1994-1998, 1999-2003 and 2004-2008[6]. The authors conclude by providing evidence in support of the increasing interdependencies between agricultural commodities and crude oil prices in particularly for the period between 2004-08. Although these studies have shed light into the some of the circumstances in which oil and commodity prices were linked during this period, none of these studies accounted for the fundamental factors affecting both oil and commodity prices as possible explanations.

On the other hand, some other studies have provided contradictory evidence (even when using similar methodologies, data spans and frequency) showing that the literature is far from in agreement on the empirical evidence of the causal links between crude oil and agricultural commodities. For example, in a study by Yu et al. (2006), the authors apply a CVAR/VECM analysis using weekly data from crude oil and four edible oils prices from January 1999 to March 2006 in order to estimate the long-run relationship between these commodities. Contrary to previous finding, Yu et al. (2006) conclude that crude oil prices do not have a significant impact on the long-run behaviour of vegetable oil prices. In another study, Zhang et al. (2009), also using a VECM, study the relationship between fuel (crude oil, gasoline and ethanol), corn and soybeans prices in the U.S. using weekly data from 1989 to 2007 and conclude that there exists no long-run relationship among these variables. Similarly, Zhang et al. (2010), using a longer data set and with a monthly frequency for a similar set of commodities, concludes that there is no long-run price relationship between crude oil and agricultural commodity prices and very weak evidence of short-run dynamics. Moreover, Saghaian (2010), also using cointegration analysis, rejects the hypothesis of cointegration between crude oil, corn, soybeans and wheat prices and is only able to show that crude oil prices Granger cause corn, soybeans and wheat prices. In summary, one of the main drawbacks of the current state of the literature, is the insufficient understanding of the transmission mechanisms in the relationship between crude oil and agricultural commodities. Consequently, there exists the need to move away from bivariate models and incorporate fundamental economic factors in the analysis in order to capture the possible dynamic forces, if any, in this relationship.

---

[6]The authors do not offer a clear methodology used to determined the break points nor the events that might have induced these. The authors state that: "The segmentation of the sample corresponds roughly to structural beaks."

The literature has been dominated by studies that simplify the complex relationship between crude oil and commodity prices and consequently are unable to offer sound empirical economic justifications for the transmission mechanism between these variables. Concretely, a significant number of the studies have limited their attempts to modelling the long relationship between these variables using bivariate models and using relatively short period of times that can not possibly measure long-run relationships. Moreover, there is not only theoretical reason to believe that there is a direct causal relationship between crude oil and commodity prices (other than as an input or transportation process). However, it is possible that the same underlined forces driving global crude oil prices (e.g. world demand factors) are also affecting commodity price cycles. Therefore, efforts to model long-run relationships between these variables using bivariate models may offer a misleading and incomplete understanding of the relationship dynamics. This might explain why some studies are able to get different conclusions even when using similar time periods, since the underlining forces driving this relationship might be strengthened and weakened independently of the variation between these individual variables.

In determining the transmission channels, many studies have singled out macroeconomic variation as significant factors driving oil as well as agricultural commodity prices. The link between the U.S. dollar exchange rate and commodity price cycles is very well discussed in the literature. Nevertheless, the underlying forces driving the dynamics of this relationship are yet to be understood. Commodity prices are quoted in U.S. dollars, thus it suggests that exchange rate fluctuations in this currency would be associated with commodity price fluctuations (Frankel, 2008). Yet again, there is no reason to believe in a direct causal relationship between U.S. dollar currency fluctuations and world commodity prices in the long-run. In fact, it is possible that the factors influencing the exchange rate (e.g. interest rates and the current account deficit) also affect the state of commodity price cycles. For instance, Gilbert (2010) argues that the co-movement experienced between exchange rates and commodity markets in general have the business cycles as a common component. Moreover, Abbott et al. (2011) point out that economic growth and the forces driving U.S. dollar exchange rate fluctuations are also important forces currently driving food prices. Similarly, Hamilton (2009, 2011) states that loose monetary policy is one of the drivers of the 2007/08 oil price shock. Therefore, in order to estimate any relationship between crude oil and commodity price it is crucial to also include macroeconomic variables since any causal effect from crude oil to commodity markets is likely to be overestimated or underestimated (depending on the relationship) in a simple bivariate CVAR model. Nevertheless, there is no study in the current literature that has incorporated the impact of all these macroeconomic dynamics in efforts to understand its effect in the global agricultural and food markets.

The effects of the U.S. dollar exchange rate have been addressed by a number studies in the recent past, but only a few have been able to draw meaningful conclusions in the global market. Distinctly, there is the work by Nazlioglu and Soytas (2012) by which the authors examine both short and long-run relationships between crude oil, the lira-to-U.S. dollar exchange rate and a number of individual agricultural commodity prices in Turkey. They apply a panel cointegration analysis as well

as Toda-Yamamoto causality to test monthly data from January 1994 to March 2010. Nazlioglu and Soytas conclude that there is no long-run transmission mechanism between fluctuations in the lira-to-U.S. dollar exchange rate and world oil prices to agricultural commodity prices in Turkey. On the other hand, Baek and Koo (2010) using an ARDL cointegration approach show that for the U.S. market, the exchange rate helps to explain variation in the short and long-run food markets. Other studies have concentrated their efforts on the effects of the U.S. dollar exchange rate on the world agricultural and food prices such as Harri et al. (2009); Kwon and Koo (2009); Gohin and Chantret (2010). In the study by Harri et al., the authors are interested in estimating the relationship between world crude oil future prices and the U.S. dollar exchange rate with individual corn, soybean, soybean oil, wheat and cotton world future prices using cointegration analysis with data spanning from January 2000 to September 2008. The authors, split the data sample into two periods in order to find a cointegrating relationship between the dollar exchange rate and oil future prices with corn from early 2006 to 2008. Kwon and Koo use a vector moving average (VMA) analysis and conclude that "unexpected movements of the exchange rate as well as interest rate are the main macroeconomic shocks causing fluctuations in the agricultural sector." On the other hand, Gohin and Chantret employ a Computable General Equilibrium (CGE) model of the macroeconomic linkages between a series of world food and energy prices. Gohin and Chantret's simulations conclude that macroeconomic variables provide a substantial explanatory power for the global energy and food price fluctuations.

Currently, the extent to which the global market for agricultural commodities have been affected by fundamental factors as well as energy price shocks remains a subject of great debate and often with conflicting results. Nevertheless, the most frequent methodologies to model this long-run relationship have been by using a Cointegration Vector Autoregressive (CVAR) approach. The advantage from using this methodology is that it allows researchers to determine both short and long-run parameters through a Vector Error Correction Model (VECM) as well as determining the pulling and pushing forces of the system of interest (Juselius, 2006). However, the majority of these studies have relied on oversimplified relationships in order to measure the long-run dynamics of agricultural prices. In this study, we will use a CVAR approach by incorporating a series of macroeconomic variables in addition to crude oil prices in an attempt to capture both short as well as long-run driving forcers of three agricultural commodity prices (maize, soybean and sugar) used in the production of biofuels. Also, as in Gilbert (2010), we argue that fundamentals in the market for agricultural food commodities are the principal dynamic drivers of global prices. This study differs from those in the current literature not only by using observations covering thirty years at a monthly frequency, but more importantly by exploiting the information dynamics including macroeconomic and energy prices together with agricultural commodities into one estimating system.

3. Methodology

The modeling approach used in this analysis is a cointegrated vector autoregressive (CVAR) proposed by Johansen and Juselius (1990) where first the cointegration space (i.e. long-run relationship) is estimated and subsequently we proceed by testing specific economic hypothesis within this space (Johansen, 1992). Lets consider a general $p - dimensional$ VAR model with order k lags in its vector error correction (VECM) form:

$$\Delta X_t = \Pi X_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \mu_0 + \mu_1 t + \Phi_1 D_t + \varepsilon_t,$$

$$\forall\, t = 1, 2, \ldots, T$$

(1)

where the difference between the mean and the actual realization is a white-noise process with mean zero and covariance Ω (i.e. $\varepsilon_t \sim NI_p(0, \Omega)$). Thus, Equation (1) is consistent with agents who are rational in the sense that they do not make systematic errors based on previous realizations (Juselius, 2006). The dimensions of parameters Π and Γ_i for $i = 1, 2, \ldots, k-1$ are $(p \times p)$ and $(p \times 1)$ for the parameters μ_0 and μ_1. The parameter Φ has dimension $(p^D \times p)$ and dimension $(p \times p^D)$, where D_t can include seasonal centered and intervention dummy variables in the form of transitory shock as well as mean-shift dummies[7].

Equation (1) provides a convenient formulation to analyze the dynamics of the system. In this case, the short-run effects are given by Γ_i and the long-run effects (levels) of the model are captured by the parameters in Π. Provided that Π has a reduce rank $(r < p)$, that is, assuming Equation (1) contains a mixture of stationary and non stationary components, then there exists $p \times r$ matrices α and β, each of them with rank r such that $\Pi = \alpha \beta'$ and $\beta' X_t$ is stationary. In this case the number of cointegrating relationships is determined by the rank 'r', while the adjustment parameters are found in the α matrix and $\beta' X_t$ represents the r number of cointegration relations. The cointegrating relationships determine the deviations front he long-run dynamics between the variables and the coefficients in α measure the rate of adjustments to any deviations from the long-run relationship.

Cointegration is a powerful tool in order to understand both the short and long-run dynamics of agricultural commodities, crude oil and macroeconomic variables. This is because as long as there exists a cointegrating relationship among these variables, it means that there is a stationary long-run equilibrium relationship between the individual non-stationary variables and in the cases where these diverge (pushing-forces) from this long-run equilibrium at least one of the variables in the system returns (pulling-forces) to the long-run equilibrium level (Juselius, 2006). The fact that deviations from the long-run equilibrium are stationary, ensures that deviations from the long-run equilibrium (i.e. cointegration relationship) of individ-

---

[7]Transitory shock dummy variables (D_{p_t}) is a vector of $(\ldots, 0, 0, 1, -1, 0, 0, \ldots)$; the vector of permanent blip dummies is defined as $(\ldots, 0, 0, 1, 0, 0, \ldots)$; and the vector of mean-shift dummies (D_{s_t}) as $(\ldots, 0, 0, 0, 1, 1, 1, \ldots)$.

ual variables are bounded despite these presenting path-dependent behavior. Therefore, by estimating a CVAR model we will be able to determine the pushing and pulling forces of this system of variables, thus helping us to predict both short and long-run behavior of food prices considering the crude oil and macroeconomic dynamics.

4. Data Description

The objective of this analysis is to determine the long-run relationship as well as short-run dynamics between a number of macroeconomic variables as well as energy prices on the world price of three major agricultural commodities: maize, soybean and sugar. Particularly, the study aims to determine the effects of four macroeconomic variables: the inflation rate, real exchange rate, short-term interest rate, and the Kilian index for global economic activity as defined in Kilian (2009) in addition to the world crude oil prices on these three agricultural commodities.

All data series have a monthly frequency and the observations span from January 1982 until December 2012 (See Table 1). In addition, all variables have been transformed using natural logarithms in an effort to obtain stable series in percentage terms and to be approximately linear (In and Inder, 1997). The agricultural commodities of interest are the price of maize (MZ_t), soybean (SB_t) and sugar (S_t), which were all obtained from the IFS database and are measured in U.S. dollars per metric tonne. All these agricultural price variables are world benchmark price series which are representative of the global market and are determined by the largest exporter of this specific commodity[8].

As a measure for the inflation rate, we have used the U.S. Producer Price Index (PPI) for all commodities (not seasonally adjusted) since the variables of interest are widely used as intermediate goods in industrial production. The real exchange rate (XR_t) was obtained and constructed by the Board of Governors of the Federal Reserve System, and is defined as the weighted average of the foreign exchange values of the U.S. dollar against the currencies of major U.S. trading partners converted to real terms. For the short-term interest rate (i_t) we have used the three-month Treasury bill secondary market rate as reported by the Federal Bank of St. Louis in the FRED database (See Table 1). Moreover, as a measurement of real global economic activity (Y_t) we use Lutz Kilian's index of global real economic activity in the industrial commodity market as defined in Kilian (2009). This index considers ocean freight rates as an observable real activity variable since in the short run the fleet of transport vessels is essentially fixed. The index is constructed by Lutz Kilian and it primarily represents the average freight rates for cargoes of grain, oilseeds, coal, iron ore, fertilizer, and scrap metal as reported by Drewry's Shipping Monthly[9]. Additionally, we also include the world price of crude oil (O_t) measured as the trade weighted average price of crude oil in U.S. dollars per barrel, obtained

---

[8] For more details on the definition of reference world prices and its benchmarks, please visit http://www.imf.org/external/np/res/commod/faq#q6

[9] For more details on the construction of this index, please see Kilian (2009)

from the IMF International Financial Statistics (IFS) (See Table 1).

The beginning of the data period for this study has been selected based on data availability for all series, in addition to avoid a period of price as well as economic and political instability from 1975 until the end of 1981. On the other hand, the end period was based on the latest data point available at the time of the analysis being written. The sample period covers the most recent macroeconomic, crude oil and commodity price shocks as well as price collapses within the past thirty years. This way, our study ensures significant variability in the observations to estimate the dynamic fluctuations between macroeconomic factors, world crude oil and maize, soybean and sugar prices. However, this very same feature presents a significant challenge to the modelling techniques since we will have to incorporate into the empirical estimation a significant number of instability periods typical of commodity price series (Figure 1 and 2).

In this analysis, all price series have been deflated by dividing each of them by the PPI series and subsequently taken the natural logarithms. Also, the monthly inflation rate has been calculated by taking the natural logarithms of PPI and subtracting the current value with that from the previous period (i.e. $\Delta p_t = ln(PPI_t) - ln(PPI_{t-1})$). It is worth noting that a brief graphical analysis from Figures 1 and 2 appear to show that these series do not have a long-run deterministic trend or cyclical component across this time period in agreement with the arguments presented in Sarris and Hallam (2006) as well as the apparent non-stationary nature of these commodity series.

Tables 2-4 (bottom), summarize the univariate analysis for the skewness, kurtosis, normality and ARCH effects of all variables (already deflated) in each system. From these tables, it is evident that all variables show signs of skewness (both positive and negative skewness) and kurtosis. In the case of maize and crude oil, high values of positive skewness imply that there are very high positive spikes (i.e. upward movements in price) which are rarely matched with downward movements in the price of these commodities. Also, from Tables 2-4 the calculated values for kurtosis indicate all of these series suffer from thicker tails than a normal distribution. As a result, not surprisingly, the tests for normality is rejected for all six variables. These are some characteristics of the data which will have to be considered in the specification of the model since it can lead to violations of the assumptions of the statistical model and consequently unreliable estimates.

5. Empirical Analysis

In order to derive the full-information maximum likelihood (FIML) estimator it is required to produce an explicit probability formulation of the initial estimated VAR model. Consequently, in estimating the model we assume multivariate normality (e.g. Homoscedasticity and no significant serial autocorrelation in the residual errors.). In case the model is unable to fulfill these assumptions we may be not able to provide any conclusive evidence since the parameter estimates are based on

an incorrectly derived estimator. Consequently it is essential that "to claim that conclusions are based on FIML inference is to claim that the empirical model is capable of accounting for all the systematic information in the data in a satisfactory way" (Juselius, 2006).

In our study, we are interested in estimating three models corresponding to each commodity of interest (i.e. Maize, Soybeans and Sugar). Our empirical analysis begins by defining the three models of interest and subsequently estimating all three unrestricted VAR (UVAR) models. Then, we conduct a series of analyses in order to find a well-specified empirical model. Once we have statistical well-specified models, we proceed to test for the rank and stability of the system. Finally, we proceed to identify the long-run relations by conducting a series of theoretical and empirical restrictions in the equations of interest.

The initial three unrestricted models consist of six variables each ($p = 6$) in the following form:

$$X'_{1,t} = [mz_t, o_t, y_t, \Delta p_t, xr_t, i_t]' \ - \ \text{Model 1}$$
$$X'_{2,t} = [sb_t, o_t, y_t, \Delta p_t, xr_t, i_t]' \ - \ \text{Model 2}$$
$$X'_{3,t} = [s_t, o_t, y_t, \Delta p_t, xr_t, i_t]' \ - \ \text{Model 3}$$

where mz_t, sb_t, s_t, o_t, y_t, xr_t and i_t are the logarithm prices of maize, soybeans, sugar, crude oil, the Kilian index of global economic activity, real exchange rate, nominal interest rate and inflation rate[10]. For each one of the three models, the initial number of lags has been selected to three (i.e. $k = 3$) by minimizing the standard Schwarz (SC) and Hanna-Quinn (HQ) information criteria as well as based on the model misspecification of 'no serial autocorrelation' on the first and k^{th} lag. Additionally, since none of the series present linear trends (at least not until the very end of the sample), we assume that the only deterministic component in all three models are the intercepts, which have been restricted to the cointegrating space. In general, the estimated unrestricted VECM (UVECM) of order $k - 1$ with the six variables ($p = 6$) for maize[11] can be written in matrix form as:

$$
\begin{bmatrix} \Delta mz_t \\ \Delta o_t \\ \Delta y_t \\ \Delta^2 p_t \\ \Delta xr_t \\ \Delta i_t \end{bmatrix} = \sum_{i=1}^{k-1} \Gamma_i \begin{bmatrix} mz_{t-1} \\ \Delta o_{t-1} \\ \Delta y_{t-1} \\ \Delta^2 p_{t-1} \\ \Delta xr_{t-1} \\ \Delta i_{t-1} \end{bmatrix} + \alpha\beta' \begin{bmatrix} mz_{t-1} \\ o_{t-1} \\ y_{t-1} \\ \Delta p_{t-1} \\ xr_{t-1} \\ i_{t-1} \end{bmatrix} + \Phi_1 \begin{bmatrix} D_{t,1} \\ \vdots \\ D_{t,d} \\ 1 \end{bmatrix} + \begin{bmatrix} \hat{\varepsilon}_{mz,t} \\ \hat{\varepsilon}_{o,t} \\ \hat{\varepsilon}_{y,t} \\ \hat{\varepsilon}_{\Delta p,t} \\ \hat{\varepsilon}_{xr,t} \\ \hat{\varepsilon}_{i,t} \end{bmatrix} \quad (2)
$$

where α is a $p \times r$ and β' is an $r \times p$ matrix with $r \leq p$ vector of stationary

---

[10]All prices have been deflated using the price level as previously described.

[11]For Soybean and Sugar the Unrestricted VECM is the same as in Equation 2.

cointegrating relations[12].

According to Juselius (2006), any misspecification of the model assumption will have fundamental effects on the parameters estimates and interpretation from the model. Therefore, it is essential to apply misspecification tests that shed light on the model constancy and normality of the residuals as assumed in the VAR model. The unrestricted model for all three commodities presents a number of misspecifications, which are primarily derived from the non-normality of the residuals in addition to ARCH effects (See Table 2, 3 and 4). The hypothesis of normality of the residuals for each individual variable is also rejected for most cases due to primarily the excess skewness and kurtosis the variables present. Nevertheless, Juselius (2006) argues that one can achieve a well specified statistical model by modifying some of the initial specifications of the UVAR using the following structure:

- including intervention dummies to account for significant political or institutional changes;

- conditioning on weakly exogenous variables;

- splitting or changing the sample period;

- checking the information set by adding new variables;

- examining the parameter constancy of the model (e.g. structural shifts in the model parameters);

- checking the adequacy of the measurements of the chosen variables;

- increasing the lag length.

In an effort to achieve a well-specified model, the first step we take is to detect periods of instability and structural changes in individual series by detecting those residuals larger than three standard deviations ($\pm 3\hat{\sigma}$). Table 5 provides a list of intervention dummy structural shifts that have been detected and has helped define a well specified VAR model for all three specification models. The periods of instability detected here coincide with major price fluctuations in commodity markets from the early 80's and early 2000's, monetary policy interventions (e.g. expansion the Federal Reserve Bank of the U.S. in 2003) as well as macroeconomic instability related to the recent global financial crisis of 2007/08 to the present.

Another possible source of misspecification is the inclusion of weakly exogenous variables in the models. Weakly exogenous variables do not have a long-run effect in the variables of interest and therefore tests should be performed in order to identify these variables. In Table 6 we present the tests for weakly exogenous variables for

---

[12]In the case of $r = n$, then Π is a singular matrix and we can model all variables as stationary; Φ_1 is a matrix of $p \times d$, where d is the number of intervention dummies in addition to the constant.

all possible rank selections. From the test, we conlcude that, for all possible rank (in all three specifications) the real exchange rate and the nominal short-run interest rate appear to be weakly exogenous. Although, test results for 'Model 1' are not as determinant as is the case on the other two models, we argue that in the first place we do not expect these variables to have a long-run effect on the agricultural commodities and also there is no reason to believe the system contains more than three cointegrating vectors. Thus, we focus the analysis and in the evidence presented on the test results related to a rank of less than four. For this reason, we condition both, the real exchange rate and nominal short-run interest rate, to be weakly exogenous in the model and thus to remain outside the cointegrating space.

After determining possible weakly exogenous variables and considering extraordinary events (as intervention dummies) in the models, the distributions of the residuals became closer to a normal distribution than in the initial estimated model. In Tables 2-4 we show evidence that the empirical VAR (EVAR) do not suffer from any serial autocorrelation and the individual variables with signs of non-normality have significantly improved from the unrestricted model. Even though in Tables 2-4 the models present signs of non-normality in the residuals, this is primarily due to the excess kurtosis, none of these present a threat to the properties of the estimates (Juselius, 2006). Therefore, the preferred specified models, shown in these tables, consist of VAR($k = 2$, $p = 4$) with interventions dummies and structural shifts as specified above as well as considering the real exchange rate and nominal short-run interest rate as weakly exogenous in the model with the intercept as the only deterministic component. As before, we can rewrite the EVECM for Maize[13] as:

$$
\begin{bmatrix} \Delta mz_t \\ \Delta o_t \\ \Delta y_t \\ \Delta^2 p_t \end{bmatrix} = \Gamma_1 \begin{bmatrix} \Delta mz_{t-1} \\ \Delta o_{t-1} \\ \Delta y_{t-1} \\ \Delta^2 p_{t-1} \\ \Delta xr_{t-1} \\ \Delta i_{t-1} \end{bmatrix} + \alpha\beta' \begin{bmatrix} mz_{t-1} \\ o_{t-1} \\ y_{t-1} \\ \Delta p_{t-1} \\ xr_{t-1} \\ i_{t-1} \\ D_{s_{0403}} \end{bmatrix} +
$$
$$
\mathbf{A}_1 \begin{bmatrix} \Delta xr_t \\ \Delta i_t \end{bmatrix} + \Phi_1 \begin{bmatrix} D_{t,1} \\ \vdots \\ D_{t,d} \\ \Delta D_{s_{0403}} \\ 1 \end{bmatrix} + \begin{bmatrix} \hat{\varepsilon}_{\{mz,sb,s\},t} \\ \hat{\varepsilon}_{o,t} \\ \hat{\varepsilon}_{y,t} \\ \hat{\varepsilon}_{\Delta p,t} \end{bmatrix}
$$

(3)

where in this case, the CVAR system contains four variables ($p = 4$) and as before Γ_1 is a $p \times 6$, α is a $p \times r$ matrix while β' is $r \times 7$[14] with $r \leq p$ vector of

---

[13]For Soybean and Sugar the Empirical VECM is the same as in Equation 3.

[14]From Equation 3, β' is $r \times 7$ matrix for both Maize and Soybean specifications while for Sugar, β' is $r \times 6$ since we have not determined any shift in the cointegrating space. $\mathbf{A}_1$ has dimensions

stationary cointegrating relations and the real exchange rate and nominal short-run interest rate restricted as weakly exogenous.

Thus, the three models we empirically test are those associated with our preferred specifications described in Equation 3. In this case, each system has a constant restricted to the cointegrating space, in addition to the two weakly exogenous variables (i_t and $\Delta^2 p_t$), a series of intervention dummies outlined in Table 5 with no shift dummy for the specification associated with Model 3 (Sugar) and including the shift dummy for April 2003 (for Model 1- Maize) as well as a shift dummy for August 2004 (for Model 2-Soybean).

5.1. Determining the rank.

According to Juselius (2006), once we have identified a well-specified empirical model then we we can test for the rank of the system. We proceeded to test the rank for each and everyone of the specified empirical models as in Equation 3. In determining the choice of the cointegrating rank we have considered: (1) the trace test for cointegrating rank; (2) the critical values of the α coefficients; (3) the recursive graphs of the trace statistic; (4) the graphs of the cointegrating relations as well as the economic interpretability of each system.

In Table 7 we present a summary of the rank test for all three models. based on the trace test for cointegration rank. It is important to note that in all three cases all methods arrive to the same conclusion and that we have used the Bartlett trace test corrected for small sample behaviour and dummies as proposed by Johansen (2000, 2002). The top of Table 7 presents the rank test conducted for Model 1 where we are able to reject the null hypothesis of one and two unit roots (at the 10% level) and fail to reject three unit roots in the system. Thus, we conclude that for the system including Maize (as one of the agricultural commodity price) we have three cointegrating relationships, which is consistent with the equations associated with the price level, global economic activity and that of maize in terms of macroeconomic variables and crude oil prices[15]. On the other hand, Model 2 from Table 7, corresponds to the specification related to the relationship between Soybeans, crude oil and the macroeconomy where we fail to reject two unit roots (at the 10% level) in the system. For Model 2, we conclude that we have two cointegrating relationships, which is consistent with the equations for Soybeans, oil prices and macroeconomic variables and the price level. The third and last model tested is that which corresponds to Sugar, crude oil and the same macroeconomic variables used before. From Table 7, the rank test indicates that we cannot reject the null hypothesis of 3 unit roots at the 10% level. Consequently, we conclude that for Model 3 we have a total of three possible cointegrating relationships.

---

of $p \times m$, where m corresponds to the number of weakly exogenous variables in the system (two in this case).

[15]This hypothesis is later tested and results will be presented.

5.2. Stability of the system.

After choosing the rank of the system we want to check the constancy of the estimated long-run parameters (Juselius, 2006). The parameter constancy of the long-run parameters β can be tested by the 'Max Test for Constancy of β' by applying the Hansen and Johansen (1999) procedure as shown in Figure 3. This is a recursive test, which consists of comparing the likelihood ratio test with that of the likelihood function from each sub-sample with the restriction that the cointegration vectors estimated from the full sample fall within the space spanned by the estimated long-run vectors. The test statistic is χ^2 distributed with $p - r$ and r degrees of freedom.

Figure 3a shows the constancy test of the slope coefficients for Model 1. In this case the constancy test is safely below the rejection area for most of the sample period, except for the period between late 2006 and early 2007, which clearly coincides with the early part of the century commodity price instability and speculation. However, it is evident that even during the period of instability the relationship returns to safe levels below the rejection area of constancy. This provides strong evidence that even though there was significant instability during this period the long-run constancy of the parameters did not permanently change. Additionally, there exists another period of similar short-run rejection of constancy in the long-run parameters, which coincides with the period of great macroeconomic volatility due to the great recession in late 2008. In summary, Model 1 only suffers from non-constant parameters in the short-run, but not in the long-run structure. This is given by the stability derived from the analysis from both the R and X-form models[16]. The only time where the model present periods of relative instability is during the peak of the commodity price shock where commodities suffered from speculative attacks. Nevertheless, soon after it appears that the system returns back to its long-run equilibrium. After this analysis there exists very little evidence that the instability experienced in the 2006/2007 commodity price market did not permanently affect the long-run parameters of the maize, crude oil price and macroeconomic variables. Therefore, the recursive constancy parameter presents strong evidence in favour of constant long-run parameters in the cointegrating relationship of maize and crude oil prices and the remaining macroeconomic variables.

Figure 3b shows the the parameter constancy of the long-run parameters β for the soybean relationship. As in the case for maize, also for soybean the constancy test is safely below the rejection area for most of the sample period, except for the period between late 2006 and early 2007, which clearly coincides with the early part of the century commodity price instability and speculation. Also, the long-run parameters present, as expected, instability around the peak of the commodity price boom, but soon returning to levels below the rejection area before showing another period of instability associated with the financial crisis of 2008 before returning to levels safely below the rejection line. In this case, soybean presents periods of temporary instability around late 2002 and early 2003, which could be associated with

---

[16]The X-form of the Model 1 contains the full model version; R-form is the concentrated version, which only contains the short-run variations.

the monetary expansion that the U.S. employed during this period. In both models for Maize and Soybeans, the long-run parameters present similar periods of instability, which essentially differ by as much as a few months from those previously highlighted. This is consistent with the literature which highlights that the peaks in the individual commodity price shock did not occur all at the same time (Headey and Fan, 2008; Gilbert, 2010). At the same time, these periods show evidence of a great deal of temporary instability due to recent events alone, but in general the relationship appears to be provide signs of stability by the end of the sample. Thus, the recursive parameter test shows that the cointegrating relationship estimated here presents long-run parameter constancy despite the periods of instability described.

Figure 3c shows the the parameter constancy of the long-run parameters β for the model associated with world sugar prices. As in the case for maize, also for soybean the constancy test is safely below the rejection area for most of the sample period, except for the period between late 2007 and early 2008, which coincides with the period of instability associated with the financial crisis of 2008 before returning to levels safely below the rejection line. Sugar long-run parameters, as supposed to maize and soybeans, did not suffer from so much instability during the same period of time. Nevertheless, as in the previous cases the recursive parameter test shows that cointegrating relationships for sugar also present long-run constant parameters.

5.3. Long-Run estimates for Models 1-3

In this section, we are going to impose restriction on the empirical specifications for each one of the models of interest in order to identify the long-run structure of the system. As suggested by Juselius (2006), we can use two approaches in order to obtain a correct identification: (1) We can impose just-identifying restrictions on the β vectors and subsequently we impose further restrictions by restricting insignificant coefficients in β; (2) We can also test the theoretical relations searching for an identified structure by combining stationary theoretical relations.

Recall that for Maize we have determined the empirical VAR specification as a VAR(2) model with $r = 3$. That is, we have identified in this model three cointegrating relationships. We proceed to impose restrictions on the long-run parameters in order to identify each cointegrating equation and interpret its economic meaning. In our case, we suspect from previous analysis that inflation (Δp_t) appears to be stationary and together with the index of global economic activity (y_t)[17] represent the first and second stationary relationships. The third cointegrating relationship is suspected to be that of interest, which consists of the real price of Maize (mz_t) explained by fluctuations in the real price of crude oil (o_t), the real exchange rate (xr_t) and real short-term interest rate ($i_t - \Delta p_t$) as exogenous[18]. From Table 8 we can see that we fail to reject LR-test with a p-value of 0.237 and thus conclude that

---

[17]By construction the index of global economic activity is stationary and as such it should be reflected as a stationary relationship with itself or other variables.

[18]An I(2) analysis was also conducted in order to corroborate the correct price transformations and no I(2) variables were found in the model given the specification used in this model.

the restrictions imposed correctly identify three cointegrating equations and proceed to analyze the economic meaning of these long-run relationships.

Table 8 shows that as suspected the first cointegrating relationship is associated with the stationary nature of the inflation rate. The second linear combination is that of the index of global economic activity, together with the real price of crude oil. This equation predicts that real shocks to the crude oil price have a negative impact on the global economic activity index. These results are corroborated in the literature by Kilian (2009) and He et al. (2010). Our estimates indicate that approximately a dollar increase in the world real price of crude oil has a long-run negative impact of about 30% in the global output, with a very low equilibrium towards the long-run relationship of 2% a month. These estimates emphasize the magnitude and long-lasting effects of shocks the world real energy market on real economic activity. Moreover, we also find substantial evidence indicating a shift in the level of global output by approximately 50% since 2003, which might explain why a number of authors have blamed the unforeseen global demand as a factor in the commodity price crisis. However, in our case there is no conclusive evidence supporting this finding.

From Table 8, the long-run relationship of interest is that associated with $\hat{\beta}_3$, where the real price of maize is expressed as a function of the remaining variables. In this case, all the coefficients have the expected sign and are highly statistically significant. The estimated long-run coefficient of the real world price of crude oil is estimated to be 0.889, which after testing the null hypothesis of a one-to-one long-run effect we fail to reject at the 1% level with a p-value of 0.727[19]. Therefore, from this analysis we conclude that a one dollar increase in the real long-run world price of crude oil is associated with a one dollar increase in the long-run real world price of maize. Looking at the coefficient alone, the implications of this estimate are profound in the Maize-Oil relationship. However, it's important to highlight that the index of global economic activity (aggregate demand) is not included in the cointegrating space. In order to understand the extent to which shocks to real crude oil prices on affect the real price of Maize, we would need to impose further restrictions to determine the common long-run trends, which analysis is presented further ahead.

In addition, also from Table 8, as the U.S. dollar appreciates with respect to major currencies, it is estimated to have a negative impact on the long-run real price of maize. This finding falls within our expectations since we (and other authors) have argued the possibility that in the long-run as the index of U.S. dollar with major currencies appreciates with respect to the major currencies, the real price of maize (USD/ton) decreases with respect to these currencies. Furthermore, we are able to accept the presence of long-run homogeneity between inflation and the nominal short-run interest rate (i.e. the real interest rate). This finding implies that increases in the real interest rate have a negative and significant impact on the real price of maize. This finding reflects the trade-off between the returns to

---

[19]The test corresponds to a $\chi^2(7) = 4.447$.

capital and investment from holding commodities. On the other hand, the index for global economic activity (as a proxy for the world business cycles) appears to have no effect on the long-run real price of maize. This finding is somewhat surprising since the index for global economic activity is thought to be a proxy for the world business cycles and increasing aggregate demand would in theory have a positive effect in the long-run real price of maize. However, it is possible that the way in which this variable has been constructed and detrended does not necessarily correspond to that variation associated with the commodity market of maize. Finally, our estimates show that since April 2003, the real price of maize has increased in real terms by approximately 63% compared to the entire previous period, which is a substantial increase considering this has occurred after a lapse of eight years or less.

The second area of interest is the speed of adjustment towards the equilibrium (α coefficients in Table 8). By closely examining α_3 in Table 8, it shows that all variables, except the growth rate of inflation ($\Delta^2 p_t$) are adjusting forces towards the long-run equilibrium of the real price of maize. Moreover, maize itself, appears to be adjusting to its long-run equilibrium at a very slow pace of approximately 4% a month from a disequilibrium state. However, analyzing maize's price adjustment coefficient alone can be misleading since oil prices and Maize are both helping the system return to its long-run equilibrium. Furthermore, these adjustment coefficients are not implausible given the nature of commodity price markets and are very similar to those previously estimated in the literature (see for example Cashin et al. (2004)).

On the other hand, the long-run estimates for the soybean specification are presented in Table 9. We determined two cointegrating relationships for this model and similar to Maize, the first relationship is that of inflation and the index for global economic activity. The second cointegrating relationship (that is $\hat{\beta}_2$ from Table 9) is that associated with the long-run fluctuations between the real world price of crude oil and macroeconomic variables. The estimated long-run coefficient for the real world price of crude oil is 0.937, which implies that approximately a one-to-one relationship between the long-run real world price of soybeans and crude oil exists. Additionally, we have tested the restriction of a one-to-one long-run relationship between these two variables and we fail to reject the null hypothesis with a p-value of 0.364. Therefore, we cannot reject the hypothesis that in the long-run, an increase in the real world price of crude oil has a one-to-one relationship with the real world price of soybeans. This finding is categorically equivalent to that found for the Maize model in that we cannot attribute solely this effect to the real world price of crude oil, but to a combination of demand and oil market effects. Also from Table 9, we obtain similar results as to those from the maize relationship. As in the case for maize, also soybeans present a negative relationship between the real exchange rate of the U.S. dollar and major currencies. Additionally, since August 2004 to December 2012, the real world price of soybeans has increased approximately 90% compared to increases in the previous period (1982-2003). In summary, our estimates for the Soybean cointegrating relationship show that the real world price of crude oil, the U.S. exchange rate and the real interest rate are important determinants of the long-run dynamics and its long-run price stability.

The long-run results for the third specification are shown in Table 10. As supposed to the previous two relationships, sugar appears not to have a long-run relationship with crude oil and only the real U.S. exchange rate is the only variable from the model to matter in the long-run world price of sugar dynamics .

5.4. Common driving trends

In this section we will try to identify the common stochastic trends of the system. In that sense, for example, we are interested in determining to what extent shocks to the real world price of crude oil have an impact on both Maize and Soybean prices. In the case of Maize, we have identified three cointegrating relationships and p-r=1 common stochastic trends. From Table 11, the common stochastic trends associated with the real world price of Maize are associated with itself and long-run shocks to the real price of crude oil alone. More precisely, permanent shocks to the real price of crude oil have approximately a permanent effect and are transmitted to the real price of maize by a factor of 0.67, everything else being constant. On the other hand, it appear that inflationary shocks only have a short-run effect on the real world price of maize. In other words, only shocks to the price of crude oil and maize itself have a long-run permanent effect in the price of maize. This is an important conclusion to draw since the literature on the effects of crude oil prices on agricultural commodities (particularly that of biofuel-commodities) argue that since the early part of the century, and as a consequence of record high prices of oil, agricultural commodities (such as maize) are more susceptible to crude oil price shocks.

In Table 11 (bottom), we present the stochastic trends associated with the soybeans specification. Here, we are able to identify two stochastic trends primarily derived from shocks to the real price of crude oil and marginally from the global economic activity. In fact, permanent shocks to real crude oil prices appears to have a permanent effect by a factor of approximately 0.67, which is the same as estimated for Maize. Additionally, permanent shocks to the real global economic activity index appears to also have a marginal permanent effect on the world price of soybeans by a factor of approximately 0.13 (at the 10% level). On the other hand, shocks to inflation and the exchange rate do not appear to have a long-run effect to the real world price of soybeans. Thus, in summary, shocks to crude oil, the global economic activity index and soybeans have a permanent effect on the real world price of soybeans.

6. Policy Implications

Food security is among the fundamental steps outlined in the Millennium Development Goal (MDG) to halved the proportion of people suffering from extreme poverty and hunger around the world by 2015. In order to achieve this goal it is essential to understand the factors and dynamics affecting global food prices. One important factor in determining global food prices has been the recent implementation of biofuel energy policies, which have been blamed (together with market

speculation) as the leading cause to the world food price crisis of 2006/07 (Nazli-oglu and Soytas, 2012).Consequently, in order to achieve these goals, any global food policy needs to consider the transmission mechanisms from crude oil to global food prices as well as the short and long-run economic and market forces influencing both energy and food commodities simultaneously.

A global policy concern in terms of food security has been the introduction of corn ethanol in the early years of the century. And this concern is founded from our results where we show that global real prices of maize and soybeans share a one-to-one long-run relationship with real crude oil prices. Zilberman et al. (2013) show that given an increase in biofuel prices due increases in the gasoline market, food prices also increase. This is, since the increase in the price of gasoline exacerbates the incentives to produce more biofuel while channeling out resources from food production towards the energy market. This in turn, causes further upward pressure on food prices. This is particularly problematic for policy makers since our empirical evidence shows that only real shocks to real crude oil prices are transmitted to both corn and soybean real prices. Therefore, policy makers and advocates for biofuel policies need to closely consider and weight out the factors that affect real crude oil prices which then is transmitted to food prices and influence biofuel markets, to minimize the unintended costs of increasing biofuel production.

Furthermore, fluctuations in the global market for agricultural products can also have significant fiscal impact across countries. Net exporters countries of oil, maize and soybeans might see significant improvements in their terms of trade and rise in revenues from higher international prices, while net importers of these commodities might see an equally, but worsening scenario. Even in commodity exporting countries higher food prices will see poor consumers being affect, particularly in developing countries, where proportion of income destined for food in much higher than in developed nations. This impact has been addressed through export bans and internal price controls, but as markets become more internationally linked, these policies will become less efficient. Therefore, the need for policy makers to pay close attention to the economic fundamentals affecting the price dynamic of energy markets and commodity prices as outlined in our research in conjunction with increasing investment in agricultural production will assist in making these policies far more resilient.

Finally, out results indicate that investors can use information in the global crude oil market to infer directional price moevements in the maize and soybeans global markets. Out long-run results indicate a one-to-one long-run relationship between real crude oil prices with maize and soybeans since 1982 to 2012. Therefore, traders can benefit from the information in the crude oil market to understand fluctuations in the markets of these agricultural goods. Additionally, investors can also gain information from movements in the macroeconomic and monetary policies as well as fluctuations in the currency markets for U.S. dollars.

7. Conclusion

This study examines the long-run relationship between the real world price of maize, soybeans and maize with the real world price of crude oil and a series of macroeconomic variables. We apply a cointegration analysis for a monthly series from January 1982 until December 2012. The main empirical results support a strong relationship between fundamentals and the world price of Maize and Soybeans. Particularly, we document significant causal long-run relationships between these agricultural commodities and the real world price of crude oil, the real interest rate and the real U.S. exchange rate. In fact, we show that in the long-run, crude oil prices and these agricultural commodities share a one-to-one relationship. In other words, a one-percent increase in the price of real crude oil is associated with a one-percent increase in the price of maize and soybeans. Despite the literature suggesting the neutrality between agricultural commodities and energy prices (mainly crude oil) our findings show that there is a very strong long-run causal relationship between these and macroeconomic factors. Moreover, we find that permanent shocks to crude oil prices are transmitted to both maize and soybeans by a factor of 0.67 for both commodities. In this aspect, our study contradicts the results from Yu et al. (2006); Campiche et al. (2007); Zhang et al. (2009) and Zhang et al. (2010) where no long-run (or partial) relationship was estimated among these variables. These results emphasize that in order to produce informative estimates on a very complex long-run relationship, as is the case of agricultural commodities, one needs to consider in a long-run data series more than just one possible factor affecting these price series. In addition, our results show that despite the instability associated with the period between 2007/08, the long-run relationship between crude oil and these agricultural commodities has remained stable during the entire sample period. These results contradict those found in studies by Harri et al. (2009); Ciaian and Kancs (2011a,b) and Natanelov et al. (2011) where instabilities in individual variables are used to model instability in the cointegrating space. Our findings suggest that despite the period of instability between 2007/08, the long-run estimated coefficients are stable along the sample period. The only period of instability associated with these relationships is around the peak of the commodity shock of 2007/08, which provides some evidence confirming the findings of Tang and Xiong (2010) and Pen and Sevi (2013) that speculation might have played an important role in explaining some of the price increases during this period. Moreover, our results also support the arguments that although the real interest rate and the U.S. exchange rate are cointegrated with these commodities, it's only permanent shocks to real crude oil prices that have a permanent effect on the behavior of these commodity prices. Therefore, we can conclude that fundamentals (in particular demand factors) are the crucial determinants of the long-run dynamics of both these two agricultural food commodities.

Appendix Figures

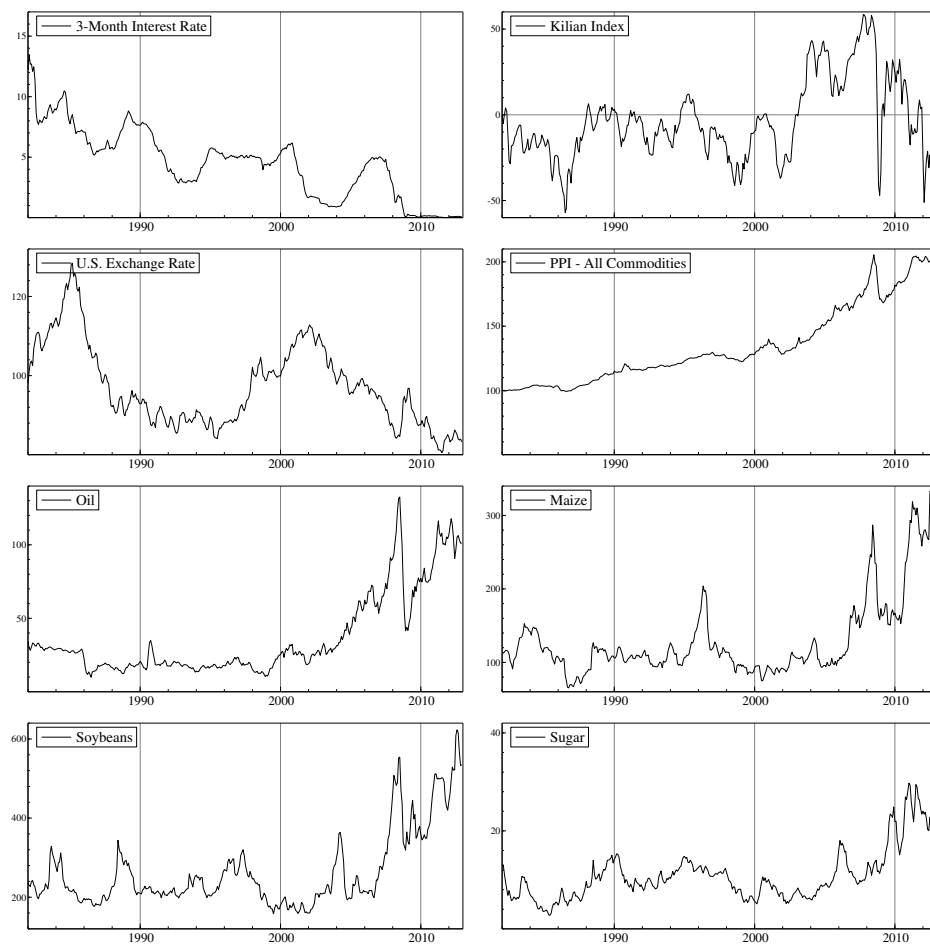Figure 1: Commodity prices and macroeconomic series in levels.

Figure 2: Logarithm of real commodity prices and transformed macroeconomic series.
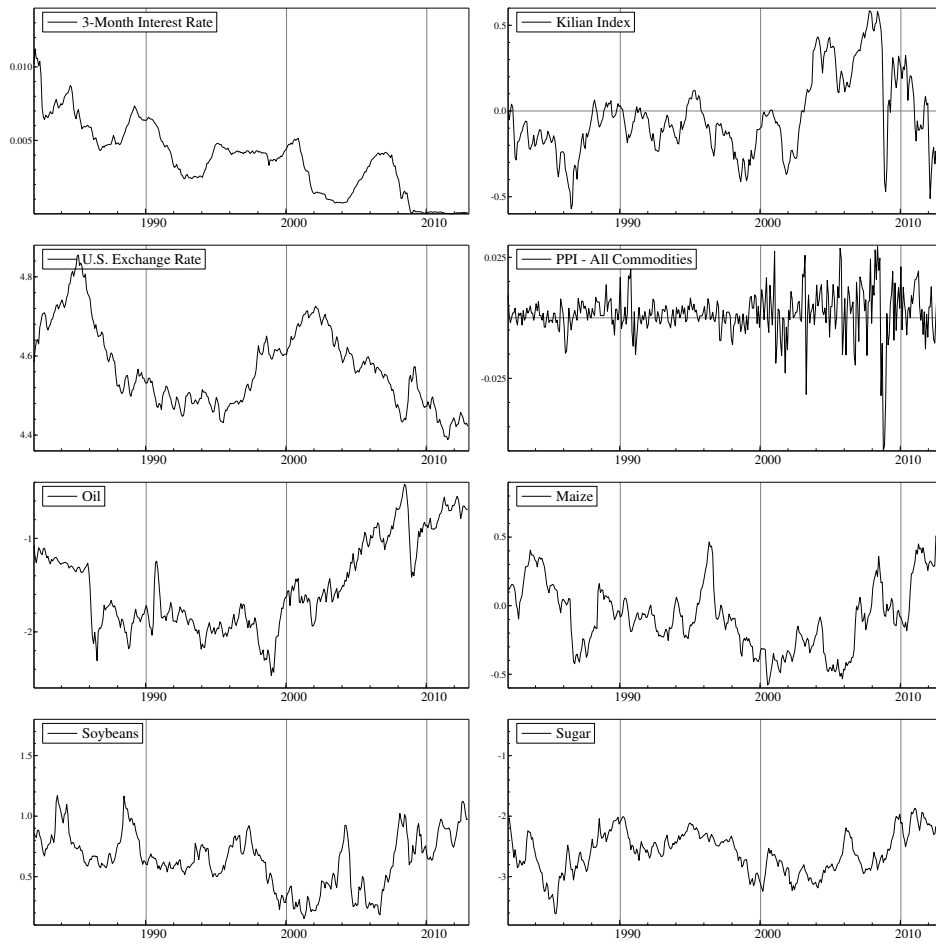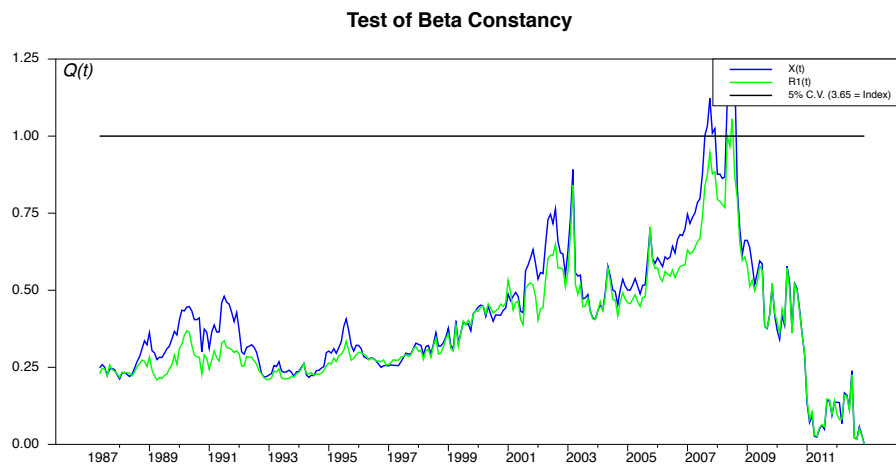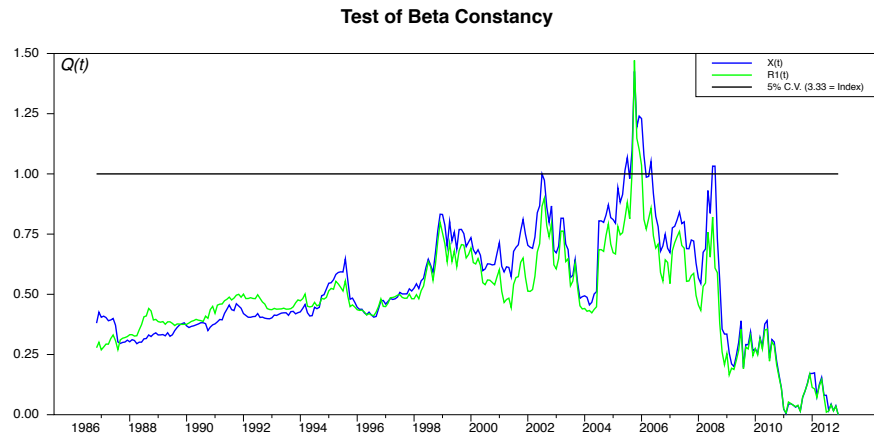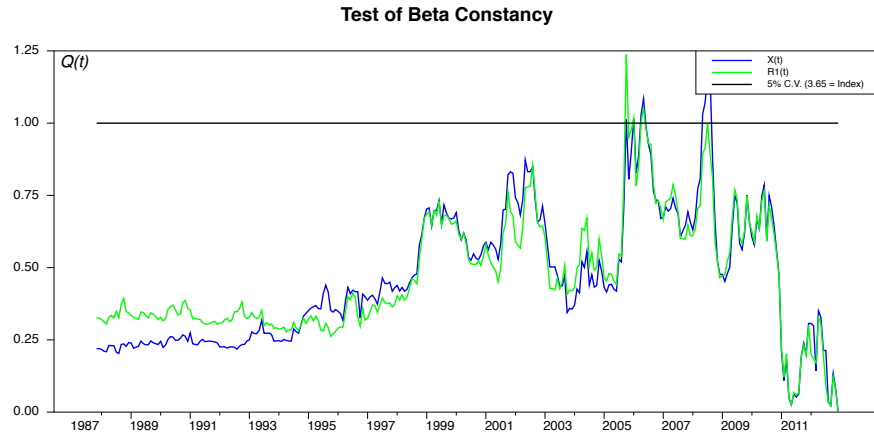
Figure 3: Recursive of constancy tests. The max test of β constancy.

Test of Beta Constancy

(a) Model 1 - Maize

Test of Beta Constancy

(b) Model 2 - Soybean

Test of Beta Constancy

(c) Model 3 - Sugar

Appendix Tables

Table 1: Data Definition and Source

Variable	Frequency	Range	Units	Source	Code
Maize (MZ_t)	Monthly	Jan 1982 Dec 2012	U.S. Dollars per Metric Ton	IFS	PZPIMAIZ
Soybean ($SOYB_t$)	Monthly	Jan 1982 Dec 2012	U.S. Dollars per Metric Ton	IFS	PZPISOYB
Sugar (S_t)	Monthly	Jan 1982 Dec 2012	U.S. Dollars per Metric Ton	IFS	PZPISUG
Crude Oil (O_t)	Monthly	Jan 1982- Dec 2012	U.S. Dollars per Barrel	IFS	PZPIOIL
PPI (P_t)	Monthly	Jan 1982- Dec 2012	Index (1982=100)	FRED	PPIACO
Three-Month T-Bill (I_t)	Monthly	Jan 1982- Dec 2012	Percentage	FRED	TB3MS
Trade Weighted USD Index (XR_t)	Monthly	Jan 1982- Dec 2012	Real Index (1997=100)	FRED	TWEXBMTH
Indx. Global Econ. Activity (Y_t)	Monthly	Jan 1982-	Index	Lutz Kilian	-

Table 2: Diagnostic statistics for Model 1 (Maize) on the unrestricted VAR($k = 3$) and empirical VAR ($k = 2$).

Multivariate Analysis

	Unrestricted VAR ($k = 3$)		Empirical VAR ($k = 2$)	
Tests for Autocorrelation:				
LM(1):	$\chi^2(36)$	$= 37.25$	$\chi^2(16)$	$= 23.20$
LM(2):	$\chi^2(36)$	$= 48.46^*$	$\chi^2(16)$	$= 22.47$
Test for Normality:	$\chi^2(12)$	$= 712.83^\ddagger$	$\chi^2(8)$	$= 156.68^\ddagger$
Test for ARCH:				
LM(1):	$\chi^2(441)$	$= 972.23^\ddagger$	$\chi^2(100)$	$= 310.41^\ddagger$
LM(2):	$\chi^2(882)$	$= 1688.63^\ddagger$	$\chi^2(200)$	$= 459.48^\ddagger$

Univariate Analysis

	ARCH[1]		Normality[2]		Skewness		Kurtosis	
Variables	U. VAR[3] ($k=3$)	E. VAR[4] ($k=2$)	U. VAR ($k=3$)	E. VAR ($k=2$)	U. VAR ($k=3$)	E. VAR ($k=2$)	U. VAR ($k=3$)	E. VAR ($k=2$)
Δmz_t	3.02	$6.39^\dagger$	$107.38^\ddagger$	$7.98^\dagger$	0.41	0.01	7.02	3.69
Δo_t	$15.54^\ddagger$	$30.61^\ddagger$	$55.15^\ddagger$	$7.14^\dagger$	0.23	-0.24	5.40	3.58
Δy_t	$27.17^\ddagger$	$29.44^\ddagger$	$194.34^\ddagger$	$49.72^\ddagger$	-0.64	0.01	9.72	5.15
$\Delta^2 p_t$	$42.31^\ddagger$	$49.30^\ddagger$	$128.67^\ddagger$	$69.94^\ddagger$	-0.82	-0.17	8.11	5.75
Δi_t	$53.78^\ddagger$		$171.20^\ddagger$		-1.89		17.45	
Δxr_t	1.57		$16.85^\ddagger$		0.31		4.14	

Notes: Significance at the the 1%, 5% and 10% level are denoted by $^{(\ddagger)}$, $^{(\dagger)}$ and $^{(*)}$ respectively.
[1] ARCH (k) is a test for autoregressive heteroskedasticity approximately distributed as $\chi^2(3)$ for the unrestricted VAR and $\chi^2(2)$ for the empirical VAR.
[2] Is the Doornik and Hansen (2008) test for univariate normality distributed as $\chi^2(2)$.
[3] Unrestricted VAR analysis.
[4] Empirical VAR analysis (restrictions imposed).

Table 3: Diagnostic statistics for Model 2 (Soybean) on the unrestricted VAR($k = 3$) and empirical VAR ($k = 2$).

	Multivariate Analysis	
	Unrestricted VAR ($k = 3$)	Empirical VAR ($k = 2$)
Tests for Autocorrelation:		
LM(1):	$\chi^2(36)$ = 37.23	$\chi^2(16)$ = 21.27
LM(2):	$\chi^2(36)$ = 50.38*	$\chi^2(16)$ = 25.80
Test for Normality:	$\chi^2(12)$ = 680.93‡	$\chi^2(8)$ = 178.90‡
Test for ARCH:		
LM(1):	$\chi^2(441)$ = 1025.51‡	$\chi^2(100)$ = 391.92‡
LM(2):	$\chi^2(882)$ = 1694.61‡	$\chi^2(200)$ = 564.34‡

	Univariate Analysis							
	ARCH[1]		Normality[2]		Skewness		Kurtosis	
Variables	U. VAR[3] ($k = 3$)	E. VAR[4] ($k = 2$)	U. VAR ($k = 3$)	E. VAR ($k = 2$)	U. VAR ($k = 3$)	E. VAR ($k = 2$)	U. VAR ($k = 3$)	E. VAR ($k = 2$)
Δsb_t	3.51	0.542	57.43‡	7.59†	0.66	0.21	6.16	3.64
Δo_t	14.44‡	43.97‡	55.00‡	7.73†	0.23	-0.28	5.40	3.60
Δy_t	26.50‡	27.66‡	198.13‡	60.98‡	-0.64	0.13	9.72	5.15
$\Delta^2 p_t$	49.53‡	31.42‡	118.01‡	82.50‡	-0.82	-0.29	8.11	6.20
Δi_t	54.67‡		168.30‡		-1.89		17.45	
Δxr_t	1.57		16.93‡		0.31		4.14	

Notes: Significance at the the 1%, 5% and 10% level are denoted by (‡), (†) and (*) respectively.
[1] ARCH (k) is a test for autoregressive heteroskedasticity approximately distributed as $\chi^2(3)$ for the unrestricted VAR and $\chi^2(2)$ for the empirical VAR.
[2] Is the Doornik and Hansen (2008) test for univariate normality distributed as $\chi^2(2)$.
[3] Unrestricted VAR analysis.
[4] Empirical VAR analysis (restrictions imposed).

Table 4: Diagnostic statistics for Model 3 (Sugar) on the unrestricted VAR($k = 3$) and empirical VAR ($k = 2$).

	Multivariate Analysis	
	Unrestricted VAR ($k = 3$)	Empirical VAR ($k = 2$)
Tests for Autocorrelation:		
LM(1):	$\chi^2(36)$ = 31.30	$\chi^2(16)$ = 22.44
LM(2):	$\chi^2(36)$ = 41.29	$\chi^2(16)$ = 23.70
Test for Normality:	$\chi^2(12)$ = 611.75‡	$\chi^2(8)$ = 277.59‡
Test for ARCH:		
LM(1):	$\chi^2(441)$ = 1031.26‡	$\chi^2(100)$ = 381.80‡
LM(2):	$\chi^2(882)$ = 1636.77‡	$\chi^2(200)$ = 528.86‡

	Univariate Analysis							
	ARCH[1]		Normality[2]		Skewness		Kurtosis	
Variables	U. VAR[3] ($k = 3$)	E. VAR[4] ($k = 2$)	U. VAR ($k = 3$)	E. VAR ($k = 2$)	U. VAR ($k = 3$)	E. VAR ($k = 2$)	U. VAR ($k = 3$)	E. VAR ($k = 2$)
Δs_t	4.10	1.89	12.90‡	4.67*	0.14	0.04	3.93	3.49
Δo_t	12.84‡	30.67‡	50.40‡	5.77†	0.25	-0.14	5.28	3.54
Δy_t	29.60‡	33.46‡	197.66‡	128.64‡	-0.64	-0.36	9.83	7.43
$\Delta^2 p_t$	51.60‡	35.86‡	100.81‡	100.94‡	-1.03	-0.62	8.89	7.36
Δi_t	56.02‡		181.08‡		-1.70		16.18	
Δxr_t	1.33		15.38‡		0.29		4.06	

Notes: Significance at the the 1%, 5% and 10% level are denoted by (‡), (†) and (*) respectively.
[1] ARCH (k) is a test for autoregressive heteroskedasticity approximately distributed as $\chi^2(3)$ for the unrestricted VAR and $\chi^2(2)$ for the empirical VAR.
[2] Is the Doornik and Hansen (2008) test for univariate normality distributed as $\chi^2(2)$.
[3] Unrestricted VAR analysis.
[4] Empirical VAR analysis (restrictions imposed).

Table 5: Description of intervention dummies.

Year	Month	Type	Events	Oil	Maize	Soybeans	Sugar	Inflation	Kilian Index
1983	May	Transitory	Drought in South Africa & Frost in Cuba				✓		
1983	August	Transitory	Drought Midwest U.S.			✓			
1986	February	Transitory	Collapse of Oil Prices	✓					
1988	June	Transitory	Drought Midwest U.S.		✓	✓			
1990	August	Transitory	Gulf War	✓					
1993	July	Transitory	Flooding Midwest U.S.			✓			
1996	Sept.	Transitory	Drought Midwest U.S.		✓				
2003	April	Level Shift	Peak U.S. Monetary Policy		✓				
2004	August	Level Shift	U.S. Monetary Policy			✓			
2008	June	Transitory	2007/08 commodity shock		✓				
2008	August	Transitory	Financial Crisis 2008					✓	
2009	January	Transitory	Financial Crisis 2008						✓
2012	January	Transitory	Eurozone Debt Crisis		✓				
2012	July	Transitory	Drought Midwest U.S.						✓

Table 6: Test of Weak Exogeneity.

| | | | | Model 1 - Maize | | | |
Rank	5% C.V.	mz_t	o_t	y_t	Δp_t	xr_t	i_t
1	3.841	2.655	0.155	7.575	83.203	0.565	0.000
2	5.991	7.146	7.426	20.000	104.543	5.521	4.011
3	7.815	10.228	10.706	34.094	114.159	5.654	7.322
4	9.488	13.191	11.013	38.394	115.385	5.654	11.190
5	11.070	16.707	12.722	41.870	118.324	7.601	15.019

| | | | | Model 2 - Soybean | | | |
		sb_t	o_t	y_t	Δp_t	xr_t	i_t
1	3.841	2.613	88.546	6.126	0.280	0.501	0.367
2	5.991	5.713	99.262	8.558	5.929	4.217	3.804
3	7.815	5.928	105.308	16.573	8.177	4.265	4.461
4	9.488	11.105	110.439	21.834	11.837	5.746	4.695
5	11.070	15.652	113.066	25.762	11.847	9.106	8.601

| | | | | Model 3 - Sugar | | | |
		s_t	o_t	y_t	Δp_t	xr_t	i_t
1	3.841	0.343	119.285	5.298	1.686	0.710	0.498
2	5.991	15.762	120.214	5.325	1.687	1.298	1.035
3	7.815	19.732	123.252	12.465	1.748	1.300	1.125
4	9.488	25.620	128.403	18.237	3.319	3.943	3.298
5	11.070	28.547	128.409	21.039	3.585	8.000	6.816

Table 7: The I(1) rank analysis for all models based on the simulated critical values.

| | Model 1 - Maize | | | |
$Rank$	$Trace$	$Trace_{Bart.}$	$C_{.95}$	P-$Value$
$r = 0$	223.228	218.247	71.302	$0.000^{(\ddagger)}$
$r \leq 1$	67.613	66.187	48.621	$0.000^{(\ddagger)}$
$r \leq 2$	28.482	27.815	30.667	$0.095^{(*)}$
$r \leq 3$	6.193	6.058	15.039	0.638
	Model 2 - Soybean			
$r = 0$	195.005	190.718	71.550	$0.000^{(\ddagger)}$
$r \leq 1$	51.255	50.209	48.556	$0.034^{(\dagger)}$
$r \leq 2$	27.024	26.053	29.738	0.133
$r \leq 3$	7.616	7.349	14.984	0.494
	Model 3 - Sugar			
$r = 0$	211.461	207.012	53.358	$0.000^{(\ddagger)}$
$r \leq 1$	54.933	53.863	35.371	$0.000^{(\ddagger)}$
$r \leq 2$	21.196	20.760	20.874	$0.047^{(\dagger)}$
$r \leq 3$	2.529	2.451	8.327	0.707

Notes: Simulated Rank Test distribution. Significance at the the 1%, 5% and 10% level are denoted by $^{(\ddagger)}$, $^{(\dagger)}$ and $^{(*)}$ respectively.

Table 8: Model 1 - Identified long-run structures (P-values in brackets).

	mz_t	Δp_t	y_t	o_t	xr_t	i_t	Ds_{2003}	μ_0
				Model 1 - Maize				
$\hat{\beta}_1$	0.000 [NA]	1.000 [NA]	0.000 [NA]	0.000 [NA]	0.000 [NA]	0.000 [NA]	0.000 [NA]	−0.002 [−3.301]
$\hat{\beta}_2$	0.000 [NA]	0.000 [NA]	1.000 [NA]	0.317 [3.588]	0.000 [NA]	0.000 [NA]	−0.490 [−5.525]	0.676 [4.227]
$\hat{\beta}_3$	1.000 [NA]	38.621 [5.382]	0.000 [NA]	−0.889 [−5.049]	1.415 [2.848]	−38.621 [−5.382]	0.625 [3.428]	−1.341 [−4.305]

	α_1	α_2	α_3
Δmz_t	0.602 [0.997]	0.010 [0.632]	−0.041 [−3.492]
$\Delta^2 p_t$	−0.772 [−8.728]	0.004 [1.519]	0.002 [0.935]
Δy_t	2.376 [3.864]	−0.101 [−6.112]	−0.023 [−1.950]
Δo_t	−1.506 [−1.759]	−0.002 [−0.097]	0.037 [2.235]

Test of Restricted Model: $\chi^2(3) = 8.011$ [0.237]

Table 9: Model 2 - Identified long-run structures (P-values in brackets).

	sb_t	Δp_t	y_t	o_t	xr_t	i_t	Ds_{2004}	μ_0
				Model 2 - Soybean				
$\hat{\beta}_1$	0.000 [NA]	1.000 [NA]	−0.007 [−3.440]	0.000 [NA]	0.000 [NA]	0.000 [NA]	0.000 [NA]	−0.002 [−4.562]
$\hat{\beta}_2$	1.000 [NA]	0.000 [NA]	0.000 [NA]	−0.937 [−4.662]	2.089 [3.582]	0.000 [NA]	0.899 [3.925]	−2.223 [−6.268]

	α_1	α_2
Δsb_t	−0.826 [−2.248]	−0.036 [−3.904]
$\Delta^2 p_t$	−0.687 [−11.069]	0.001 [0.871]
Δy_t	1.095 [2.462]	0.013 [1.184]
Δo_t	−0.265 [−0.447]	0.036 [2.411]

Test of Restricted Model: $\chi^2(6) = 7.624[0.267]$

Table 10: Model 3 - Identified long-run structures (P-values in brackets).

	s_t	Δp_t	y_t	o_t	xr_t	i_t	C
				Model 3 - Sugar			
β_1	1.000 [NA]	0.000 [NA]	0.000 [NA]	0.000 [NA]	2.545 [6.504]	0.000 [NA]	2.690 [62.269]
β_2	0.000 [NA]	1.000 [NA]	0.000 [NA]	0.000 [NA]	0.000 [NA]	0.000 [NA]	−0.002 [−4.395]
β_3	0.000 [NA]	0.000 [NA]	1.000 [NA]	0.000 [NA]	0.000 [NA]	0.000 [NA]	0.000 [NA]

	α_1	α_2	α_3
Δs_t	−0.096 [−5.660]	0.366 [0.567]	−0.005 [−0.280]
$\Delta^2 p_t$	0.001 [0.701]	−0.764 [−12.300]	0.004 [2.359]
Δy_t	−0.014 [−1.189]	1.352 [3.098]	−0.061 [−4.687]
Δo_t	−0.006 [−0.423]	−0.750 [−1.315]	0.004 [0.215]

Test of Restricted Model: $\chi^2(9) = 3.450[0.944]$

Table 11: The MA representation when β is restricted - The Long-Run Impact

	$\hat{\varepsilon}_{mz_t}$	$\hat{\varepsilon}_{\Delta p_t}$	$\hat{\varepsilon}_{y_t}$	$\hat{\varepsilon}_{o_t}$
	Model 1 - Maize			
mz_t	0.571	−0.827	0.014	0.674
	[3.384]	[−1.361]	[0.075]	[5.188]
$\Delta^2 p_t$	−0.000	0.000	−0.000	−0.000
	[−3.384]	[1.361]	[−0.075]	[−5.188]
y_t	−0.204	0.295	−0.005	−0.240
	[−3.384]	[1.361]	[−0.075]	[−5.188]
o_t	0.643	−0.931	0.015	0.758
	[3.384]	[−1.361]	[0.075]	[5.188]
	Model 2 - Soybean			
sb_t	0.690	−0.879	0.133	0.678
	[3.904]	[−1.356]	[1.788]	[3.986]
$\Delta^2 p_t$	0.003	0.013	0.011	−0.001
	[1.853]	[1.965]	[13.718]	[−0.567]
y_t	0.482	1.875	1.506	−0.142
	[1.853]	[1.965]	[13.718]	[−0.567]
o_t	0.737	−0.938	0.142	0.723
	[3.904]	[−1.356]	[1.788]	[3.986]

Abbott, P. C., Hurt, C., Tyner, W. E., Mar. 2009. What's driving food prices? Issue Reports 48495, Farm Foundation.

Abbott, P. C., Hurt, C., Tyner, W. E., Jul 2011. Whats driving food prices in 2011? Issue reports, Farm Foundation.

Baek, J., Koo, W. W., September 2010. Analyzing factors affecting u.s. food price inflation. Canadian Journal of Agricultural Economics 58 (3), 303320.

Baffes, J., September 2007. Oil spills on other commodities. Resources Policy 32 (3), 126–134.

Barsky, R. B., Kilian, L., Jul. 2001. Do we really know that oil caused the great stagflation? a monetary alternative. NBER Working Papers 8389, National Bureau of Economic Research, Inc.

Bernanke, B. S., Gertler, M., Watson, M., 1997. Systematic monetary policy and the effects of oil price shocks. Brookings Papers on Economic Activity 28 (1), 91–157.

Cairns, A. P., Meilke, K. D., Apr 2012. The next-11 and the brics: Are they the future markets for agrifood trade? Working Papers 122737, Canadian Agricultural Trade Policy Research Network.

Campiche, J. L., Bryant, H. L., Richardson, J. W., Outlaw, J. L., 2007. Examining th between petroleum prices ande evolving correspondence agricultural commodity prices. 2007 Annual Meeting, July 29-August 1, 2007, Portland, Oregon TN 9881, American Agricultural Economics Association (New Name 2008: Agricultural and Applied Economics Association).

Carter, C. A., Rausser, G. C., Smith, A., October 2011. Commodity booms and busts. Annual Review of Resource Economics 3 (1), 87–118.

Cashin, P., Cspedes, L. F., Sahay, R., 2004. Commodity currencies and the real exchange rate. Journal of Development Economics 75 (1), 239 – 268.

Céspedes, L. F., Velasco, A., December 2012. Macroeconomic performance during commodity price booms. IMF Economic Review 60 (4), 570–599.

Chakravorty, U., Hubert, M.-H., Moreaux, M., Nostbakken, L., Mar 2012. Do biofuel mandates raise food prices? Economics Working Paper Archive (University of Rennes 1 & University of Caen) 201214, Center for Research in Economics and Management (CREM), University of Rennes 1, University of Caen and CNRS.

Chaudhuri, K., 2001. Long-run prices of primary commodities and oil prices. Applied Economics 33 (4), 531–538.

Chen, X., Khanna, M., 2013. Food vs. fuel: The effect of biofuel policies. American Journal of Agricultural Economics 95 (2), 289–295.

Ciaian, P., Kancs, d., October 2011a. Food, energy and environment: Is bioenergy the missing link? Food Policy 36 (5), 571–580.

Ciaian, P., Kancs, d., January 2011b. Interdependencies in the energy-bioenergy-food price systems: A cointegration analysis. Resource and Energy Economics 33 (1), 326–348.

Colin Carter, G. R., Smith, A., 2012. The effect of the u.s. ethanol mandate on corn prices, working Paper.

Condon, N., Klemick, H., Wolverton, A., August 2013. Impacts of ethanol policy on corn prices: A review and meta-analysis of recent evidence, selected Paper prepared for presentation at the Agricultural & Applied Economics Associations 2013 AAEA & CAES Joint Annual Meeting.

Cooke, B., Robles, M., 2009. Recent food prices movements: A time series analysis. IFPRI discussion papers 942, International Food Policy Research Institute (IFPRI).

Doornik, J. A., Hansen, H., December 2008. An omnibus test for univariate and multivariate normality. Oxford Bulletin of Economics and Statistics 70 (s1), 927–939.

Frankel, J., Rose, A. K., Sep. 2010. Determinants of agricultural and mineral commodity prices. Working Paper Series rwp10-038, Harvard University, John F. Kennedy School of Government.

Frankel, J. A., October 2008. The effect of monetary policy on real commodity prices. In: Asset Prices and Monetary Policy. NBER Chapters. National Bureau of Economic Research, Inc, pp. 291–333.

Gilbert, C. L., 2010. How to understand high food prices. Journal of Agricultural Economics 61 (2), 398–425.

Gohin, A., Chantret, F., 2010. The long-run impact of energy prices on world agricultural markets: The role of macro-economic linkages. Energy Policy 38 (1), 333–339.

Hamilton, J. D., May 2009. Causes and consequences of the oil shock of 2007-08. Working Paper 15002, National Bureau of Economic Research.

Hamilton, J. D., February 2011. Historical oil shocks. Working Paper 16790, National Bureau of Economic Research.

Hamilton, J. D., Wu, J. C., 2014a. Effects of index-fund investing on commodity futures prices. Working Paper 19892, National Bureau of Economic Research.

Hamilton, J. D., Wu, J. C., April 2014b. Risk premia in crude oil futures prices. Journal of International Money and Finance Volume 42, 9–37.

Hansen, H., Johansen, S., 1999. Some tests for parameter constancy in cointegrated var-models. Econometrics Journal 2 (2), 306–333.

Harri, A., Nalley, L. L., Hudson, D., August 2009. The relationship between oil, exchange rates, and commodity prices. Journal of Agricultural and Applied Economics 41 (02), 501–510.

He, Y., Wang, S., Lai, K. K., 2010. Global economic activity and crude oil prices: A cointegration analysis. Energy Economics 32 (4), 868 – 876, policymaking Benefits and Limitations from Using Financial Methods and Modelling in Electricity Markets.

Headey, D., Fan, S., November 2008. Anatomy of a crisis: the causes and consequences of surging food prices. Agricultural Economics 39, 375–391.

Henderson, J., 2011. Building u.s. agricultural exports: one bric at a time. Federal Reserve Bank of Kansas City, Economic Review I, 63–82.

In, F., Inder, B., December 1997. Long-run relationships between world vegetable oil prices. Australian Journal of Agricultural and Resource Economics 41 (4), 455–470.

Jacks, D. S., Mar. 2013. From boom to bust: A typology of real commodity prices in the long run. CEH Discussion Papers 013, Centre for Economic History, Research School of Economics, Australian National University.

Johansen, S., June 1992. A representation of vector autoregressive processes integrated of order 2. Econometric Theory 8 (02), 188–202.

Johansen, S., October 2000. A bartlett correction factor for tests on the cointegrating relations. Econometric Theory 16 (05), 740–778.

Johansen, S., September 2002. A small sample correction for the test of cointegrating rank in the vector autoregressive model. Econometrica 70 (5), 1929–1961.

Johansen, S., Juselius, K., May 1990. Maximum likelihood estimation and inference on cointegration- with applications to the demand for money. Oxford Bulletin of Economics and Statistics 52 (2), 169–210.

Josling, T., Anderson, K., Schmitz, A., Tangermann, S., 2010. Understanding international trade in agricultural products: One hundred years of contributions by agricultural economists. American Journal of Agricultural Economics 92 (2), 424–446.

Juselius, K., 2006. The Cointegrated VAR Model: Methodology and Applications. Oxford University Press.

Kilian, L., December 2008a. The economic effects of energy price shocks. Journal of Economic Literature 46 (4), 871–909.

Kilian, L., May 2008b. Exogenous oil supply shocks: How big are they and how much do they matter for the u.s. economy? The Review of Economics and Statistics 90 (2), 216–240.

Kilian, L., June 2009. Not all oil price shocks are alike: Disentangling demand and supply shocks in the crude oil market. American Economic Review 99 (3), 1053–69.

Kilian, L., Hicks, B., 08 2013. Did unexpectedly strong economic growth cause the oil price shock of 20032008? Journal of Forecasting 32 (5), 385–394.

Kilian, L., Murphy, D. P., October 2012. Why agnostic sign restrictions are not enough: Understanding the dynamics of oil market var models. Journal of the European Economic Association 10 (5), 1166–1188.

Kilian, L., Murphy, D. P., April/May 2013. The role of inventories and speculative trading in the global market for crude oil. Journal of Applied Econometrics 29 (3), 454–478.

Kwon, D.-H., Koo, W. W., 2009. Interdependence of macro and agricultural economics: How sensitive is the relationship? American Journal of Agricultural Economics 91 (5), 1194–1200.

Leung, G. C., 2010. China's oil use, 1990-2008. Energy Policy 38 (2), 932 – 944.

Mitchell, D., Jul. 2008. A note on rising food prices. Policy Research Working Paper Series 4682, The World Bank.

Mutuc, M. E. M., Pan, S., Hudson, D., 2010. What drives commodity prices more: Oil demand or supply shocks? 2010 Annual Meeting, July 25-27, 2010, Denver, Colorado 61070, Agricultural and Applied Economics Association.

Natanelov, V., Alam, M. J., McKenzie, A. M., Van Huylenbroeck, G., September 2011. Is there co-movement of agricultural commodities futures prices and crude oil? Energy Policy 39 (9), 4971–4984.

Nazlioglu, S., Soytas, U., 2012. Oil price, agricultural commodity prices, and the dollar: A panel cointegration and causality analysis. Energy Economics 34 (4), 1098 – 1104.

Pen, Y. L., Sevi, B., Jan 2013. Futures trading and the excess comovement of commodity prices. Working Paper 1301, Aix-Marseille School of Economics, Marseille, France.

Phillip, R. C., Friederich, U. J., October 2013. Monetary policy shocks and real commodity prices. The B.E. Journal of Macroeconomics 13 (1), 35.

Roberts, M. J., Schlenker, W., October 2013. Identifying supply and demand elasticities of agricultural commodities: Implications for the us ethanol mandate. American Economic Review 103 (6), 2265–95.

Saghaian, S. H., August 2010. The impact of the oil sector on commodity prices: Correlation or causation? Journal of Agricultural and Applied Economics 42 (03), 477–485.

Sarris, A., Hallam, D. (Eds.), March 2006. Agricultural Commodity Markets And Trade: New Approaches to Analyzing Market Structure And Instability. Edward Elgar Pub, Ch. Properties of International Commodity Prices: Identifying Trends, Cycles and Shocks, pp. 16–30.

Spatafora, N., Tytell, I., Sep. 2009. Commodity terms of trade: The history of booms and busts. IMF Working Papers 09/205, International Monetary Fund.

Tang, K., Xiong, W., Sep. 2010. Index investment and financialization of commodities. NBER Working Papers 16385, National Bureau of Economic Research, Inc.

Tyner, W. E., November 2010. The integration of energy and agricultural markets. Agricultural Economics 41, 193–201.

Wright, B. D., 2011. The economics of grain price volatility. Applied Economic Perspectives and Policy 33 (1), 32–58.

Yu, T.-H. E., Bessler, D. A., Fuller, S. W., 2006. Cointegration and causality analysis of world vegetable oil and crude oil prices. 2006 Annual meeting, July 23-26, Long Beach, CA 21439, American Agricultural Economics Association (New Name 2008: Agricultural and Applied Economics Association).

Zhang, W., Law, D., Jul. 2010. What drives china's food-price inflation and how does it affect the aggregate inflation? Working Papers 1006, Hong Kong Monetary Authority.

Zhang, Z., Lohr, L., Escalante, C., Wetzstein, M., June 2009. Ethanol, corn, and soybean price relations in a volatile vehicle-fuels market. Energies 2 (2), 320–339.

Zhang, Z., Lohr, L., Escalante, C., Wetzstein, M., January 2010. Food versus fuel: What do prices tell us? Energy Policy 38 (1), 445–451.

Zilberman, D., Hochman, G., Rajagopal, D., Sexton, S., Timilsina, G., 2013. The impact of biofuels on commodity food prices: Assessment of findings. American Journal of Agricultural Economics 95 (2), 275–281.