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On the Evolving Relationship between Corn and Oil Prices

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Abstract:

The relationship between corn and oil prices is not a stable one. We identified three breaks in the relationship between corn and oil prices. The first break coincides with the second oil crisis. The second break marks the end of the agricultural export subsidy war between the EU and the US in the mid 1980s while the third one occurred at the beginning of the ethanol boom at the very end of the 1990s. The relationship between corn and oil prices tends to be stronger when oil prices are highly volatile and when agricultural policies create less distortion. The ethanol boom strengthened the relation between corn and oil prices which are (were not) cointegrated in the fourth regime (first three) regime(s). Impulse response functions confirm that corn prices systematically respond to oil price shocks, but the converse is not observed.

Keywords: Oil, Corn, Structural changes, Cointegration, Ethanol, Protectionism

Résumé:

La relation entre le prix du maïs et celui du pétrole n'est pas stable dans le temps. Trois changements structurels ont été identifiés. Le premier coïncide avec le premier choc pétrolier, le deuxième marque la fin de la guerre des subventions à l'exportation entre l'Union européenne et les États-Unis au milieu des années 80 et le troisième s'est produit au début de la phase d'expansion du marché d'éthanol à la fin des années 90. La relation tend à être de plus en plus forte en période de grande volatilité des prix du pétrole et lorsque les distorsions créées par les politiques agricoles sont plus faibles. L'expansion du marché d'éthanol a renforcé la relation entre le prix du maïs et celui du pétrole qui sont devenus (n'ont pas été) cointégrés au cours du quatrième régime (trois premiers régimes). Les fonctions de réaction aux impulsions confirment que les prix du maïs réagissent systématiquement aux chocs des prix du pétrole, mais l'inverse n'est pas vrai.

Mots clés: Pétrole, maïs, changements structurels, cointégration, éthanol, protectionnisme

Classification JEL: C32, Q11, Q17, Q40

On the evolving relationship between corn and oil prices

1. Introduction

The massive production of energy from agricultural resources during the last decade is viewed as a contributing factor behind the spectacular surge in commodity prices observed early in 2008 and in the Fall of 2010.[1] Increases in commodity prices were severe enough to trigger food security concerns in many less developed countries (FAO, 2009; Von Braun and Torero, 2009). The FAO estimates that the spike in food prices in 2008 added 115 million persons to the pool of people afflicted by chronic hunger. Similarly, the spectacular increases in the price of oil led some politicians, reporters and economists to talk about a third oil crisis.[2] The price of oil rose to \$60/bl in August 2005, reached \$92/bl in October 2007 and then hit \$147.02 /bl on the 11th of July in 2008. Whether the high prices observed in the agricultural and energy sectors are temporary or permanent is a source of contention and so are the causes for the high prices. A popular explanation for the 2008 food crisis is the expansion of the biofuel sector. In the United States, ethanol production increased by 460 % between 2000 and 2008 while the proportion of the national corn production used to produce ethanol increased from 6% to 37% during the same period (RFA, 2009).[3]

The rapid expansion of the ethanol industry has stimulated interest in the relationship between energy and agricultural commodities. Koizumi (2003) developed a dynamic partial equilibrium model to analyze the impact of the ethanol-gasoline blend ratio in Brazil on the world markets of ethanol and sugar. The maintenance of the blending ratio allows Brazil to exert much control over its domestic market and the world sugar market and to have a moderate, but persistent, impact on the world ethanol market.

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[1] Abbott, Hurt and Tyner (2009) discuss several other factors that contributed to high commodity prices, including demand from rapidly growing low income countries, the weaker US dollar and low inventories.

[2] Many commentators have been referring to a third oil crisis throughout the last decade whenever oil prices were rising. Paul Krugman was among the first to anticipate the surge in the price of oil, as documented in his April 2002 NY Times column entitled "The Third Oil Crisis?".

[3] The ethanol expansion was encouraged by the Renewable Fuel Standards (RFS) of the Energy Policy Act of 2005, which requires that gasoline sold in the U.S. should contain a minimum volume of renewable fuel (UZEPA 2006 and Zhang et al, 2009) and by a volumetric tax credit for ethanol blenders, income tax deduction for flexible fuel vehicles (FFV) and federal tax incentives initially targeting 10% ethanol-gasoline blend and extended to cover biodiesel. In addition to federal blending credits, state and federal subsidies and imports tariffs provided incentives to increase production. Koplow (2006) estimated that all of the measures amounted to a subsidy of \$1.42-\$1.87 per gallon of gasoline equivalent in 2006.

Along similar lines, Koizumi and Ohga (2007) examined the domestic and international implications of the Chinese bio-ethanol program.[4] They argue that the introduction of the E10 program was going to increase the world price of corn by 1.6%. Tyner and Farzad (2008) relied on an integrated partial equilibrium framework to analyze scenarios about the promotion of ethanol production. A fixed subsidy could induce an increase in crude oil price from $40/bbl to $120/bbl and boost ethanol production from 3.3 billion gallons to 17.3 billion gallons. This would result in a much higher corn price, higher corn production and increase the proportion of the domestic supply of corn used in ethanol production from 12% to 52%. These studies are useful because they explicitly model the linkages between agricultural commodity prices and energy prices. Hence, they can make predictions about the implications of real or hypothetical energy and agricultural policies.

Others have exploited recent advances in time series econometrics to gain new insights. Balcombe and Rapsomanikis (2008) developed a Bayesian approach to identify the nature of the cointegrating relation governing price pairs in the oil-ethanol-sugar complex. They focused on possible non linear dynamic price adjustments and found that the relationship between oil and ethanol prices is characterized by a threshold effect while that between oil and sugar prices exhibit asymmetries. Ethanol and sugar prices are linearly cointegrated and they respond to oil price shocks, but the oil price was found strongly exogenous. Rapsomanikis and Hallam (2006) found similar results by adopting the discrete two-regime threshold cointegration approach developed by Hansen and Seo (2002). Serra et al (2010) computed a smooth transition error correction model to investigate the changing price dynamics in the US corn-ethanol-oil-gasoline nexus between 1990 and 2008 on monthly data. They uncovered two cointegrating relations and found that all prices "error-correct" to deviations from at least one cointegrating vector. The link between corn and ethanol prices is particularly strong. Most of the above studies use relatively short samples. As such, they are limited in their ability to precisely identify structural changes and describe the evolution of the relationship between corn and oil prices over different regimes.[5] Furthermore, the arbitrary beginning of the sample may

[4] In October of 2004, the Chinese government introduced a compulsory 10% bio-ethanol-gasoline (E10) blend and announced several ethanol plant expansions.

[5] Serra et al. (2010) relied on a smooth structural change approach which is most especially useful when dealing with a small sample.

have an incidence on the characterization of the most recent relationship. As Andrews (2003, p.1662) puts it, a structural change "test can be used to determine the start of a sample period that is most appropriate for a given model." Structural change tests allowing for multiple endogenous breaks are also useful to make meaningful inter-temporal regime comparisons.

This paper characterizes the relationships between international corn and crude oil prices over the January 1957-April 2009. We show that the cointegration finding in other studies applies only to the recent past. The second oil shock of 1979 marks the beginning of a new regime. This event had far reaching macroeconomic implications and was often identified as a break point in many structural change investigations (e.g., Zeileis et al. 2003). The end of our second regime occurs at the end of the agricultural export subsidy war between the European Union and the United States, a year after the launch of the Uruguay Round. The level and nature of agricultural protectionism in the 1980s did much to exacerbate world price volatility as countries used policies to shield their domestic markets (Larue and Ker, 1993). As a result, the influence of the oil price on the corn price was strengthened by the second oil crisis but thwarted by policy distortions. The progress achieved in the liberalization of agricultural trade since 1995 and the energy policies encouraging the expansion of the ethanol industry have greatly strengthened the influence of the oil price on the corn price, hence the cointegration finding for the most recent regime (1999-2009). Unlike Rapsomanikis and Hallam (2006), we did not find support for the kind of non-linearities generated by a discrete two-regime threshold cointegration model, but impulse response functions confirm that oil price shocks impact strongly on corn and ethanol prices. However, the converse is not true. The implications are that corn prices will keep on being influenced by political events happening in the Middle East, unless there is a WTO meltdown and a return to protectionist policies to dampen the influence of market forces on the world corn price.[6]

The rest of the paper is organized as follows. We investigate the stochastic properties of the data and test for cointegration using the full sample in the next section. We then implement the Bai and Perron (2003) procedure to endogenously identify

[6] The comparison of regime 2 and regime 4 impulse functions is revealing. In the last (second) regime, a temporary oil price shock has a permanent (temporary) effect on the corn price.

structural break dates. We rely on these dates and their corresponding confidence intervals to uncover the events/policies that changed the relationship between corn and oil prices. A procedure that allows for multiple breaks at endogenously determined dates is warranted because events and policies are sometime anticipated or else trigger delayed responses. Section 4 reports results about the corn and oil price dynamics in the first three regimes. Section 5 focuses on the interactions between pairs of prices for corn, oil and ethanol in the most recent regime. Concluding remarks are offered in Section 6.

2. The Relationship between the Corn and Oil Prices under the Null of No Structural Change

We rely on monthly data on international crude oil and corn prices from the International Financial Statistics (IFS) of the International Monetary Fund (IMF). Our sample goes back to January of 1957 and ends in April of 2009. We begin by analyzing the time series properties of each variable. The literature is divided on this issue. Evidence of non stationnarity for corn prices is reported by Babula, Ruppel and Bessler (1995) and Newbold, Rayner and Kellard (2000) with respectively sample covering the 1978-1989 and 1900-1995 periods. Lanza, Manera and Giovanni (2005), Wlazlowski (2007) and Maslyuk and Smyth (2008) find that petroleum prices are also non stationary. Serra et al (2010) found that both corn and petroleum prices are non stationary. In contrast, Wang and Tomek (2007) argue that agricultural commodity prices should be stationary and provide empirical evidence to support their view. German and Shih (2009) contend that energy commodities prices are stationary prior to 2000 and became non stationary after.

We perform the augmented Dickey and Fuller's (1979) ADF unit root test along with the Kwiatkowski et al.'s (1992) KPSS test, which has a null of stationarity. Because these tests do not perfectly complement one another (see Maddala and Kim, 1998 p.126), we also implemented Carrion et al.'s (2001) joint confirmation analysis, which can be construed as a synthesis of both aforementioned tests[7]. We used the Bayesian Information Criterion (BIC) and the Akaike Information Criterion (AIC) to select the optimal lag length in our tests. Stationarity could not be rejected for all of the first-differenced series

[7] Charemza and Syczewska (1998) were among the first to work on confirmatory analyses. Carrion et al. stress the importance of the joint use of stationarity and non-stationarity tests to avoid to prior one null assumption over the other by setting the type I error of a test equal to the type II error of the other test.

and the confirmatory analysis validated this finding. This ruled out the possibility that the series be I(2). Table 1 reports test results for the series in levels. Both ADF and KPSS tests point toward a unit root in the oil price and corn price series and so did the confirmatory analysis. Given the limited size of our sample, the residual-based stationary boot strap procedure of Parker et al (2006) is also used. Bootstrap techniques have overwhelmingly better power than the usual asymptotic tests in finite sample. The bootstrap p-values reject the presence of unit root in both series of prices and data are considered as stationary.

Crude oil and corn prices are not cointegrated. However, the Granger causality test[8] suggests that causality is running both ways between both of them. Our full-sample analysis produced results that are quite different from the aforementioned results from the literature based on much shorter samples. However, if there were several structural changes, the full sample results could be construed as some kind of weighted average of sub-period results that would not be useful to understand the past or what is currently going on. We implement the multiple structural changes procedure of Bai and Perron (2003), BP henceforth, in the next section to ascertain the likelihood of one or more structural breaks.

3. Endogenous detection of structural breaks

There is a vast literature on structural change tests. Since the 1990s, new testing procedures have allowed for the endogenous determination of multiple breaks (eg., Andrews et al. (1996), Garcia and Perron (1996) and Liu at al. (1997) and BP (2003)). The BP procedure endogenously determines the date of each break point and generates a confidence interval around each break under rather mild assumptions. For example, errors are allowed to have different or identical distributions across regimes, to be correlated and to be stationary or non-stationary. The method consists of estimating by ordinary least squares (OLS) a linear regression with *n* breaks (*n*+ 1 regimes):

$$y_t = z_t^{'}\alpha_j + u_t;\ \ t = T_{j-1}+1,...,T_j;\ j=1,...,n+1 \qquad (1)$$

[8] The specification of the VAR was chosen with various criteria, such as the BIC, AIC, LR, HQIC and FPE and diagnostic checks on the residuals of the model. We use a lag length of 11 in the present case which is intuitive given that corn is harvested once a year.

where y_t is the dependent variable (corn price) at time *t* in the model, $z_t(m \times 1)$ is vector of regressors (the crude oil price, an intercept and seasonal dummies variables), and α_j the correspondent vector of coefficients; u_t is the disturbance at time *t*. The indices $(T_1,...,T_n)$, or the breaks points, are explicitly considered as unknown and we adopt the convention $T_0 = 0$ and $T_{n+1} = T$. Our objective is to estimate simultaneously the coefficients of the regression and the breaks points when T observations (y_t, z_t) are available.[9] The estimation procedure allows for different variances across regimes.

We also estimated restricted versions of the model where some regressors are assumed not to be regime-specific. A restricted version of model (1) can be written as:

$$y_t = x_t^{'}\beta + p_t^{'}\delta_j + u_t; \quad t = T_{j-1}+1,...,T_j;\, j = 1,...n+1 \tag{2}$$

where y_t is usually the dependent variable at time t in the model, $x_t(p \times 1)$ and $p_t(q \times 1)$ are vector of regressors, and β and δ_j the correspondent vector of coefficients. For each n-partition $(T_1,...,T_n)$ of $[T_0,...,T]$, the estimates of $\hat{\beta}$ and $\hat{\delta}_j$ are obtained by minimising the sum of squared residual:

$$\sum_{j=1}^{m+1} \sum_{T_{J-1}+1}^{T_j} \left[y_t - x_t^{'}\beta + p_t^{'}\delta_j \right] \tag{3}$$

If we denote by $\hat{\beta}(\{T_j\})$ and $\hat{\delta}(\{T_j\})$ the relative estimates of a given partition $\{T_j\}_{j=1}^{n}$ of n elements, substitute them in the objective function and denote by $S_T(T_1,...,T_n)$ the resulting sum of squared residuals (SSR), the estimated breaks dates $(\hat{T}_1,...,\hat{T}_n)$ must satisfy the condition: $(\hat{T}_1,...,\hat{T}_n) = \arg\min_{T_1,...,T_n} S_T(T_1,...,T_n)$, where the minimisation is conducted over all partitions $(T_1,...,T_n)$ under the constraint $T_j - T_{j-1} \geq d^2$ with d defined as the minimal length of a regime. Thus the break point estimator must be the global

[9] If we use the price of crude as the dependent variable and vice versa, only the δ_j coefficients will changes and the break dates will remain generally the same.

minimiser of the objective function and the parameters estimates $\hat{\beta} = \hat{\beta}(\{\hat{T}_j\})$ and $\hat{\delta} = \hat{\delta}(\{\hat{T}_j\})$ are the coefficients associated with the optimal *n*-partition.

Since break points are discrete parameters and take just a finite number of values, they can be estimated by a grid search, but this method quickly becomes computationally cumbersome when *n* exceeds 2. Fortunately, a dynamic programming-based algorithm can be used to accurately identify the break dates. It evaluates which partition achieves a global minimization of the overall SSR. If we denote by $w(i, j)$ the recursive residual at time j and $SSR(i, j)$ the SSR estimated by OLS for a segment that starts at date *i* and finish at the date *j*, then the recursive SSR for the sample would be given by $SSR(i, j) = SSR(i, j-1) + w(i, j)^2$ (Brown, Durbin and Evans,1975). Let $SSR(\{T_{r,t}\})$ be the SSR associated with the optimal partition using the first *t* observations in the case of r changes and let the minimum length between two breaks be fixed at d, then the optimal partition solves the following recursive problem:

$$SSR(\{T_{n,T}\}) = \min_{nd \le j \le T-d} [SSR(\{T_{n-1,j}\}) + SSR(j+1, T)] \qquad (4)$$

BP propose two tests about the null of absence of structural change against the alternative of an arbitrary number of changes, given an upper limit U: the $UD\max$ and $WD\max$ tests. To check for the presence of multiple structural changes, BP propose a test of the null hypothesis of no structural break against an unknown number of breaks l. The rejection of the null rationalizes the implementation of the test of the null of l breaks against the alternative of $l+1$ breaks. We can then iterate to find the endogenously determined number of breaks. We reject the null assumption of l changes in favour of the alternative of $l+1$ changes if the global SSR for the $l+1$ breaks model is sufficiently lower then that of the l breaks model. We considered a maximum number of break points of $U = 5$. We have 628 observations and setting $\varepsilon = 0.15$ as the convergence criterion for the objective function, we fix $h = \varepsilon T$ =94 months for the minimum length of any regime.

The results are presented in table 3. The UDmax and WDmax tests are highly significant and we can infer that there is at least one break point/2 regimes. Because the $SupF_T(2|1)$ statistic is 106.574 and hence highly significant and the $SupF_T(3|2)$

statistic is only 1.193, we conclude that the sequential procedure selects 2 breaks. The BIC and LWZ information criteria select 3 breaks. The simulations carried out by BP (1998) to assess the reliability of competing procedures to estimate the number of breaks indicate that the information criteria typically tends to underestimate, nor overestimate, the number of breaks. The two dates identified by the sequential procedure were close to the first and third dates suggested by the information criteria reported in table 3 (i.e., May 1977 versus January 1979 and May 2000 versus November 1999), but they were not estimated as precisely. For this reason and because we found the 1987 break plausible, we consider from this point on that there are three breaks.[10] Figure 1 illustrates the four resulting regimes.

The first break corresponds to the second oil crisis that occurred in 1979. The crude oil price almost tripled between 1978 and 1981 and this had major repercussions on the world economy. Thus, it is not surprising that the relationship binding corn and oil prices was strengthened. The second break occurred just one year after the launch of the Uruguay Round of multilateral trade negotiations in 1986. Agricultural markets were heavily distorted by protectionist policies during the 1980s.[11] It was recognized that it was time to discipline agriculture and this is why the Uruguay Round was dubbed the Agriculture Round. The EU's Common Agricultural Policy was relying extensively on variables levies and export subsidies to achieve domestic price targets. Such policies had a negative impact on the level of world prices and exacerbated world price variability (Vousden, 1992 p. 100-103, Larue and Ker, 1993). The United States was also very active through its Export Enhancement Program. Other exporters of agricultural commodities, like Canada, Australia, Argentina and New Zealand, were calling for GATT disciplines on export subsidies and because the export subsidy war was very costly to the European and US treasuries, there was far less resistance to progress in this area than on market access.

[10] The hypothesis that some regressors are not regime-specific is strongly rejected. We also tested for seasonal effects, but they turned out to be insignificant. Break dates are confirmed by Chow and CUSUM tests.

[11] OECD corn producer support estimates (in %) for the United States are available between 1986 and 2004: 44, 41, 30, 23, 19, 16, 20, 20, 17, 6, 12, 14, 28, 34, 34, 27, 20, 13, 27. The 1987-1999 period saws PSEs follow a U-shaped trend, but the high PSEs of 1999 and 2000 were substantially lower than the 1986 and 1987 PSEs.

The third break occurred in November of 1999. It coincides with the emergence of the new ethanol sector and its spectacular expansion during the last decade. The number of ethanol producing countries has more than doubled between 1993 and 2003. US production went from 1470 billion gallons in 1999 to 9000 billion gallons in 2008, due in part to the banning of Methyl Tertiary Bethyl Ether (MTBE) and its substitution by ethanol.

4. Corn-oil price dynamics between 1957 and 1999

Before comparing the price dynamics across regimes, we must characterize the stochastic properties of the data for each regime. As in the previous section ADF and KPSS tests are implemented along with the joint confirmatory analysis and, given the limited sample size of each regime, the residual-based stationary bootstrap procedure of Parker et al. (2006) is used too. The BIC and AIC information criteria are employed to select the optimal lag length and insure that residuals are white noise. The results reported in table 4 cannot reject non stationarity at the 5% level in the first and fourth regimes for both prices. Stationarity appears more plausible in the third regime for both prices and in the second regime for the corn price. Tests results for the oil price in the second regime are mixed, but the joint confirmatory analysis and the bootstrap suggest stationarity.[12] Consequently, we will consider that corn and oil prices have a unit root in the first and fourth regimes and no unit root in the intermediate regimes.

The price dynamics is analyzed with a VAR (in levels) for the regimes in which prices are stationary. Because corn and oil prices are not stationary during the first and fourth regimes, we must first ascertain whether they are cointegrated during the sub-sample periods. The results reported in Table 2 provide no support for cointegration in the first regime. However, the null of no-cointegration can be rejected at the 10% level in the fourth regime. Therefore, we rely on a VAR estimated in first difference to gain some insights about the first regime and a VEC for the last regime. Both AIC and BIC information criteria are used together with residual autocorrelation tests to select optimal lag lengths for the VARs and the VEC. The VAR analyses for the first three regimes

[12] Asymptotic tests suggest that the oil price is not stationary in the second regime, as opposed to the bootstrap procedure. We side with the latter because asymptotic tests tend to under reject unit roots (Maddala and Kim, 1998).

confirm that the corn price exhibits a strong autoregressive component. The hypothesis that the crude price does not cause the corn price is rejected by the Granger test at the 5% level in the first regime and at the 10% level in the second and third regimes. The absence of causality from the corn price to the crude price is not rejected at conventional levels in the second and third regimes, but it is rejected in the first regime.

We can now proceed with the analysis of impulse response functions (IRF).[13] We rely on the Choleski factorization procedure. The oil price is first ordered because it is more likely to affect the corn price than to be affected by it as petroleum is an important input in grain production and in the manufacturing of chemicals used in agriculture. We changed the ordering for sensitivity purposes and found similar impulse responses. The crude response to its own shocks reaches a peak in the first month and fades away after the fourth month (Fig 3.1.a). The return to zero after a few months is expected of impulse response functions derived from stationary systems. Similarly, the corn price response to an own shock is highest in the first month and vanishes after the second one (Fig 3.1.b). The corn price and crude price responses to each other's shock are very small and insignificant (Fig 3.1.c and 3.1.d). All first regime responses exhibit an oscillating pattern unlike the ones for the second regime.

In the second regime, the crude price response to an own shock reaches a peak after two months, but vanishes only after eight months (Fig 3.2.a). The corn price response to its own shock also reaches a peak after two months, decreases gradually and disappears after the ninth month (Fig 3.2.b). Both crude and corn price responses in this regime are more important in magnitude then their first regime counterparts. Crude price impulse functions to a corn innovation are not significant at any horizon (Fig 3.2.c). However, a positive response of the corn price to a crude price shock is observed. This response remains significant for almost two years (Fig 3.2.d). The adjustment in the corn price is characterized by a 3-month delay. Clearly, an oil price spike can induce a long series of corn price increases. The impulse responses for own shocks plotted for the third regime are not unlike the ones plotted for the second regime (Fig 3.3.a and 3.3.b). However, responses to a shock in the other price are not significantly different from zero,

[13] Generalised impulse response functions (GIRF) developed by Pesaran and Shin (1998) deal with the Wold-ordering problem affecting orthogonal impulse functions, but Hyeongwoo (2009) advise against the use of GIRF because they are based on extreme identifying assumptions. See Kim (2009) for more details.

even at short horizons (Fig 3.3.c and 3.3.d). In the late eighties and early nineties, energy prices remained rather stable while grains prices fluctuated moderately, but were distorted by policies. We can conclude that crude contributions to corn price volatility were very limited.

The forecast error variance decomposition analysis for corn shows that the contribution of the crude price is highest in the second regime as it explains about 30% of corn price innovations. Crude price contributions are particularly low in the first and third regimes (12% and 2% respectively). The implication is that price transmission from the oil price to the corn price is strongest when oil prices are highly volatile. Table 5 shows the variances in the corn and oil prices across regimes as well as the contribution of the oil price to the variance of the corn price. It is in the fourth regime that the oil price is most volatile and not surprisingly corn and oil prices are cointegrated only in this regime.

5. The oil-ethanol-corn price linkages after 1999

The fourth regime coincides with the ethanol boom and this is why we enlarge the scope of our analysis to include ethanol prices. Since data on international ethanol prices are not available, we use the US ethanol price as a proxy because of the US leading position in the world market for ethanol.[14] We rely on three bivariate models as opposed to a 3-variable model because the bivariate approach treats each pair of prices as having their own adjustment to equilibrium and isolates their mutual interactions; this flexibility is not possible in a simultaneous analysis of the three prices (Balcombe, 2008). Using a bivariate approach also insures consistency with the approach adopted in our analysis of the first three regimes. Of particular interest is the nature of the cointegration relationships between pairs of prices, especially the ones involving the ethanol price. Interest in the analysis of non-linear long run relationships was spurred by the contributions of Tsay (1989), Balke and Fomby (1997), Hansen and Seo (2002), Seo (2004), Choi and Saikonnen (2004) and Kapetanios, Shin and Snell (2006). Non-

[14] The US also dominates with the largest share of world production (i.e., US production accounted for 50% of world total production in 2008 and is expected to continue to grow).

linearities have also been investigated in several agriculture studies (e.g., Goodwin and Piggott (2001) and Balcombe, Bailey and Brooks (2007)).[15] .

Table 4 shows that the ethanol price is I(1) while table 6 confirms the presence of a linear cointegration relation in all three price pairs. Non linear price adjustments between oil and sugar in Brazil were reported by Rapsomanikis and Hallam (2006) and Balcombe and Raspomanikis (2008). Serra et al. (2010) also found non-linearities in the relationship linking oil and corn prices in the US (Serra and al., 2010). Accordingly, it seems pertinent to entertain the possibility of non-linear cointegration relations. It is worth reiterating that the aforementioned studies pertain to domestic price adjustments, that they use different methodologies on different periods. Serra et al (2010) argue that the smooth transition specification is most suitable when adjustment costs differ among individuals. Balcombe and Raspomanikis (2008) favour abrupt regime changes over a smooth transition in modeling the sugar-ethanol-oil nexus because of the manner with which macroeconomic variables react to oil price shocks. Based on this argument, we adopt a standard threshold approach featuring discontinuous adjustments. The model can be specified as follows:

$$\Delta_{Pt} = A' P_{t-1}(\beta) + u_t \tag{5}$$

where P_t is a n dimensional I(1) time series, $w_{t-1}(\beta) = \beta' p_t$ is the I(0) error correction term $P_{t-1}(\beta) = (1 \;\; w_{t-1}(\beta) \;\; \mathrm{p}_{t-1} \;\; ... \;\; \Delta p_{t-l})'$ is a $k \times 1$ regressor, u_t a vector martingale difference sequence with a finite covariance matrix $\Sigma = E(u_t u_t')$ and l is the optimal lag. In our study $n = 2$, the optimal lag selected by information criterion is $l = 2$, hence $k = 4$. The parameters (β, A, Σ) are estimated by maximum of likelihood and the iid Gaussian errors hypothesis is assumed. Some normalization on β are needed to achieve identification. In a bivariate context, this can be achieved by setting one element of β to unity.

Thus a two-regime threshold cointegration model can be written as follows:

$$\Delta_{Pt} = \begin{cases} A_1' P_{t-1}(\beta) + u_t & if \;\; w_{t-1}(\beta) \le \gamma \text{ below the threshold} \\ A_2' P_{t-1}(\beta) + u_t & if \;\; w_{t-1}(\beta) \ge \gamma \text{ above the threshold} \end{cases} \tag{6}$$

[15] For an example of an agricultural application on stationary data, see Bonroy, Gervais and Larue (2007).

The threshold (asymmetric) hypothesis refers to the fact that adjustment mechanisms toward the long run relationship differ, in a discrete manner, with respect to the magnitude (the sign) of the deviation. To test for linearity against the threshold cointegration hypothesis, Hanson and Seo (2002) developed sup-like statistics. Two tests can be used. The first is based on a fixed-regressor bootstrap and the other relies on a residual bootstrap. The value of the $SupLM^0$ test statistics in conjunction with the fixed-regressor and residual bootstraps p-values reported in table 6 cannot reject the linear cointegration hypothesis for all three pairs of prices. We conclude that for our variable definition and sample period, a linear VECM is more appropriate than a threshold model to analyze price linkages between the oil, corn and ethanol prices.[16]

The linear VECM reported in tables 6 are characterized by strong autoregressive behaviour. This is especially true for the corn price. This is consistent with our theoretical prior because crude oil production is far more elastic in the short run than corn production. Corn is harvested only once a year and as such its supply is very inelastic within a year which implies that demand shocks can induce much price volatility. Transitory effects between prices are significant, so the presence of short run effects in the VEC specifications are important. In the corn-oil pair of prices, the error-correcting coefficient is significant at the 5% level in the corn equation, but not in the crude equation. Put differently, corn prices adjust to deviations in the cointegration relation binding corn and oil prices, but the oil price does not. The oil price is weakly exogeneous and all of the adjustment needed to restore the long run equilibrium between corn and oil prices is done by the corn price.

For the corn-ethanol price pair, the error-correction parameter is significantly different from zero only in the ethanol price equation. The corn price is weakly exogenous and all of the so-called long run adjustments fall on the ethanol price. The estimated cointegrating parameter is 0.61 which means that 60 % of a shock on the long run relation between the corn price and the ethanol price is transmitted to next month's ethanol price.

[16] We have also tested the linear hypothesis against a restricted version of the threshold model, where the threshold parameter is set to zero, but still we could not reject the linear hypothesis.

In the oil-ethanol pair, the statistical significance of the error-correction parameters suggests that the oil price is at least weakly exogenous. Additional information from Granger causality tests reveals that the oil price is actually strongly exogenous, in both the corn-ethanol and the oil-ethanol bivariate analyses. The highest speed of adjustment is observed in the oil-ethanol bivariate model, as the ethanol price adjusts to long run deviations triggered by oil price shocks twice as fast as it adjusts to corn price shocks.

The impulse response functions (IRF) provide a visual assessment of how variables adjust to one-time shocks. Unlike in a VAR, IRF derived from a VEC are not expected to necessarily converge to zero after a few periods. Granger causality tests can be used to deal with the order of the variable that condition the impulse responses. The oil price is a logical candidate to be the first variable in the VECs in which it is paired with the corn price and the ethanol price. By the same token, the corn price is the first variable in corn-ethanol VEC. Figure 3.4a shows that the response of the oil price to an own shock quickly converges, but not to zero. This is a case where a temporary shock has a permanent effect. The corn price response to an own shock reaches a peak after four months and ceases to be statistically significant after nineteen months (see Figure 3.4.b). The corn price reverts to its mean in less then two years and its shocks are said to be transitory. The crude price response to a corn price shock is very weak (see Fig 3.4.c), which stems from the previous finding about the crude oil price not error-correcting. In contrast, a one-time increase in the crude price triggers a permanent positive effect on the corn price (Figure 3.4.d), but the adjustment is relatively slow. The corn price IRF reveals that six months are needed to observe a statistically significant price effect. An oil price innovation permanently raises the corn price by about 7% over its pre-shocks level. These results confirm the strong influence of the energy sector on agriculture.

Figure (3.4.e) shows that an increase in the corn price causes ethanol price increases. Ethanol price responses increase during the first four months and converge to a statistically positive value, 5%, after fifteen months. The corn shock has a permanent effect. It is also transmitted quickly as statistically significant IRF are observed two months after the shock. This sort of response is expected to be observed as long as corn will remain a crucial input in ethanol manufacturing (OCDE, 2006). The response of the ethanol price to a crude price shock is positive and immediate, as the IRFs are statistically significant from the first month. The IRFs peak after four months and converge to 6.3 % after ten months (Figure 3.4.f). The impact of an oil shock on the ethanol price is permanent and has the expected sign: a higher oil price leads to a higher ethanol price because ethanol becomes more competitive and the blending demand increases. The forecast error variance decomposition for the ethanol price estimates the contribution of the oil price to the variance of the ethanol price at 89 %. The crude response to an ethanol shock is not significant at any horizon (Figure 3.4.g) which was expected given strong exogeneity of the oil price.

The response of the corn price to an ethanol shock has the expected sign (Fig 3.4.h). A positive ethanol shock should trigger an increase in the aggregate demand for corn which should translate into higher corn prices. However, the lower bound of the confidence interval around the IRF lays below zero which imply that what could be construed at first glance as a permanent positive effect is actually not significant. Accordingly, ethanol contributions to corn innovations are small (below 9%). Balcombe and Raspomanikis (2008) reached a similar conclusion in their Brazilian study about the ethanol-sugar linkage. This result might be robust at this point in time, but one would expect that future growth in corn-based ethanol production should make it easier to find a significant feedback effect from ethanol price shocks.[17]

In short, the linkages between the energy and agricultural sectors are much stronger now than in the past as shown by the absence of cointegration relations prior to November of 1999. Our results clearly indicate that oil price shocks have very strong and long lasting effects on the prices of corn and ethanol. This is likely to continue in the foreseeable future. The fact that ethanol and corn prices are strongly impacted by the oil price does not necessarily translate into equally strong linkages between corn and ethanol prices.

[17] Production is expected to triple between 2012 and 2022 according to the Energy Independence and Security Act (EISA, 2007).

6. Conclusion

The relationship between corn and oil prices has attracted much attention lately because of the spectacular increases in commodity prices observed in the first part of 2008 and in the second part of 2010. Oil is an important direct input in the production of agricultural outputs, but also in the manufacturing of key agricultural inputs. Thus, one would expect that a relationship exists between corn and oil prices since the mechanization of agriculture. However, the recent ethanol boom is diverting an increasing portion of corn production from food and feed end uses. This suggests that the relationship between corn and oil prices has experienced at least one structural change. We argue that it is crucial to identify all of the structural breaks. First, it is important to understand the repercussions of past events and/or policies. Second, even if our interest is specific to the most recent regime, a more precise estimation of the parameters will result if one can rely on the endogenous identification of multiple breaks to determine the beginning of the most recent subsample than if an arbitrary date is chosen by researchers.

We used structural change tests proposed by Bai and Perron (1998, 2003) to determine the number of breaks in the relationship between the corn and oil prices. Using monthly data going back to January of 1957, we found three breaks. The first one corresponds to the second oil crisis in the late 1970s. This event had major macroeconomic repercussions and obviously had a strong incidence on agriculture. Our second break occurred at the end of 1987. The 1979-1987 period was one during which grain prices were heavily distorted by domestic and export subsidies. The beginning of the Uruguay Round of multilateral trade negotiations put an end to the export subsidy war between the European Union and the United States. The evidence from the impulse response functions for the third regime indicates that the oil-corn linkage was at its weakest, thus reflecting both the high but declining market distortions created by protectionist agricultural policies and the relative stability of the oil price during this period. The last break marks the beginning of the ethanol boom in the United States in 1999.

We could not find support for cointegration between corn and oil prices in the first three regimes. Cointegration is supported in the fourth regime, but we could not reject linear cointegration when tested against threshold cointegration. Apart from this, our results are quite close to the results of Balcombe and Raspomanikis (2008) that pertain to the oil-sugar-ethanol nexus in Brazil. We find from our bivariate models that the oil price is either weakly or strongly exogenous as it does not error-corrects and in some cases is not Granger caused by the corn price or the ethanol price. However, we found that the corn price and the ethanol price systematically react to oil price shocks. Thus, disturbances in the Middle East impacting on the oil price should continue to have a strong impact on the world price of corn, but less so if major corn producing countries were to revert back to their old trade distorting welfare decreasing policies.

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Table 1. Stochastic properties of prices over the whole sample

	ADF test		KPSS test	Joint confirmation of a unit root	Bootstrap p-value
	Selected lag	Statistic			
Corn	2	-2.752	1.070	Yes	0.000
Oil	6	-2.103	1.320	Yes	0.000

Critical values for the ADF test were obtained from Davidson and Mackinnon (1993) and the KPSS critical values were obtained from Kwiatkowski *et al.* (1992). The critical values for the joint hypothesis of a unit root were taken from Carrion-i-Silvestre *et al.* (2001). A time trend was present in all the tests.

Table 2. The Johansen likelihood cointegration test

Period	Number of cointegrating relations	Trace Statistic
Corn-oil (1957:1 / 1978:11)	0 versus 1	3.946
	1 versus 2	0.935
Corn-oil (1999:11 / 2009:4)	0 versus 1	16.903*
	1 versus 2	6.062
Oil-Ethanol (1999:11 / 2009:4)	0 versus 1	21.830**
	1 versus 2	7.605
Corn-Ethanol (1999:11 / 2009:4)	0 versus 1	20.069**
	1 versus 2	7.740

* and ** denote significance at the 10% and 5% levels respectively.

Table 3. The BP procedure

The double maximum tests			
	Computed values		Critical value at the 5.00 % level
UDmax	106.857		8.880
WDmax	153.832		9.910
The $SupF_T(k+1\|k)$ statistics			
$SupF_T(k+1\|k)$	SupF tests	Critical values at the 5.00 % level	Critical values at the 10.00 % level
$SupF_T(2\|1)$	106.574	8.580	7.040
$SupF_T(3\|2)$	1.193	10.130	8.510
$SupF_T(4\|3)$	0.027	11.140	9.410
$SupF_T(5\|4)$	0.000	11.830	10.040
Break dates and their confidence intervals			
Changes	Date	Confidence interval at the 10% level	
1	January 1979	June 1977	March 1981
2	December 1987	January 1983	January 1990
3	November 1999	June 1999	April 2003

Table 4. Tests about the stochastic properties of the data within each regime

Regime		ADF test		KPSS test	Joint confirmation of a unit root	Bootstrap p-value
		Selected lag	Statistic			
Regime 1	Corn	4	-2.274	0.638	Yes	0.787
	Oil	1	-1.540	1.040	Yes	0.969
Regime 2	Corn	2	-1.562	0.439	No	0.000
	Oil	3	-4.077	0.194	No	0.000
Regime 3	Corn	2	-2.678	0.283	No	0.000
	Oil	2	-3.603	0.326	No	0.000
Regime 4	Corn	1	-2.099	0.319	Yes	0.696
	Oil	2	-2.726	0.198	Yes	0.646
	Ethanol	3	-2.817	0.173	Yes	0.190

Critical values for the ADF test were obtained from Davidson and Mackinnon (1993) and the KPSS critical values were obtained from Kwiatkowski *et al.* (1992). The critical values for the joint hypothesis of a unit root were taken from Carrion-i-Silvestre *et al.* (2001).

All variables are time trended in the last regime, corn and oil are also time trended in the first regime, but not in the both intermediary regimes. The null hypothesis of the bootstrap is the unit root.

Table 5. Corn and oil price variance in different regimes

Regime	Regime 1	Regime 2	Regime 3	Regime 4
Corn variance	731.801	548.486	484.294	2004.237
Oil variance	17.802	53.645	13.628	668.375
Oil contribution to corn variance	2 %	30 %	12 %	78 %

Table 6. Estimation of a the linear VECM for the three pair of prices

Regressors	Corn-oil		Corn-ethanol		Oil-ethanol	
	Corn	Oil	Ethanol	Corn	Ethanol	Oil
EC_{t-1}	- 0.061** (0.027)	0.003 (0.041)	- 0.121*** (0.039)	0.025 (0.023)	- 0.262*** (0.058)	0.062 (0.050)
ΔP_{Corn}	0.291*** (0.093)	- 0.332** (0.137)	0.341** (0.152)	0.275*** (0.924)	–	–
ΔP_{Oil}	0.063 (0.067)	0.248*** (0.093)	–	–	0.195* (0.107)	0.259** (0.104)
ΔP_{Eth}	–	–	0.0008 (0.009)	-0.004 (0.005)	0.290*** (0.090)	-0.030 (0.087)
Cons	0.002 (2.234)	-0.003 (0.008)	0.001 (0.009)	-0.004 (0.005)	0.001 (2.234)	-0.004 (0.008)
Cointegrating vector	(1, - 0.648)		(1, - 0.616)		(1, 0.517)	
Test statistic value	16.903		10.223		14.1688	
Fixed regressor p-value	0.108		0.853		0.320	
Residual-bootstrap p-value	0.050		0.353		0.234	

*, **and *** denote significance at 10%, 5% and 1% levels, respectively. Standard errors are between parentheses.

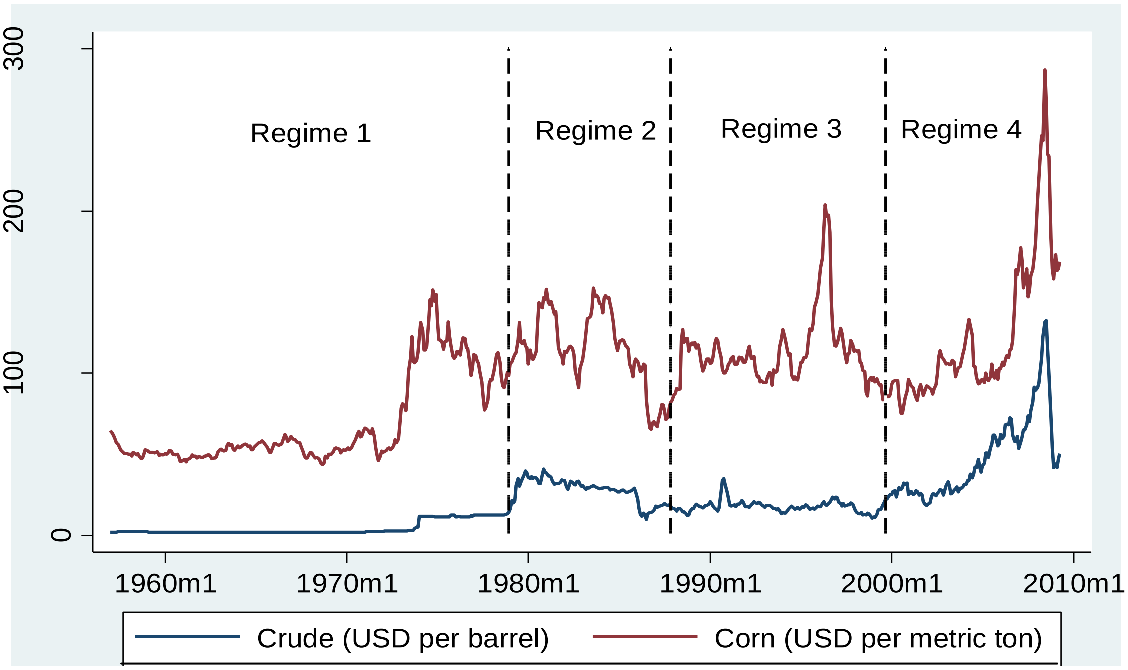


Figure 1. Corn and oil price evolution over the whole sample and identified regimes

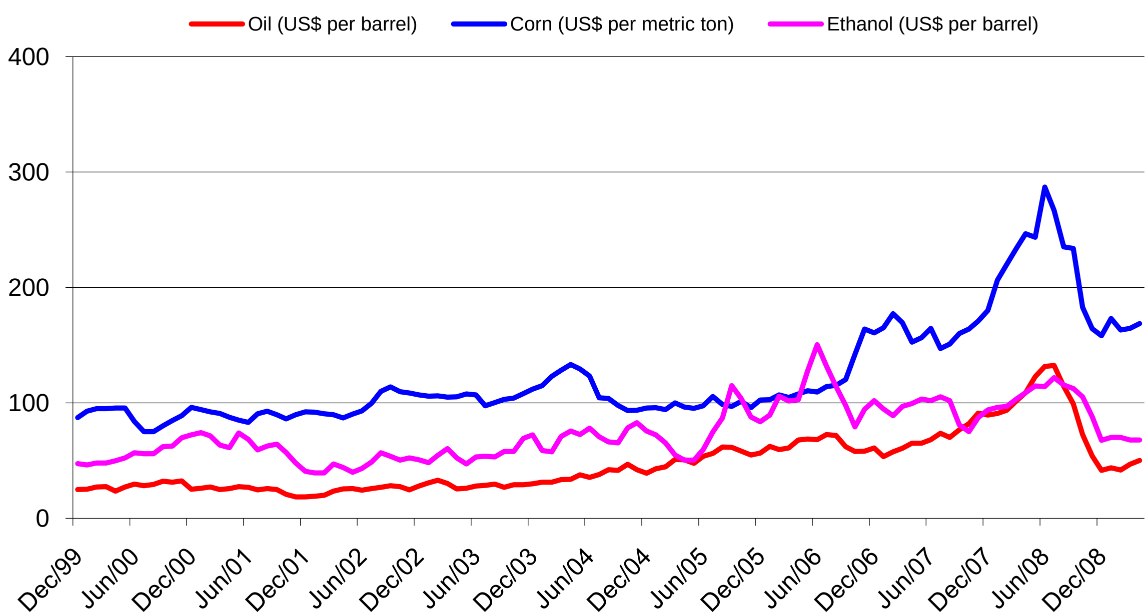


Figure. 2 Oil, ethanol and corn prices evolution over the period October 1999- April 2009

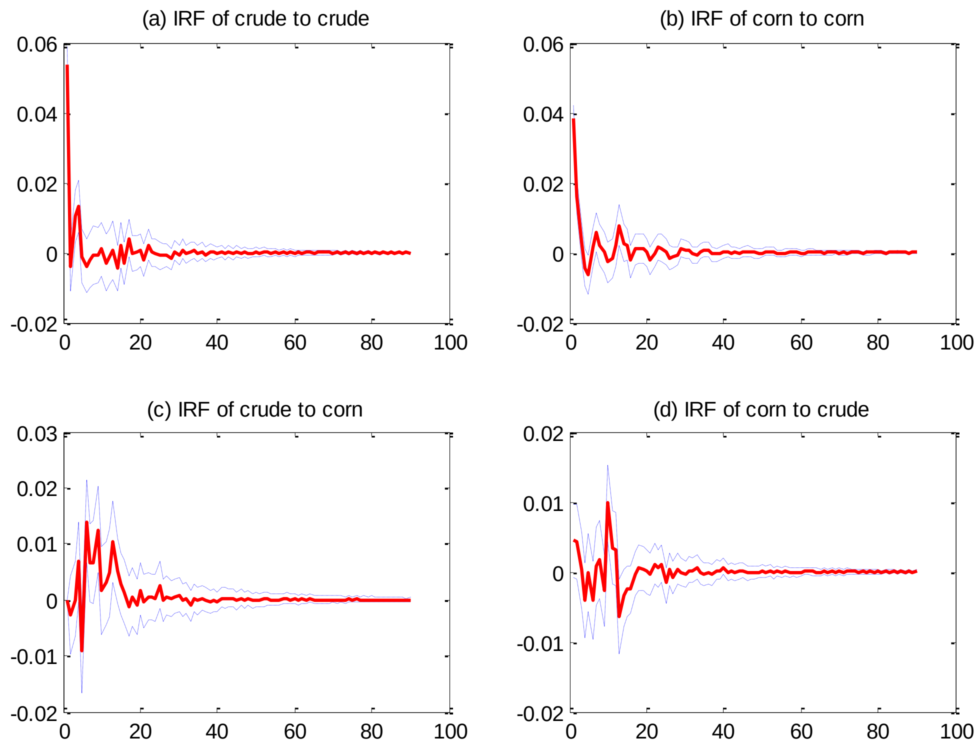


Figure 3.1 Impulse response functions for the first regime

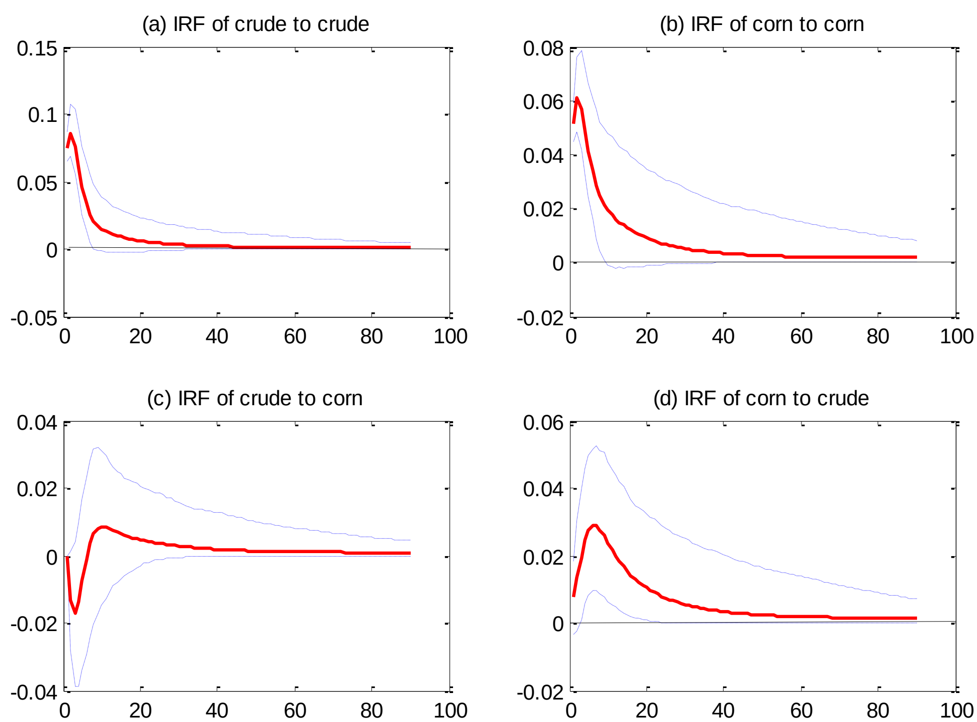


Figure 3.2 Impulse response functions for the second regime

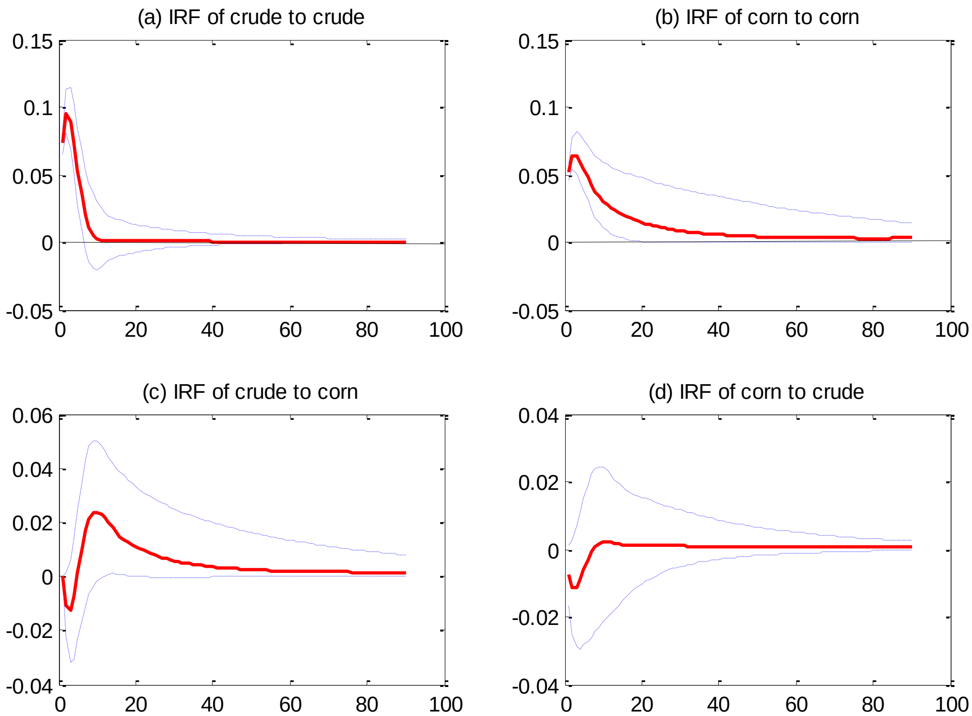
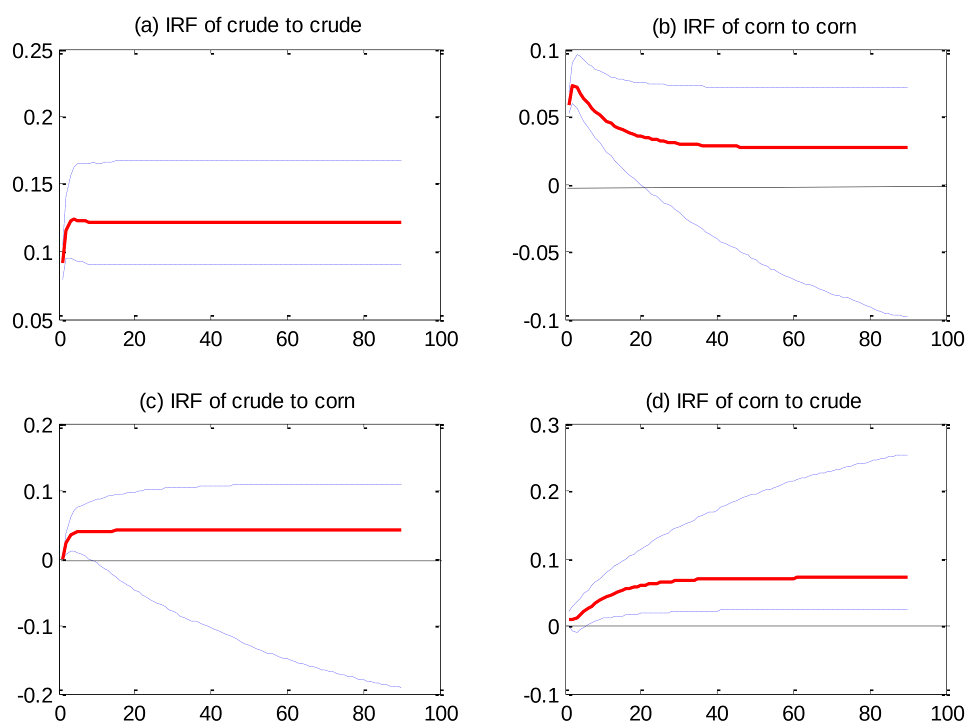


Figure 3.3 Impulse responses functions for the third regime



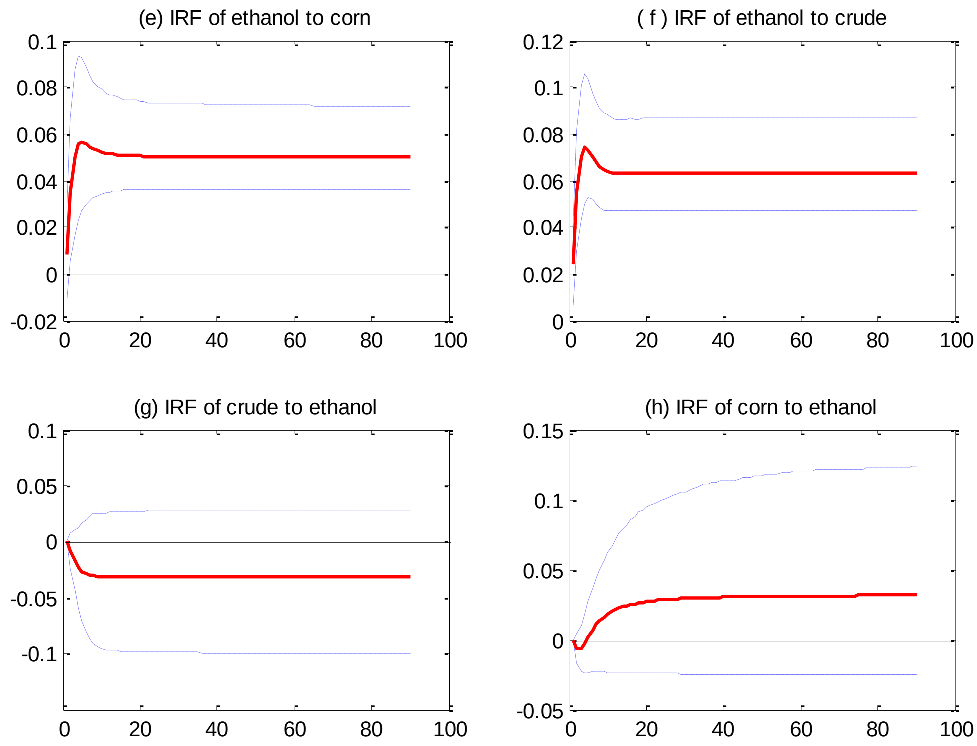


Figure 3.4. Impulse response functions for the recent regime