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Threshold cointegration in the sugar-ethanol-oil price system in Brazil: evidence from nonlinear vector error correction models

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ABSTRACT

In this paper, the possibility of nonlinear dynamic adjustment in the sugar-ethanol-oil nexus in Brazil is examined. Threshold vector error correction models are employed to test for linearity in the adjustment of prices of sugar and oil, ethanol and oil and ethanol and sugar. Strong evidence of threshold type nonlinearity is found. The results suggest that sugar and oil and ethanol and oil prices are characterised by discrete threshold behaviour, whereas sugar and ethanol can be thought of as being linearly cointegrated. Threshold estimates suggest that sugar prices adjust rapidly to a long run equilibrium, determined by oil prices, in an asymmetric manner, when disequilibria are negative. The dynamic adjustment of ethanol prices is faster when the oil-ethanol price spread widens and ethanol prices are below a critical threshold. Both sugar and ethanol prices are found to be determined by oil prices and no evidence for a causal relationship that runs from oil to ethanol to sugar is found.

JEL Classification: C12, C22, Q11.

Keywords: Cointegration, Threshold, Oil, Ethanol, Sugar, Prices

RÉSUMÉ

Les auteurs étudient la possibilité d'un ajustement non-linéaire dynamique du rapport sucre-éthanol-pétrole au Brésil. Des modèles de correction d'erreurs de vecteurs de seuil sont utilisés pour tester la linéarité de l'ajustement des prix du sucre et du pétrole, de l'éthanol et du pétrole, et de l'éthanol et du sucre. Des preuves solides de non-linéarité de seuil ont été observées. Les résultats indiquent que les prix du sucre et du pétrole et ceux de l'éthanol et du pétrole se caractérisent par un léger comportement de seuil, alors qu'on peut considérer que ceux du sucre et de l'éthanol présentent une cointégration linéaire. Selon les estimations de seuil, les prix du sucre s'ajustent rapidement à un équilibre à long terme, déterminé par les prix du pétrole, de façon asymétrique lorsque les déséquilibres sont négatifs. L'ajustement dynamique des prix de l'éthanol est plus rapide lorsque l'écart entre les prix du pétrole et de l'éthanol s'élargit et que les prix de l'éthanol sont inférieurs à un seuil critique. Il apparaît que les prix du sucre et de l'éthanol sont tous deux déterminés par les prix du pétrole et aucune preuve d'un rapport de cause à effet n'a été observée entre le pétrole et l'éthanol et le sucre.

RESUMEN

En este documento, se analiza la posibilidad de un ajuste dinámico no lineal en el nexo azúcar-etanol-petróleo en Brasil. Se emplean modelos de corrección de errores de vectores de umbral para probar la linealidad del ajuste de precios del azúcar y petróleo, etanol y petróleo y etanol y azúcar. Se hallaron marcados indicios de no linealidad del tipo de umbral. Los resultados sugieren que los precios del azúcar y el petróleo y del etanol y el petróleo se caracterizan por un discreto comportamiento de umbral, mientras que se puede considerar que el azúcar y el etanol están cointegrados linealmente. Las estimaciones de umbral sugieren que los precios del azúcar se ajustan rápidamente a un equilibro de largo plazo, determinado por los precios del petróleo, de una manera asimétrica, cuando los desequilibrios son negativos. El ajuste dinámico de los precios del etanol es más rápido cuando se amplía el margen de los precios del petróleo y el etanol y los precios del etanol se encuentran por debajo de un umbral crítico. Se descubrió que los precios del petróleo determinan tanto los precios del azúcar como del etanol y no existen indicios de que exista una relación causal que vaya del petróleo al etanol y del etanol al azúcar.

CONTENTS

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1 INTRODUCTION

Nonlinear vector error correction models (VECMs), since their introduction by Balke and Forby (1997), became increasingly popular in empirical applications that often focus, in agricultural economics and in commodity analysis, on testing the Law of One Price and assessing the nature and the extent of market integration (Abdulai, 2000; Fackler and Goodwin, 2001; Goodwin and Piggott, 2001; Barrett and Li, 2002; Sephton, 2003, Balcombe, forthcoming). In mainstream economics, nonlinear VECMs examine the term structure of interest rates and asset pricing. The popularity of nonlinear VECMs has led to contributions that developed an analogue to Granger's representation theorem for nonlinear vector autoregressions (see Corradi *et al.*, 2000; Escribano and Mira, 2002; Bec and Rahbec, 2004; Saikkonen, 2005).

Threshold behaviour and discrete adjustment characterise many economic relationships that determine commodity and asset prices, inventories, interest and exchange rates and employment. Thresholds are normally thought of as functions of transaction and adjustment costs, or economic risk that prevent agents from adjusting continuously to changes in markets, as reflected by the empirical notion of cointegration and the related linear VECMs. Threshold cointegration, as considered by Balke and Forby (1997) includes such discrete adjustment to long run equilibrium. In this model, the cointegrating relation between two variables is inactive within a certain threshold, resulting in the variables not adjusting to deviations from the equilibrium, with adjustment taking place only when deviations become large and exceed the threshold.

There is a number of important issues related to nonlinear VECMs. These include testing for the null of linear cointegration against the alternative of threshold cointegration, the identification and estimation of threshold parameters that govern regime switching, as well as that of the model's slope parameters and the provision of the corresponding standard errors. These issues have been addressed in differing ways by a growing literature that encompass discrete and smooth threshold VECMs and Markov-chain VECMs, classical and Bayesian estimation and inference.[1]

In this paper, the discrete two regime threshold cointegration approach is adopted, utilising an algorithm developed by Hansen and Seo (2002) to test for threshold behaviour in the oil-ethanol-sugar price system in Brazil. The objective of this paper is to contribute in the current debate on the behaviour of commodity prices, and especially sugar prices, in the wake of the oil price hike. Prices of oil reached historically high levels in 2005-2006, resulting in increases in the Brazilian price of ethanol and sugar. Casual observation and cost-accounting exercises suggest the possibility of nonlinear behaviour in this system with sugar and ethanol prices adjusting to oil prices in a discrete manner when disequilibria exceed a critical threshold.

The paper is organised as follows. Section 2 provides a description of the relationship between the ethanol and sugar markets in Brazil and the manner this has been shaped by government intervention and technical change. Threshold behaviour and the estimation of TVECMs are discussed in section 3. Section 4 presents the empirical application and results, whilst section 5 presents the concluding remarks.

2 THE SUGAR-ETHANOL-OIL MARKETS

Demand for biofuels, such as methanol, ethanol and biodiesel, has been growing worldwide mainly due to environmental reasons. Biofuels, being derived from cellulosic biomass, emit lower carbon dioxide levels, as compared to fossil fuels and, for many Kyoto Protocol signatory countries, consist of the main option for mitigating greenhouse emissions. Ethanol, is currently the most widely used

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[1] Discrete and smooth threshold behaviour has been examined by Hansen and Seo (2002) and Kapetanions *et al.* (2006) respectively. For Bayesian estimation and inference of threshold VECMs see Balcombe (2006)

biofuel. Its production is based on commodities that are traditionally produced for food, such as sugarcane in Brazil, maize in the United States and wheat in the European Union.

Brazil is the most important producer and consumer of ethanol in the world. In 2005-2006, approximately half of Brazil's sugarcane production was processes into ethanol, sufficing for approximately one fifth of the country's total transport fuel energy consumption. The growth of the Brazilian ethanol market has been realised due to a combination of factors, including government policies and technical change both in the processing of sugarcane into ethanol and in the manufacturing of vehicles that can use high level blends of ethanol with petrol. The national alcohol programme began in 1975 with the aim of reducing the country's oil import bill. The programme consisted by a number of different policy instruments that included production quotas and institutional setting of the price for ethanol at a level lower than that of petrol, combined with subsidies to ethanol distillers. The ethanol programme was effectively eliminated in the 1990's and a transition to full liberalisation took place between 1996 and 2000. Although, the government no longer exercises direct control over ethanol production and exports, it sets an official blending ratio of anhydrous ethanol with petrol to 20-25 per cent and periodically provides support in the form of purchases and sales from ethanol strategic reserves (International Energy Agency, 2004).

In most countries, production costs of ethanol are higher relative to the costs of extraction of fossil fuels and government policies are essential in providing incentives to the industry in order to produce adequate quantities of biofuels (Goldemberg *et al.,* 2004; OECD, 2006). However, the Brazilian alcohol programme assisted the industry to achieve economies of scale which, in conjunction with technology improvements, have resulted in Brazil being the only country where ethanol production costs are lower than the regional supply costs of petrol in volume terms. In energy terms, when prices per litre are converted into prices per energy unit, ethanol production costs are, in general, higher that regional supply costs for petrol. However, substantial increases in the price of oil not only during the period 2005-2006, but also for a number of months in the years 2002 and 2003 have resulted in increasing ethanol's competitiveness (International Energy Agency, 2004; Goldemberger *et al.*, 2004; Hamelick and Faaj, forthcoming). In a similar manner, cost-accounting exercises suggest that ethanol is competitive with petrol over a price level of US$39 per barrel (Organisation for Economic Development and Cooperation, 2006). With crude oil prices reaching historically high levels of US$60 per barrel since 2005 onwards, the demand for ethanol has strengthened considerably, with the price of ethanol in Brazil increasing by over 60 percent from May 2005 to the same month in 2006 (Food and Agriculture Organisation, 2006). Strong derived demand for sugarcane in Brazil, the world's leading sugar producer and exporter, affected the world sugar market balance, thus fuelling world sugar prices that, during the first half of 2006, reached a level almost twice as high relative to the same period in 2005.

Both formal cost-accounting and casual observations in market developments suggest that the oil price and the prices of sugar and ethanol in Brazil may co-move outside a certain threshold, that is defined over a certain price spread and mainly determined by adjustment costs in the sugar-ethanol processing industry, as well as by technical factors, such as the extent to which petrol can be substituted for ethanol. The introduction of *flex-fuel* vehicles that can use any combination of petrol-ethanol blends, but also pure ethanol only, has enhanced considerably the substitution possibilities between these fuels and the demand prospects for ethanol. The sale of flex-fuel cars, assisted by some tax incentives, has increased dramatically since their introduction in 2003. Flex-fuel cars comprised 40 percent of total car sales in Brazil in 2005, with forecasts for 2006 suggesting that their share in total car sales will increase to over 60 percent of total car sales, allowing petrol and ethanol to be purchased on the basis of relative prices (Goldemberger *et al.,* 2004; Tokgoz and Elobeid, 2006). At the sugar mill – sugar distillery level, the decision to utilise sugarcane for the production of ethanol or sugar is also made on the basis of relative prices, as the industry consists of a large number of firms with a dual-processing structure that can that can switch easily between the production of sugar and ethanol. The surge in prices, in conjunction with the continuously increasing substitution possibilities between ethanol and oil, provides good economic reasoning for the existence of threshold effects in the oil-ethanol-sugar price complex in Brazil.

3 THRESHOLD VECTOR ERROR CORRECTION MODELS

Threshold vector error correction models (TVECMs) have been initially applied by Blake and Forby (1997) and Hansen and Seo (2002), having their origin in the Self-Exciting Autoregressive model (SETAR) by Tong (1978, 1983). In TVCEMs, threshold effects follow variants of the specification of Balke and Forby (1997) where the rate of adjustment to the long run equilibrium of two variables $y_t' = \left(y_{1,t}, y_{2,t}\right)$ differs between regimes as follows:

$$\begin{aligned}
&\Delta y_t = \mu_t + F\left(e_{t-1}\right) I + \sum_{i=1}^{k} \gamma_k \Delta y_{t-k} + u_t \\
&F\left(e_{t-1}\right) = \delta \alpha_1 e_{t-1} + \left(1-\delta\right) \alpha_2 e_{t-1} \qquad (1) \\
&\delta \begin{cases} = 1 \text{ if } e_{t-1} \leq \lambda_1 \\ = 0 \text{ if } e_{t-1} > \lambda_2 \end{cases}
\end{aligned}$$

where t=1,...n, and $F\left(\square\right)$ is an indicator function of the error correction $e_{t\text{-}1}$ which is assumed to be covariance stationary with zero mean. Vectors $\alpha_i' = \left(\alpha_{1,i}, \alpha_{2,i}\right)$ and $\gamma_i' = \left(\gamma_{1,i}, \ldots \gamma_{k,i}\right)$ are adjustment and short run dynamic parameters. As in linear VECMs, the error correction term is defined as $e_{t-1} = y_{1,t-1} - \beta_1 y_{2,t-1} - \beta_0 - \beta_2 t$, whilst the parameters β_0 and β_2 may assume values equal to zero. The TVECM errors u_t is assumed to be an *iid* Gaussian sequence with a finite covariance matrix, $\Sigma = E\left\{u_t u_t'\right\}$. Models such as (1) with $\lambda_1 \leq \lambda_2$ allow for a Band TVECM specification, where there is no adjustment inside the threshold, when the error abide by $\lambda_1 < e_{t-1} \leq \lambda_2$, whilst a simple two-regime threshold model is specified with $\lambda_1 = \lambda_2$. Other popular threshold models, such as the Momentum Threshold Autoregressive models of Granger and Lee (1999), Enders and Granger (1998) and Escribano and Pfann (1997) allow for the variables to adjust differently depending on whether the disequilibria are negative or positive.[2]

Balke and Forby (1997) view cointegration as a global characteristic of the series, whilst threshold behaviour consists of a local characteristic and, therefore, conduct estimation in two steps. The first step comprises of testing for non cointegration and the estimation of the cointegrating vector, by means of the Engle-Granger method. As a second step, tests for nonlinearity in the residuals of the cointegration relation are performed by means of cumulative least squares tests, as well as by the threshold autoregression tests developed by Tsay (1989). In theory, such a two-step estimation procedure is appropriate, as in the event that the residuals follow a univariate SETAR process, the threshold of this model should be identical with the threshold in the cointegrating relationship. In practice, however, such an estimation procedure may not be optimal, if threshold behaviour was present, since the likelihood function, on which the estimates of the cointegrating vector are based, depends on the threshold parameters. In addition, Balke and Forby (1997), on the basis of simulation exercises, note that non parametric tests, such as that of Tsay (1989) have, in general low power as compared to model based tests.

Nevertheless, the estimation of bivariate threshold models is quite complex. In general, the inequality constraints implied by the threshold behaviour are difficult to enforce and lead to statistical problems in both estimation and inference. Firstly, Maximum Likelihood Estimation (MLE) is complicated, as the likelihood function is jagged and not differentiable, rendering optimisation methods that are based on derivatives inadequate. Secondly, as second derivatives are non existent, inference is difficult.

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[2] Balke and Forby (1997) present more general specifications than the one reflected by equation (1) which give rise to a variety of non linear behaviour that can be characterised by more than one threshold, as well as by attractor mechanisms to the long equilibrium relationship different than the adjustment coefficients.

Finally, the threshold parameter is not identified under the null of a linear model, involving a Davies' problem (Davies, 1987). Optimal classical statistics solutions to such a problem have been proposed by Andrews (1993) and Andrews and Ploberger (1994) and involve concentrating out nuisance parameters with respect to *a priori* supremum, or weighting function and calculating the critical values for the likelihood ratio statistics, as their distribution is non-standard. Hansen and Seo (2002) (hereinafter HS) follow such an approach and propose a quasi-MLE method based on a grid-search over the cointegrating vector and the threshold parameter, in conjunction with a 'fixed regressor' bootstrap to test for threshold effects in a two regime model as in (1) with $\lambda_1 = \lambda_2 = \lambda$ and $\beta_0 = \beta_2 = 0$ .In more detail, HS consider the following likelihood function:

$$L\left(a_i', \gamma_i', \beta_1, \Sigma, \lambda\right) = -\frac{n}{2}\log|\Sigma| - \frac{1}{2}\sum_{t=1}^{n} u_t\left(a_i', \gamma_i', \beta_1, \Sigma, \lambda\right)' \sum_{t=1}^{n}{}^{-1} u_t\left(a_i', \gamma_i', \beta_1, \Sigma, \lambda\right) \tag{2}$$

The maximisation of the likelihood function is feasible if parameters $\left(\alpha_i', \gamma_i', \Sigma\right)$ are concentrated out, by holding parameters $\left(\beta_1, \lambda\right)$ fixed, thus giving rise to a constrained maximum likelihood estimator that is equivalent to OLS regressions of Δy_t on e_{t-1} and $\Delta y_{t-1}, \ldots, \Delta y_{t-k}$, for each fixed β_1 and λ, and for sub-samples for which $e_{t-1} \leq \lambda$ and $e_{t-1} > \lambda_1$. As a second step, the estimates $\left(\hat{a}_i'(\beta_1, \lambda), \hat{\gamma}_i'(\beta_1, \lambda), \hat{\Sigma}(\beta_1, \lambda)\right)$ are utilised to yield the concentrated likelihood:

$$L\left(\beta_1, \lambda\right) = L\left(\beta_1, \lambda, \hat{a}_i'(\beta_1, \lambda), \hat{\gamma}_i'(\beta_1, \lambda), \hat{\Sigma}(\beta_1, \lambda)\right) = -\frac{n}{2}\log\left|\hat{\Sigma}(\beta_1, \lambda)\right| - n \tag{3}$$

The maximum likelihood estimator $\left(\hat{\beta}_1, \hat{\lambda}\right)$ minimise $\log\left|\hat{\Sigma}(\beta_1, \lambda)\right|$ subject to a constraint that reflects the probability P that observations lie over or below the threshold:

$$\pi_0 \leq P\left(e_{t-1} \leq \lambda\right) \leq 1 - \pi_0 \tag{4}$$

with π_0 being a trimming parameter usually taking values between 0.05 and 0.15 (Andrews, 1993). As the function L is jagged and not differentiable, HS propose a numerical maximisation method based on an evenly grid-search starting with consistent estimates of the cointegrating vector from the linear model. The algorithm sets sub-regions of the parameter space for β_1 based on a confidence interval $\left[\beta_{1l}, \beta_{1u}\right]$ constructed from the linear estimate $\tilde{\beta}_1$. For the threshold parameter λ , a sub-region $\left[\lambda_l, \lambda_u\right]$ is obtained so as λ_l and λ_u represent the π_0^{th} and $\left(1-\pi_0\right)^{th}$ percentiles of the error correction term, e_{t-1}, also based on the linear estimate $\tilde{\beta}$. The null of linear cointegration is tested against the alternative of threshold cointegration by means of the supremum Lagrance Multiplier test following Andrews (1993) and Andrews and Ploberger (1994):

$$\text{SupLM} = \sup_{\lambda_l \leq \lambda \leq \lambda_u} \text{LM}\left(\tilde{\beta}_1, \lambda\right) \tag{5}$$

The function $\text{LM}\left(\tilde{\beta}_1, \lambda\right)$ is not differentiable and, therefore, is evaluated by means of a grid-search over $\left[\lambda_l, \lambda_u\right]$. HS also propose an additional test SupLM^0 with β_1 fixed at an *a priori* known, or estimated value. Both LM tests are based on the restriction that the adjustment coefficients are equal, $\alpha_1 = \alpha_2$, whilst the threshold parameter is treated as fixed to its maximum likelihood value. Since the asymptotic distribution of the test is not known, it is approximated by means of a residual bootstrap. In addition, HS perform a 'fixed regressor' bootstrap procedure, holding the regressors fixed at their sample values (Hansen 1996). The fixed regressor bootstrap is not intended to approximate better the finite sample distribution of the test statistic, as compared to the residuals bootstrap, but allows for heteroscedasticity of unknown form and provides heteroscedasticity-consistent standard errors.

4 EMPIRICAL RESULTS

The HS estimation method is applied on logarithmic transformations of weekly prices for crude oil, ethanol and sugar in Brazil, expressed in Real for the period between July 2000 and May 2006.[3] Price pairs of sugar and oil, ethanol and oil and ethanol and sugar are examined, assuming that causation runs from oil to ethanol to sugar. The method comprises of a series of tests, starting with tests for unit roots and stationarity, estimation of cointegrating relationships between price pairs and testing for non cointegration, estimation of linear and threshold VECMs and testing for the null of linearity against the alternative of threshold-type non linearity. The direction of causation in the variables is also investigated by means of Granger causality tests.

Table 1 presents the results of unit root and stationarity tests. Augmented Dickey-Fuller (ADF) and Phillips and Peron (PP) tests for unit roots were conducted along with the Kwiatkowski-Phillips-Schmidt-Shin (KPPS) test for stationarity.[4] The lag lengths were selected using the Schwarz-Bayes information criterion. The ADF and PP tests fail to reject the null of unit roots around the mean for all series at a 5 percent level of significance. Similar evidence is provided by the KPPS tests that rejects stationarity around the mean for all prices at the same confidence level.

The inclusion of a deterministic trend in the tests provides some mixed evidence. For sugar the unit root tests fail to reject the null of unit roots around a trend at a 5 percent level. For oil and ethanol the tests suggest that the price series could be consistent with being stationary around a trend at a 10 percent level. Likewise the KPPS test fails to reject the null of stationarity around a trend at 10 percent level of significance. Nevertheless, all tests indicate that, at 5 percent confidence level the series are consistent with unit roots around a trend.

The results of tests for pair-wise non cointegration between sugar-oil, ethanol-oil, and sugar-ethanol are presented in Table 2. Non cointegration is tested utilising the Johansen approach that is based on MLE and likelihood ratio testing (Johansen, 1988). The results provide sufficient evidence for cointegration for all price pairs, being consistent with variables that contain unit roots and have a stationary cointegrating relation. The tests were conducted on the basis of a cointegrating VAR with a drift and linear trends in the series levels, whilst the lag length was selected by means of the Swartz-Bayes information criterion.[5]

The TVECMs are estimated following HS, on the basis of a *a priori* known cointegrating vector, estimated by means of the Johansen MLE approach, as opposed to estimating the cointegrating vector by means of the grid-search. Consequently, testing for linear cointegration is conducted by means of the SupLM^0 test-statistic. This choice of sequential testing procedure, consisting of tests for non cointegration, estimation of the cointegrating vector, estimation of TVECMs and tests for linear cointegration, may suffer from bias and, therefore, warrants more discussion. HS restrict the deterministic components of the VAR to include an intercept in the VECM, that reflects linear deterministic trends in the levels of the variables that cancel in the cointegrating relation, but no intercept in the cointegrating relation. In the case of prices in different, but strongly related markets, an

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[3] Data on sugar and ethanol have been collected by UNICA, the Sao Paolo Sugar Cane Agro Industry Union (www.unica.com.br), data on world oil prices have been collected from the Energy Information Administration of the United States government (www.iea.doe.gov) and information on exchange rates from the Federal Reserve Bank of St Louis (http://research.stlouisfed.org). The data covers the period 24 March 2001 to 20 May 2006.

[4] For these tests see Dickey and Fuller (1979), Phillips and Perron (1988) and Kwiatkowski *et al.* (1992).

[5] For the sugar-oil and ethanol-oil price VARs, two lags were selected. For the ethanol-sugar VAR, four lags were selected. Intercept was included in both the cointegration relation and the VECM.

intercept ought to be included in the cointegration relation in order to ensure that the error correction term in the VECM has a zero mean. As a grid-search over three parameters, that is the intercept, β_1 and λ, would be very complex and possibly intractable, it is preferable that the application relies on *priori* estimated cointegrating vector. HS conduct Monte Carlo experiments in order to assess the power of both tests and report that the SupLM^0 test has slightly more power that the SupLM in which the cointegrating vector is initially unknown and is estimated by means of the grid-search.

Estimates of the linear VECMs and the threshold VECMs, obtained by means of the HS approach are presented in Tables 3 to 5. The grid-search for λ is conducted over 300 grid points, whilst both the fixed regressor and the standard residual bootstrap experiments are used to calculate the *p*-values for the SupLM^0 test. Note that in all estimated linear VECMs, the adjustment coefficients are very low, mainly due to the high frequency of the price series.

For the sugar-oil price pair, presented in Table 3, the estimated cointegrating parameter is 0.64, implying that a 10 percent increase in the price of oil brings about over 6 percent increase in the price of sugar. The value of the SupLM^0 test is 27.47 and the *p*-values are 0.028 and 0.004 for the fixed regressor and the residual bootstraps respectively, supporting the threshold cointegration hypothesis. In both linear and threshold VECMs, the statistical significance of the estimates of the adjustment parameters also reveal that oil is the dominant market. The adjustment parameter in the sugar ECM is statistically significant, whilst that in the oil ECM is not, suggesting that, in the long run, oil prices Granger-cause sugar prices, but not *vice versa,* an outcome that corresponds to long run weak exogeneity in the econometric sense (Granger, 1988). The estimated threshold is -0.02 and identifies two regimes, an extreme regime and a typical one, with 16 and 84 percent of the data respectively. The adjustment coefficients in the sugar price ECM indicate a fast adjustment to the long run equilibrium in the extreme regime with an estimated value of -0.21.[6] In the typical regime adjustment is about twenty times slower. Since the TVECM has been estimated using the logarithmic transformations of the variables, the threshold implies that adjustment is fast when sugar prices lie 2 percent below its long run equilibrium, as this is determined by the price of oil. Nevertheless, a value for the threshold that as low as -0.02 may indicate that the nonlinear relationship between the sugar and oil prices can be characterised as asymmetric, allowing the variables to adjust differently depending on whether the disequilibria are negative or positive (see Granger and Lee, 1999; Escribano and Pfann, 1997; Enders and Granger, 1998). The estimated parameters of the differenced terms that reflect the transitory effects, suggest strong autoregressive behaviour in sugar, with a change in one week been followed by a somewhat smaller change in the following week. In the short run, oil prices also influence the sugar price level, at least in the typical regime. Nevertheless, most of effect is accounted for by the error correction term.[7]

The estimates of the ethanol-oil price relationship are presented in Table 4. The cointegrating parameter estimate is 0.60, approximately equal to that in the sugar-oil price relationship. In the long run changes in oil prices are incompletely, but for the most part, transmitted to the price of ethanol. The SupLM^0 test statistic has a value of 21.47, with *p*-values the fixed regressor and the residual bootstraps at 0.02 and 0.03 respectively, strongly rejecting the null of linear cointegration. The threshold parameter λ, is estimated at -0.20, with an extreme regime accounting for 18 percent of the observations and a typical regime accounting for 82 percent of the observations. As in the case of sugar, oil appears to influence the long run development of ethanol, in a 'no levels feedback' manner. As the adjustment parameters in the oil ECMs are statistically not significant, oil can be considered as a driving trend in the system, causing ethanol prices in the Granger sense, being weakly exogenous. The threshold in conjunction with the estimated adjustment parameters in the ethanol price threshold

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[6] Figures of the negative log-likelihood function against threshold values, as well as figures depicting the relationship between the error correction term and the change in the dependent variable in the VECM are presented in the Appendix.

[7] This observation is based on the statistical significance of the estimates in the TVECMs. It is important to note that HS have no formal distribution theory and therefore warn towards cautious interpretation of the Eicker-White standard errors.

ECMs indicate that ethanol price adjusts about four times as fast to its long run equilibrium in the extreme regime, as compared to the typical regime.

The adjustment is rapid at approximately 12 percent of the disequilibrium, when ethanol and oil prices abide by the relationship, $\text{Ethanol} \leq 0.60\text{Oil} - 0.20$, indicating that for this fast adjustment to take place the price of ethanol should be about 20 percent below its long run trend, as this is determined by the price of oil. The estimated TVECMs also suggest that there is also strong autoregressive behaviour in both ethanol and oil prices. However, the evidence suggests that the transitory effects between the prices are minimal. The support for threshold cointegration, as well as the estimated threshold value confirm the *a priori* expectations for non linear behaviour between ethanol and oil prices. Oil price increases appear to draw ethanol prices at a faster rate, when the later are possibly sluggish, lingering well below their long run trend.

For the ethanol-sugar price relationship, the results are presented in Table 5. The cointegrating parameter is estimated to be equal to 0.87, a level quite close to unity. The restriction of β_1=1 was imposed and tested by means of a Likelihood Ratio test, as suggested by Johansen and Juselius (1992). The test statistic value is 1.02 and hence accepted with a *p*-value of 0.31, suggesting that changes in the price of sugar are completely transmitted to the price of ethanol in the long run. The SupLM^0 test fails to reject the null of linear cointegration. The value of the test statistic is 27.61 and the *p*-values are 0.12 and 0.19 for the fixed regressor and the residual bootstraps respectively. The estimated linear VECM suggests that sugar is weakly exogenous and influences the long run behaviour of ethanol prices without being influenced by them, as the adjustment parameter in the sugar price ECM is not statistically significant.

This is a rather surprising result as one would expect that increases in the demand of ethanol and subsequent increases in the derived demand for sugarcane would mean that ethanol prices Granger-cause sugar prices, with the direction and order of causation running from oil to ethanol to sugar. These results reflect, in conjunction with the results of the sugar-oil price relationship, that sugar prices respond to changes in the price of oil, instead of following ethanol prices.

5 CONCLUSIONS

The paper explored the possibility of threshold effects in the relation between prices of sugar and oil, ethanol and oil, and ethanol and sugar, applying a threshold cointegration approach for estimation and inference. The results supported threshold cointegration and nonlinear adjustment in the sugar and oil, and sugar and ethanol price relations. Thus, the evidence supports *a priori* expectations on the behaviour of these prices, with the dynamic behaviour of sugar and ethanol prices being governed by oil prices and more specifically by the magnitude of the spread between these prices and the price of oil. It appears that when the spread widens, with the price of oil increasing and the prices of sugar and ethanol being sluggish, adjustment of these prices to their long run trend is fast.

More specifically, the long run behaviour of sugar prices was found to be determined by oil prices and, rather surprisingly, not ethanol prices. Sugar prices were found to adjust to a long run equilibrium, determined by oil prices, in an asymmetric manner with adjustment being fast when disequilibria are negative and slow when disequilibria are positive. Ethanol prices were also found to be Granger-caused by oil prices. Evidence for threshold behaviour suggested that ethanol prices adjust fast when they are about 20 percent lower than their trend, as this is determined by oil prices. There was no evidence for threshold cointegration between sugar and ethanol prices. The results suggested that these prices were linearly cointegrated, with sugar being the dominant market. In addition, the tests suggested that changes in the sugar price are completely transmitted to the price of ethanol. The direction and order of causation is also an interesting finding. The results indicate that oil is weakly exogenous and Granger-causes both ethanol and sugar. It appears that sugar responds directly to oil price changes and not in a successive manner through ethanol prices.

The analysis suggests that if the prices of crude oil were to continue their upward trend, ethanol will become increasingly competitive with petrol. World sugar prices will also increase, as Brazil, the major sugar producer and exporter in the world, influences the world sugar balance. In the long run, sustained high sugar prices may result in an increase in the area under sugarcane not only in Brazil, but also in developing low-cost producer countries, such as Malaysia, that already invest in ethanol production. Nevertheless, it is important to note that infrastructure development is necessary to allow further expansion of the area under sugar cane and of ethanol processing capacity in many countries. Whilst, high sugar prices may benefit sugar farmers, consumers become worse off. Such a development may not result in net losses for many developing countries that produce and export sugar, given the relatively low expenditure on sugar as compared to other foods. Nevertheless, sustained increases in oil price may result, in the long run, in increases in the prices of other commodities that may be used for ethanol production, such as maize, as well as changes in the allocation of land. Given the large volume of food commodities that are necessary for the production of relatively small volumes of biofuels and the demand for food, the co-movement of oil and food commodities may result in high food prices, thus exerting additional pressure on food importing developing and least developed countries.

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TABLES

Table 1 – Unit root and stationarity tests

	without time trend			with time trend		
	ADF	**PP**	**KPPS**	**ADF**	**PP**	**KPPS**
Sugar	-1.82	-1.68	0.71	-2.44	-2.26	0.12
Oil	-1.33	-1.18	1.89	-3.39	-3.39	0.13
Ethanol	-2.48	-2.23	0.73	-3.17	-2.70	0.13

Note: Critical values for the ADF and PP tests without (with) a deterministic trend for 5 percent and 10 percent level of significance are -2.87 (-3.42) and -2.57 (-3.14) respectively. For the KPPS the corresponding critical values are 0.46 (0.15) and 0.35. (0.12).

Table 2 – Johansen test for non cointegration

	Maximum Eigenvalue test	
	Zero *vs* One Cointegrating Vector	**One *vs* Two Cointegrating Vectors**
Sugar-Oil	20.18	2.59
Oil-Ethanol	19.54	2.54
Ethanol-Sugar	21.44	2.98

Note: Critical values for 5 percent and 10 percent level of significance are 15.49 and 13.42 respectively.

Table 3 - Linear and Threshold VECMs: Sugar - Oil

Regressors	**Linear VECM**		**Threshold VECMs**			
			First regime		**Second regime**	
Percentage of observations			**0.16**		**0.84**	
	Sugar	**Oil**	**Sugar**	**Oil**	**Sugar**	**Oil**
e_{t-1}	-0.0261 *0.0061*	-0.0076 *0.0116*	-0.2099 *0.0587*	-0.0409 *0.0821*	-0.0154 *0.0067*	-0.0165 *0.0155*
ΔSugar, *t*-1	0.7799 *0.0512*	-0.0416 *0.0789*	0.9570 *0.0882*	-0.1344 *0.1214*	0.7508 *0.0557*	-0.0258 *0.0923*
ΔOil, *t*-1	0.0533 *0.0269*	0.2377 *0.0625*	0.0515 *0.1007*	0.2190 *0.1396*	0.0594 *0.0281*	0.2379 *0.0674*
Intercept	0.0061 *0.0022*	0.0048 *0.0034*	-0.0061 *0.0052*	-0.0034 *0.0098*	0.0023 *0.0026*	0.0083 *0.0051*
Cointegrating Vector	(-1, 0.6428)					
Threshold Estimate λ	-0.0169					
$\mathbf{SupLM^0}$	Test Statistic Value: 21.4723 Fixed regressor *p*-value: 0.0286 Residual bootstrap *p*-value: 0.0470					

Note: Standard errors in italics

Table 4 - Linear and Threshold VECMs: Ethanol - Oil

Regressors	**Linear VECM**		**Threshold VECMs**			
			First regime		**Second regime**	
Percentage of observations			**0.18**		**0.82**	
	Ethanol	**Oil**	**Ethanol**	**Oil**	**Ethanol**	**Oil**
e_{t-1}	-0.0389 *0.0136*	-0.0099 *0.0122*	-0.1279 *0.0594*	0.0027 *0.0386*	-0.0332 *0.0165*	0.0104 *0.0208*
ΔEthanol, *t*-1	0.5512 *0.0751*	0.0264 *0.0550*	0.5397 *0.1248*	-0.0345 *0.0699*	0.5870 *0.0898*	0.0648 *0.0680*
ΔOil, *t*-1	0.0195 *0.0425*	0.2331 *0.0638*	0.2824 *0.1805*	0.0283 *0.1490*	-0.0289 *0.0358*	0.2707 *0.0677*
Intercept	0.0006 *0.0136*	0.0029 *0.0027*	-0.0378 *0.0235*	0.0168 *0.0156*	0.0008 *0.0023*	0.0002 *0.0208*
Cointegrating Vector	(-1, 0.6004)					
Threshold Estimate λ	-0.2059					
$SupLM^0$	Test Statistic Value: 21.4686 Fixed regressor *p*-value: 0.0250 Residual bootstrap *p*-value: 0.0364					

Note: Standard errors in italics

Table 5 - Linear and Threshold VECMs: Ethanol - Sugar

Regressors	**Linear VECM**	
	Ethanol	**Sugar**
e_{t-1}	-0.0682 *0.0195*	0.0003 *0.0106*
ΔEthanol, *t*-1	0.6678 *0.0989*	0.1176 *0.0476*
ΔSugar, *t*-1	0.0571 *0.1047*	1.0262 *0.0703*
ΔEthanol, *t*-2	-0.2583 *0.1074*	0.0256 *0.0557*
ΔSugar, *t*-2	0.0606 *0.1292*	-0.3439 *0.0869*
ΔEthanol, *t-3*	0.1306 *0.0781*	-0.0499 *0.0451*
ΔSugar, *t*-3	-0.1457 *0.0712*	-0.0650 *0.0746*
Intercept	0.0007 *0.0021*	0.0008 *0.0012*
Cointegrating Vector	(-1, 0.8711)	
Threshold Estimate λ	0.0331	
$SupLM^0$	Test Statistic Value: 27.6159 Fixed regressor *p*-value: 0.1226 Residual bootstrap *p*-value: 0.1786	
H_0: β_1=1 vs H_1: $\beta_1 \neq$1 Likelihood Ratio test	Value: 1.02 *p*-value: 0.31	

Note: Standard errors in italics

APPENDIX

Figure A1 – Sugar-Oil Concentrated Negative Log-Likelihood

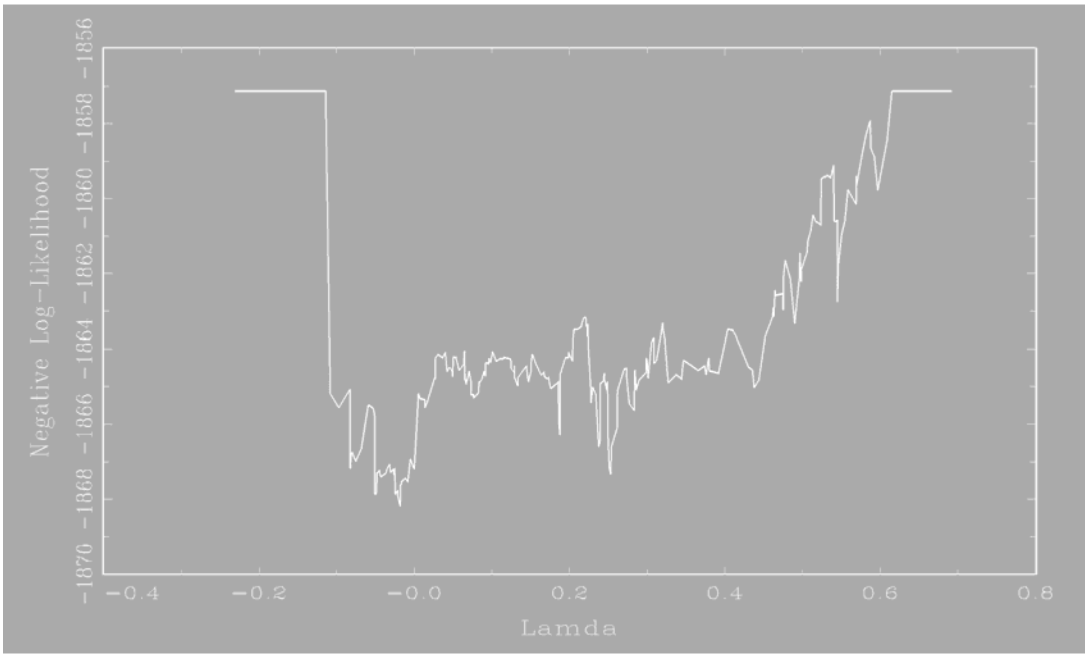


Figure A2 – Sugar Price Response to Error Correction

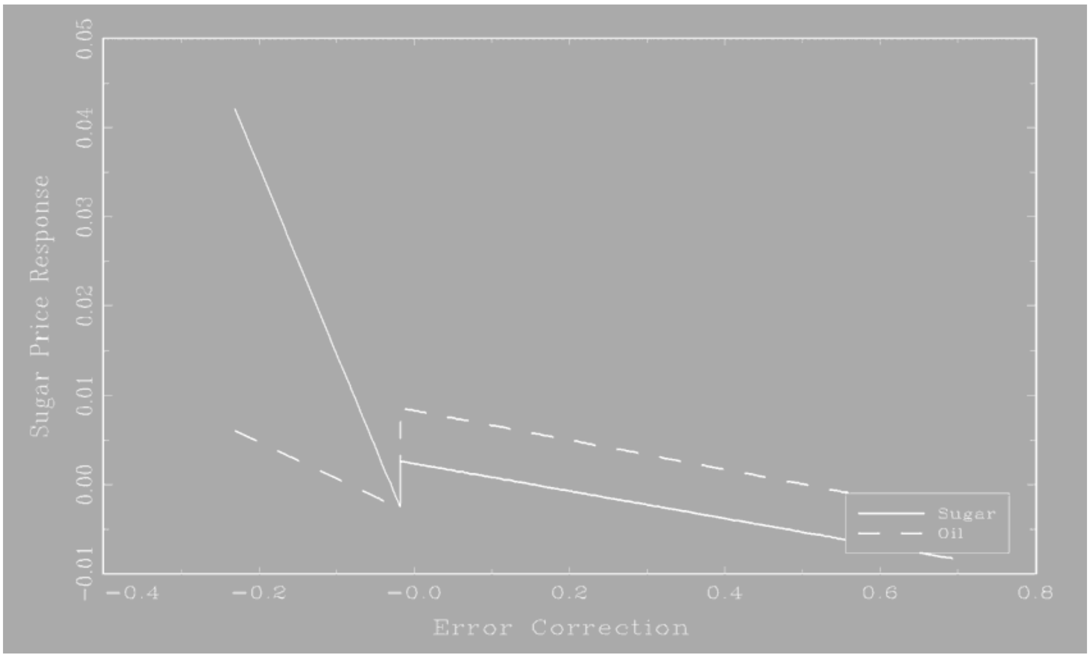


Figure A3 – Ethanol-Oil Concentrated Negative Log-Likelihood

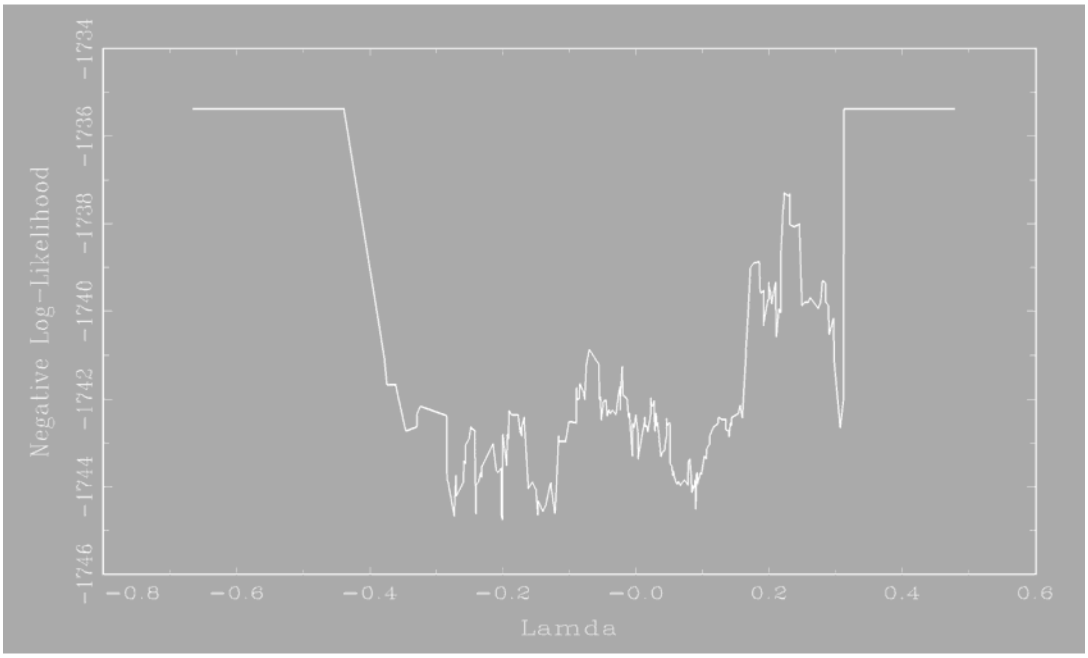
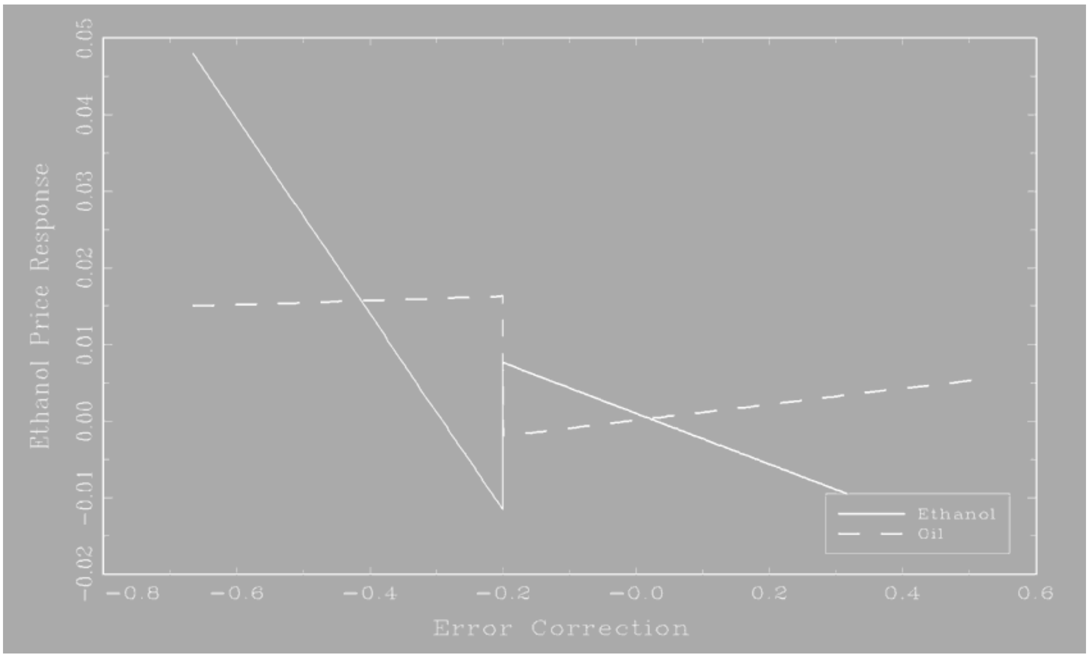


Figure A4 – Ethanol Price Response to Error Correction



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