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THE INFLUENCE OF THE EC CRUSHING SECTOR
ON WORLD PRICES FOR SOYBEAN PRODUCTS

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SEPTEMBER 1989

This paper is an attempt to highlight the influence of EC processors' activity on the long-run determination of Rotterdam prices for soybean seeds, soybean oil and soybean meal. Empirical results show that these prices are cointegrated and that their co-evolution is mainly attributable to the joint product aspect of meal and oil. It is then shown that the cointegration equation can be interpreted as the aggregate optimal pricing policy for EC crushers.

Key words: EC crushing soybean prices market cointegration.
THE INFLUENCE OF THE EC CRUSHING SECTOR

ON WORLD PRICES FOR SOYBEAN PRODUCTS
INTRODUCTION

By simultaneously buying soybeans and producing soymeal and soyoil the crushing sector plays a central role on the relationships between the various soybean product prices. The processing of a given quantity of soybean seeds produces both meal and oil in relatively fixed proportions. Once the amount of seeds to be crushed has been determined, market supplies of meal and oil are fixed. They are then sold into end use markets or held as inventory. The joint-product aspect of meal and oil is thus due to the fact that meal production and oil production are locked together through technically fixed crushing yields for meal and oil: one unit of soybean seeds includes nearly 0.78 unit of meal and 0.18 unit of oil. Concerning the demand side of the markets, soymeal and soyoil demand are independant: soymeal is exclusively used for animal feed while soyoil is for human consumption. Thus, soymeal, soyoil and soybeans are essentially linked together through the crushing industry.

The EC imports soybean cakes which are mainly used for animal feed and soybeans which are crushed within the EC. The EC is self-sufficient in vegetable oils and a great part of soybean oil produced as a by-product of the crushing process is thus exported. In 1987, EC processors globally imported about 8 million metric tons (MT) of soybean meal and 14 million MT of soybeans. These soybeans were quasi-exclusively crushed to produce 2.6 million MT of soybean oil of which nearly 24% were exported and 11 million MT of soybean meal.
The object of this paper is to show how Rotterdam prices for soybeans, soymeal and soyoil are influenced by the link created between these products by the EC crushing sector.

EC imports and exports for soybean products are essentially exchanged on the Rotterdam market. Several studies have already performed comparative static analysis of the interactions among the three market and have put in light the fact that the prices of soybeans, meal and oil are simultaneously determined due to the joint product relationship (Meyers and al ¹, Peterson-Auerback ²). However dynamic price relationships among the three products were never analysed.

According to the Efficient Market Hypothesis (Fama ³), these markets are said efficient if price changes instantaneously and fully reflect all the available information (in the absence of transaction costs). If it is so, movements in world Rotterdam prices for soybean products cannot be predicted ahead of time. At the opposite, if these markets fail to clear at each point in time (for example when firms do not make instantaneous adjustments in response to a changing market situation), then prices only slowly and gradually adjust to their equilibrium level. In this case, an analysis of this adjustment process can provide useful information on the price discovery mechanisms and the characteristics (degree of inefficiency) of these markets. This is the view adopted in this study since linkages between Rotterdam price for soybeans, soymeal and soyoil are investigated by using time series models. The analysis of the price discovery process for these commodities may highlight the influence of the EC crushing sector on the process by which equilibrium Rotterdam prices for soybean products are reached. The method used in this paper rely on integration and cointegration tests. The advantage of such a procedure is that it permits to distinguish between long and short-run mechanisms which underly the price adjustment process.
II- ANALYSIS OF THE SERIES

Data are weekly Rotterdam (CIF) prices for soybean seeds (PS), soybean oil (PO) and soybean cake - 44% protein - (PC). All prices are nominal and $ US per MT. The data period is January 1, 1981 to July 16, 1987. Data are converted to natural logs.

Usual causality analysis rely on an examination of Granger causal links within bivariate or multivariate autoregressive models. These methods only deal with stationary data, so that first differences of the data are generally taken to remove linear trends and deterministic components. However, as was pointed out in Granger, if the series are integrated of the first order and cointegrated "many of the papers discriming causality tests based on the traditional time series modelling techniques could have missed some of the forecastability and hence reached incorrect conclusions about non-causality in mean". Due to this, tests for cointegration between soybean product prices are first performed as a preliminary to the analysis of dynamic relationships among these prices. Sims, Stock and Watson show that the asymptotic distributions of causality tests are sensitive to unit roots and time trends in the series. We thus begin by characterizing the time trends and unit root properties of the data with several Dickey-Fuller (Dickey-Fuller) tests.

A series \( x_t \) is integrated of order \( d \) (denoted \( x_t \sim I(d) \)) if it is a series which has a stationary, invertible, non-deterministic ARMA representation after differencing \( d \) times \( ((1-L)^d x_t \sim I(0)) \). In other words \( x_t \) has \( d \) unit roots.
Table 1 investigates the possibility that PO, PS and PC series might have up to two unit roots and a first order linear time trend of the first order. Usual Dickey-Fuller (DF) and augmented Dickey-Fuller (ADF) tests for unit roots are used to determine the order of integration of the series while the presence of a deterministic time trend is tested with $\phi_2$ and $\phi_3$ likelihood ratio statistics described in Dickey-Fuller.\(^6\)

The general procedure for these tests is to run the regressions:

$$\Delta^d x_t = \alpha \Delta^{d-1} x_{t-1} + \sum_{i=1}^{p} \beta_i \Delta^{d} x_{t-i} + u_t \quad (1a)$$

and

$$x_t = a + b.t + c.x_{t-1} + \sum_{d=1}^{h} d_j . \Delta^{d} x_{t-j} + v_t \quad (1b)$$

where \(t\) is a deterministic time trend and $\Delta^d x_t = (1-L)^d x_t$ with $L^k x_t = x_{t-k}.u_t$ and $v_t$ are white noise process.

Equation (1a) is used to perform DF and ADF tests with $p = 1$ for DF and $p$ large enough to take account of remaining autocorrelation for ADF. Both tests are the Student $t$-statistics for the coefficients of $\Delta^{d-1} x_{t-1}$ and rely on rejecting the hypothesis that the series $\Delta^{d-1} x_t$ is I(1) in favour of stationarity.

Dickey-Fuller likelihood ratio statistics $\phi_2$ and $\phi_3$ are based on equation (1b). Null hypothesis are respectively $c=1$ and $a=b=0$ with $\phi_2$ and $c = 1$ and $b = 0$ with $\phi_3$. Distributions are non-standard for all these tests. Critical values are given in Engle-Granger\(^7\) for DF and ADF and in Dickey-Fuller\(^6\) for $\phi_2$ and $\phi_3$.\(^8\)
Table 1. Tests for unit roots and time trends

<table>
<thead>
<tr>
<th>Series</th>
<th>DF (d=1)</th>
<th>ADF (d=1)</th>
<th>DF (d=2)</th>
<th>ADF (d=2)</th>
<th>$\varphi_2$</th>
<th>$\varphi_3$</th>
</tr>
</thead>
<tbody>
<tr>
<td>PS</td>
<td>-1.01</td>
<td>-0.93</td>
<td>-11.02</td>
<td>-7.21</td>
<td>1.61</td>
<td>2.32</td>
</tr>
<tr>
<td>PC</td>
<td>-1.04</td>
<td>-0.96</td>
<td>-12.93</td>
<td>-7.36</td>
<td>1.50</td>
<td>2.20</td>
</tr>
<tr>
<td>PO</td>
<td>-0.76</td>
<td>-0.78</td>
<td>-14.28</td>
<td>-7.61</td>
<td>1.10</td>
<td>1.50</td>
</tr>
</tbody>
</table>

* indicates rejection of the respective null hypothesis at the 1% level.

The DF and ADF regressions with first differences (d = 1) lead to non significative t statistics. At the opposite, running the same models with second differences (d = 2) gave significative t statistic indicating the first difference of each serie is stationary. Both $\varphi_2$ and $\varphi_3$ are always non significative so that the hypothesis of a linear deterministic time trend is always rejected by the data. Each serie is thus I (1) and needs to be differenced once to achieve stationarity.

III- TESTING FOR COINTEGRATION

We now search for cointegration relationships between the prices for soybean products. Let $X_t = (x_{1t}, \ldots, x_{Nt})$ be a vector of N individually I (d) time series. $X_t$ is cointegrated of order d, b
(denoted $X_t \in (CI(d;b))$ if there exist a vector $\gamma$ such that $Z_t = \gamma'X_t$ is integrated of order $(d-b)$ with $b > 0$. $(\gamma'X_t)$ is the equilibrium or long-run relationship while $Z_t$ represents equilibrium error. The vector $\gamma$ is called the cointegrating vector.

Here, each price serie is $I(1)$, we thus search for a linear combination of the three series that is itself $I(0)$. As proposed by Engle-Granger, we first run regressions between each combination of the levels of the three series PS, PC and PO. We then test for the stationarity of residuals by both using the standard Durbin Watson test with the critical values presented by Sargan and Bargava under the unit root null hypothesis (this test is called the cointegrating regression Durbin-Watson (CRDW) test) and the DF and ADF (with four lags) tests. Table 2. presents results from simple OLS regressions.

Table 2. Cointegration tests and estimates

<table>
<thead>
<tr>
<th>Dep. Var</th>
<th>Cst</th>
<th>PC</th>
<th>PO</th>
<th>$R^2$</th>
<th>CRDW</th>
<th>DF</th>
<th>ADF(4)</th>
</tr>
</thead>
<tbody>
<tr>
<td>PS</td>
<td>70.67</td>
<td>0.87</td>
<td>0.57</td>
<td>0.031</td>
<td>-1.24</td>
<td>-1.69</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(8.31)</td>
<td>(21.41)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>PS</td>
<td>150.21</td>
<td>0.20</td>
<td>0.51</td>
<td>0.060</td>
<td>-2.35</td>
<td>-2.28</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(27.31)</td>
<td>(18.91)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>PO</td>
<td>397.48</td>
<td>0.51</td>
<td>0.01</td>
<td>0.040</td>
<td>-1.22</td>
<td>-1.51</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(8.49)</td>
<td>(2.27)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>PS</td>
<td>0.93</td>
<td>0.78</td>
<td>0.175</td>
<td>0.97</td>
<td>0.65*</td>
<td>-5.50*</td>
<td>-4.21*</td>
</tr>
<tr>
<td></td>
<td>(0.38)</td>
<td>(74.07)</td>
<td>(68.90)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

* significant at the 1% level.

Numbers in parentheses are $t$ values of estimated coefficients.
While we are unable to reject the hypothesis that every pair of the three variables is not cointegrated, the three variables PS, PC and PO do appear as cointegrated. The corresponding cointegrating regression is given by the last row in Table 2. In this case, the CRDW, DF and ADF statistics are well above their respective critical values so that we cannot reject the hypothesis that these variables are not cointegrated even at the 1% level of significance. However, before going on to interpreting this result it is necessary to consider the fact that in general the cointegration regression cannot be expected to be unique and we have to check that using any of the three variables as the dependant variable does not modify the cointegrating vector. The cointegration regression is thus reestimated twice by respectively using PC then PO as the dependent variable. Table 3 compares the corresponding equilibrium relationships by reporting the various regressions rearranged so that PS is on the left hand side.

Table 3. Comparison of the estimated cointegrating vectors

<table>
<thead>
<tr>
<th>Dep. Var</th>
<th>Constant</th>
<th>PC</th>
<th>PO</th>
<th>R²</th>
<th>CRDW</th>
<th>DF</th>
<th>ADF(4)</th>
</tr>
</thead>
<tbody>
<tr>
<td>PS</td>
<td>0.93</td>
<td>0.78</td>
<td>0.17</td>
<td>0.97</td>
<td>0.65 *</td>
<td>-5.5 *</td>
<td>-4.2 *</td>
</tr>
<tr>
<td>PO</td>
<td>4.01</td>
<td>0.78</td>
<td>0.19</td>
<td>0.93</td>
<td>0.66 *</td>
<td>-5.5 *</td>
<td>-4.0 *</td>
</tr>
<tr>
<td>PC</td>
<td>8.21</td>
<td>0.83</td>
<td>0.17</td>
<td>0.94</td>
<td>0.62 *</td>
<td>-5.3 *</td>
<td>-4.1 *</td>
</tr>
</tbody>
</table>

* Significant at the 1% level
Excepted for the constant term, Student t-ratios of price coefficients were always highly significative. The three variables always appear to be cointegrated and only the coefficient of the constant term is highly sensible to the choice of the dependant variable.

IV- ECONOMIC INTERPRETATION OF COINTEGRATION RESULTS

In every rearranged regression given in Table 3, the coefficients attached to variables PC and PO are always quasi equal to the soybean meal (0.78) and oil (0.18) extraction rates. These results may be interpreted as resulting from the joint-product aspects of soybean meal and oil in the profit maximization program of processors. To see that, we assume the identity of European processors' production functions and their profit-maximizing behavior in the face of perfectly competitive markets for outputs and inputs so that there is a sense in which there exists an aggregate optimization program which may be written as:

\[
\text{Max } \pi (QS) = QC \cdot PC + QO \cdot PO - QS \cdot PS - Z \cdot PI \quad (2)
\]

with \[ \begin{align*}
QC &= QC (QS, Z) \quad \text{and } QC_1 > 0 \quad QC_2 > 0 \quad (3) \\
QO &= QO (QS, Z) \quad \text{and } QO_1 > 0 \quad QO_2 > 0 \quad (4)
\end{align*} \]

where \( \pi \) are profits; QC, QO are the produced quantities of soybean cakes and soybean oil while QS is the amounts of seeds used in production; Z represents other inputs and PI stands for the price of those other inputs. QC\(_1\) and QO\(_1\) \((i = 1, 2)\) are partial derivatives of cake and oil production functions. Second order partial derivatives are assumed to be negative.
Substituting (3) and (4) into (2) gives the long-run conditions for profit maximisation:

\[
\begin{align*}
\frac{\partial m}{\partial Q_S} &= QC_1 \cdot PC + QO_1 \cdot PO - PS = 0 \\
\frac{\partial m}{\partial Q_O} &= QC_2 \cdot PC + QO_2 \cdot PO - PI = 0
\end{align*}
\] (5) (6)

Approximating the partial derivatives of meal and oil production functions with respect to seeds by the average meal and oil extraction rates (respectively 0.78 and 0.18), the first condition may be rewritten as:

\[PS = 0.78 \cdot PC + 0.18 \cdot PO\] (7)

Under competitive conditions, the previously estimated cointegration relationships between soybean product prices may thus be interpreted as the aggregate optimal pricing policy for crushers (with a technology as described above). Thus, whereas demand movements can distort the price relationship (7) in the short-run, this relationship needs to be true in the long-run since it is the solution to the optimization problem of crushers. These results tend to confirm that there exists a long-run equilibrium relationship between the prices for the three soybean products which reflects the fixed proportions of meal and oil which are locked together in the whole bean. Soybean meal and soybean oil are essentially independant of each other in use and the economic forces underlying the demand for meal are largely unrelated to those underlying the demand for oil. However the joint product-aspect of meal and oil production ensures that the supply and the price for these commodities are tied very closely together in the long-run. As was graphically shown in Meyers and al. a rise in say the price of meal (due for instance to an exogenous shift in demand) will increase the crushing demand and will ultimately lead to an higher oil supply which will push oil prices downwards.
For instance, an exogenous increase in the price of soybean cakes \( \Delta PC \) > 0 will have two distinct effects on the price of oil as can be seen from equation (8) which is obtained by differentiating equation (7) around equilibrium:

\[
\frac{\Delta P0}{\Delta PC} = \frac{1}{0.18} \cdot \frac{\Delta PS}{\Delta PC} - \frac{0.78}{0.18}
\]

\( \Delta PS/\Delta PC \)

The first effect is a direct cost-push effect: due to the increase in cake supply induced by the rise in the price of cakes, the demand for beans is increased and the price for beans is thus pushed upwards by an amount \( \Delta PS/\Delta PC \) > 0. This will tend to increase the price of oil proportionally to the oil extraction rates as shown by the first term on the right side of equation (8). The second effect comes from the joint-product aspects of soymeal and soyoil: by producing more cakes, the crushing sector automatically produces more oil and this effect results in downwards pressures on the oil price. This second effect is also directly governed by the respective oil and meal contents of beans as shown by the second member of the right side of equation (8). Globally the ultimate effect of the initial exogeneous increase in the price for soymeal on the price for oil is negative or positive depending on wether \( \Delta PS/\Delta PC \) is larger or smaller than 0.78.

V- ERROR CORRECTION REPRESENTATION AND MARKET EFFICIENCY

There is an intimate connection between cointegrated series and Error Correction Models (ECM) representations. As was shown in Engle-Granger, if the vector \( X_t \) of soybean product prices is
cointegrated of order 1,1 ($X_t \sim CI (1,1)$), then there exists a generating mechanism having the general "error correction form":

$$\Delta X_t = -\lambda Z_{t-1} + \text{Lagged} \left( \Delta X_t \right) + d(B) \Sigma_t$$  (9)

with $\lambda \neq 0$ and $\Sigma_t$ is the vector of white noise disturbances.

It clearly appears from (9) that if the components of $X_t$ are $I(1)$ and cointegrated, there must be Granger causality between these series in at least one direction. $Z_{t-1}$ must occur in at least one equation and this knowledge of $Z_{t-1}$ must improve forecastability of at least one component of $X_t$. Furthermore, the presence of the lagged terms in the ECM provides a second source of causation.

If we define $ES$, $EC$ and $EO$ to be the derived residuals from cointegrating regressions obtained with respectively $PS$, $PC$ and $PO$ as the dependent variable, we may then include these residuals in a standard ECM model. A fairly simple search procedure produced the following equations:

\[
\Delta PS_t = 0.184 \Delta PS_{t-2} - 0.178 \Delta PS_{t-4} - 0.131 \Delta PS_{t-6} \\
- 0.160 \Delta PS_{t-14} + 0.030 \Delta PO_{t-3} + 0.037 \Delta PO_{t-4} \\
+ 0.427 \Delta PC_{t-4} - 0.151 \Delta PC_{t-5} \quad (10a)
\]

\[
Q(47) = 38.69 \; \text{SEE} = 6.62
\]

\[
\Delta PO_t = -0.178 EO_{t-1} - 0.308 \Delta PO_{t-1} - 0.136 \Delta PO_{t-2} \\
- 0.79 \Delta PC_{t-14} \quad (10b)
\]

\[
Q(49) = 30.22 \; \text{SEE} = 27.69
\]
\[ \Delta P_{C_t} = -0.065 \Delta C_{t-1} + 0.269 \Delta P_{C_{t-4}} - 0.100 \Delta P_{C_{t-5}} \]
\[ (-1.80) \quad (4.97) \quad (-1.91) \]
\[ -0.111 \Delta P_{C_{t-8}} - 0.101 \Delta P_{S_{t-6}} + 0.112 \Delta P_{C_{t-5}} \]
\[ (-2.06) \quad (-2.33) \quad (2.28) \]
\[ -0.091 \Delta P_{S_{t-14}} - 0.0275 \Delta P_{O_{t-10}} \]
\[ (-2.10) \quad (-2.43) \quad (10c) \]

\[ Q(48) = 30.58; \text{ SEE } = 5.56 \]

Where \( Q \) and \( \text{SEE} \) are respectively the Ljung-Box test for a random correlogram and the standard error of estimates. Number in parentheses are the \( t \)-values of estimated coefficients.

Only the soybean oil and soybean meal equations include significant lagged levels. These terms are of the appropriate sign and size for error correction terms and are individually significant or nearly so (an \( F \) test for the nullity of the error correction term in equation (10c) yielded a \( F \) statistic close to 3.25 with a significance level equal to 0.072).

On the ground of these three equations, there appear to be general interaction among the three prices since every price is influenced by the two other ones. Within each equation, the sum of lagged terms on every variable is always lower than the coefficients derived from the cointegration regression thus showing that adjustment towards the cointegration relationships is only partial in the short-run.

Finally, the structure of delays seems to show that while variations in the prices for soybean oil or soybean meal are both transmitted to the price of seeds with an average lag close to nearly 3.5 weeks, movements in the price of seeds are more quickly transmitted to oil consumers (average lag close to 2 weeks) than to meal consumers (average lag close to 10 weeks).
Lastly, a related issue is that the Rotterdam markets for soybean seeds, soybean oil and soybean meal do not seem to be "FAMA-efficient" markets. FAMA\textsuperscript{3} defined an efficient market as one that fully reflects all available information. According to this definitions Rotterdam markets for soybean products are clearly inefficient since cointegration implies that information concerning the prices for two of these commodities can help forecast the price of the third product (Granger\textsuperscript{4}).

V- CONCLUSION

In this paper, the dynamic relationships of world prices for soybean products (seeds, meal and oil) are investigated using integration and cointegration tests.

Empirical results suggest that there is evidence of equilibrium level relationships between the prices of soybean products. In the long-run, the price level of these commodities do not significantly depart from each other. Furthermore this long-run relationship seems mainly due to the joint product aspect of soymeal and soyoil production since the two coefficients of the cointegration regression just equal the meal and oil extraction rates and cointegration regression is just the aggregate optimal pricing policy for crushers under competitive conditions. The entire system (beans, oil and meal) thus moves towards a long-run equilibrium in a simultaneous manner with all three markets interacting with one another.
Three conclusions can be drawn from this result. At first, due to the link between cointegration and an error correction mechanism, results from unrestricted vectorial autoregressive analysis of soybean product prices are to be taken with caution. Secondly, due to the joint-product aspect of meal and oil production, the long-run effect of a rise in the price for soybean meal (oil) is a decrease in the price for soybean oil (meal) which is strictly proportional to the respective oil and meal contents of soybeans. Although this result has already been theoretically discussed, it was never given an empirical basis. Lastly, Rotterdam markets for soybean products do not seem to behave as efficient markets in the sense of FAMA.
REFERENCES


