Agricultural Markets Integration in the European Union: Further Empirical Evidence on the Pork Sector

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Abstract

Market integration among the EU pork markets is studied from both a long and short-run perspective. Johansen's/1988/ multivariate cointegration procedure is used to identify long-run equilibrium relationships among pork prices. Some hypotheses about perfect integration are formulated and tested. Forecast Error Variance decomposition is used to examine the short-run interrelationships among price series. Pig carcasses (grade II) prices from EUROSTAT are used and five countries are considered: Denmark, Spain, Germany, United Kingdom and Italy. Data cover the period from January 1973 to December 1993. Structural breaks have been considered when testing for the presence of unit roots. Results suggest that a high degree of integration exists among these selected markets although the detection of only one cointegrating vector does not provide evidence enough in favor of a unique pork market in the EU. Pro-

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duction characteristics, deficiencies in information transmission and sanitary controls may explain sluggishness in prices adjustments. (JEL Classification: C32, F15, Q11) <Key Words: market integration; unit roots; multivariate cointegration; forecast error variance decomposition.>

I. Introduction

There are at least two factors that should favor strong integration relationships among agricultural markets in the European Union. First, the removal of trade barriers with the aim of achieving the Unified European Market by the beginning of 1993 and, second, the Common Agricultural Policy (CAP) and, in particular, the common institutional prices regime. However, in the last twenty years the effects of these two elements have been reduced by non-tariff barriers, such as sanitary controls and the Monetary Compensatory Amounts (MCAs), which might have limited the free flow of agricultural commodities.

Assuming homogeneous goods, perfect information, and no barriers to trade, the Law of One Price (LOP) postulates that international arbitrage should force differences in prices across locations, expressed in a common currency, to be no greater than transportation costs. Markets in which this condition holds are said to be integrated. The study of spatial price relationships contributes to explain the global performance of markets (Goodwin and Schroeder [1991b]), giving some clues about the effectiveness of arbitrage (Carter and Hamilton [1989]) and the efficiency of pricing (Buccola [1985]). In empirical studies, it is common to define market integration on the basis of price transmission. Two markets are integrated if price changes in one market are fully passed on the other markets (Faminow and Benson [1990]). However, this adjustment may take time, so market integration or the fulfillment of the LOP can be viewed as an equilibrium relationship that is satisfied in the long-run.

The purpose of this paper is to examine the degree of integration achieved by pork markets in the EU in the period 1973-1993. This industry represents about the 11% of the Final Agricultural Output. However, its relative importance is higher if narrow linkages with the cereals sector are taken into consideration. The intense trade flows within the European markets, the common institutional prices, the simultaneous operations developed by some large multinational feed companies in different countries as
well as vertical and horizontal integration practices, may have contributed to homogenize pork prices in Europe.

The explicit consideration of prices non-stationarity has led to widespread use of cointegration. Cointegrated prices do not drift apart in the long-run and tend to move towards a shared equilibrium path. Most of the existing literature has dealt with pairs of prices (Ardeni [1989], Schroeder and Goodwin [1990], Baffes [1991], Zanias [1993]). However, a multivariate framework using the Johansen's [1988] procedure is more adequate as it allows the researcher to consider all possible price linkages. This can be examined by Johansen's [1988] procedure. Goodwin [1992] used the same methodology to test the LOP in international wheat markets. He argues that a larger number of cointegrating vectors provides stronger support for market integration but no formal tests are carried out. In this paper market integration is formally tested by performing similar tests to those used in Ravallion [1986]. On the other hand, we have considered that market integration has to do not only with long-run trends but also with short-run dynamics. Short-run dynamic linkages are studied using the Forecast Error Variance (FEV) decomposition obtained from a Vector Error Correction Model (VECM) where the restriction on the number of cointegration vectors is imposed.

The paper is organized as follows. In section II some comments about spatial market integration and a review of the most recent empirical approaches are presented. In section III the econometric techniques used in this paper are exposed. In section IV univariate properties of price series are analyzed. The long period covered in the study suggests the convenience of performing unit roots tests taking into account, explicitly, structural change. In section V results of the empirical application are presented. Finally, some concluding remarks are outlined.

II. Empirical Approaches to Market Integration Testing

Tests of the LOP have often used a model similar to the following (see for example Richardson [1978]):
\[ P_{1,t} = \beta_0 + \beta_1 P_{2,t} + T_{t}^{\beta_2} \] (1)

where \( P_{1,t} \) and \( P_{2,t} \) are prices for a homogeneous good in countries 1 and 2, in time \( t \) and \( T_{t} \) represents the commodity exchange transfer costs between both countries. Estimation is carried out by taking the log transformation of prices such that equation (1) becomes linear.

The absolute version of the LOP requires price levels equalization, that is \( \beta_0 \), \( \beta_1 \) and \( \beta_2 \) have to be not significantly different from one. This version is very restrictive and rather difficult to be satisfied empirically as data on transfer costs are rarely available. Normally, this variable is assumed to be constant and, hence, its influence on spatial price relationships is reflected in the constant term in the log equation. The relative formulation only requires proportionality between both prices (\( \beta_2=1 \) or the elasticity of transmission equals one). In this case, markets are said to be perfectly integrated. Integrated markets only require the \( \beta_2 \) parameter to be significantly different from zero.

Alternative methodological approaches have been used to test market integration which have evolved in time as new econometric developments became available. The static regression between prices has been broadly criticized and rapidly replaced by dynamic approaches. Even if markets are efficient, spatial arbitrage may not be instantaneous. Time is required to recognize that prices differential offers an opportunity for trading and for physical deliver of goods (Schroeder and Goodwin [1990]). Likewise, some commodities are storable and sold under medium or long-term contracts (Delpachitra and Hill [1994]). Granger causality tests (Blank [1987]), Gordon et al. [1993], Bellégo [1992]), the Ravallion’s [1986] approach and Vector Autoregressive models (VAR) (Shroeder and Goodwin [1990], Goodwin and Schroeder [1991a]) have been introduced to recognize the dynamic dimension of market integration.

Ardeni’s [1989] paper is the first one in testing market integration within the context of cointegration. Non-stationarity of prices invalidate testing procedures among levels, whilst differencing may lose valuable long-run information contained in the data. Two series are cointegrated if, being individually I(1) (the most frequent case), there is a linear combination of them that is stationary. Stationarity can be viewed as the statistical counterpart of the economic concept of equilibrium. If two prices are cointegrated, they are
linked by a long-run relationship that precludes them to drift apart and, therefore, their respective markets are integrated\(^1\).

Bivariate cointegration has been profusely used to test market integration among international agricultural markets (e.g. Ardeni [1989], Baffes [1991], Goodwin and Schroeder [1991b], Zanias [1993]). As in the static approach, a slope coefficient equal to one is required for perfect integration in the long-run (Baffes [1991], Goodwin and Schroede [1991b]).

However, the multilateral patterns of trade are likely to draw complex interactions and simultaneous determination of market prices which can be better described in a multivariate framework, following the procedure developed by Johansen [1988]. Goodwin [1992] and Silvapulle and Jayasuriya [1994] inter alia, have applied this method to evaluate agricultural market integration. However, further effort has to be made to develop hypotheses on the long-run parameters and to derive economic implications from them. On the other hand, rejection of cointegration could be due to non-stationary transaction costs, and not to non-integrated markets (Barrett [1996]).

Moreover, long-run price transmission can be extended to evaluate patterns of dynamic linkages. Market integration will be accompanied with a greater interdependence among prices in the short-run, such that every price contributes to explain the evolution of the others. This can be analyzed with the FEV decomposition. In this approach, cointegration becomes an essential step for properly specifying the VECM to be used in the analysis of dynamics (see Goodwin et al. [1996]). Both issues, hypotheses testing on the long-run and short-run dynamics will be considered in this paper.

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1. Tests for cointegration are sensible for the purpose of studying market integration when prices are integrated of the same order. If price series are stationary, then equilibrium relationships may be established by means of traditional estimation procedures. Baffes ([1991] p.1265) points out that: ".. if prices are integrated of different order, no equilibrium exist and the LOP does not hold, because one price contains explosive components which cannot be explained by the other price." In our opinion, this fact does not preclude the existence of integration among markets, but just suggest differences on their functioning and characteristics.
III. Econometric Framework

The multivariate maximum likelihood procedure proposed by Johansen [1988] starts by specifying the following VAR model:

$$Y_t = \Psi D_t + A_1 Y_{t-1} + ... + A_{t-p} Y_{t-p} + \epsilon_t$$  (2)

which can be reparameterized in a Vector Error Correction form (VECM) as follows:

$$\Delta Y_t = \Psi D_t + \Gamma_1 \Delta Y_{t-1} + ... + \Gamma_{p-1} \Delta Y_{t-p+1} + \Pi Y_{t-1} + \epsilon_t$$  (3)

where $Y_t$ is a vector of $k$ variables (e.g. price series for each market); $D_t$ are deterministic variables (e.g. seasonal dummies); $\Gamma_i$ is a $k \times k$ matrix of short-run parameters ($i=1,\ldots,p$) where $p$ is the number of lags. The matrix $\Pi$ of order $k \times k$ contains the information about the long-run relationships among the series in $Y_t$. If $\Pi$ is of full rank ($r=k$), then $y_t$ is a vector of stationary variables, while a rank of zero implies that contains no long-run information. Finally, if the rank of $\Pi$ is a positive number $r<k$, then there are $r$ stationary linear combinations of the five prices in $y_t$ (i.e. series are cointegrated). Johansen [1988] develops two statistics, $\lambda_{\text{max}}$ and $\lambda_{\text{trace}}$ to test for the cointegration rank ($r$).

It is useful to define two $k \times r$ matrices $\alpha$ and $\beta$ such that $\Pi=\alpha\beta'$, where the columns of $\beta$ contain the parameters of the $r$ cointegrating vectors; and $\alpha$ is a matrix of adjustment parameters which measure the average speed of convergence towards the long-run equilibrium after a shock. The maximum likelihood procedure developed by Johansen [1988] allows to estimate those matrices. Hypothesis testing on the long-run parameters, $\alpha$ and $\beta$, is developed by Johansen and Juselius [1990,1992].

Perfect market integration in the long-run has usually been expressed in two different ways. The first one concludes that markets are perfectly integrated when, as a result of international arbitrage activities, it is possible to find a unique representative price for the set of markets. In this sense, Goodwin ([1992] p.337) point out that: “multiple cointegrating vectors provide even stronger support for the concept of a single price. ...a single cointegrating vector (as obtained in this study) implies that any single price may be solved for in terms of the $k-1$ prices,... and then is not fully representative
of the set of \( k \) prices”. Hence, a single cointegrating vector implies that markets are integrated, as far as a long-run equilibrium exist among prices, but preclude perfect integration. Nevertheless, any linear combination of cointegrating vectors is also an equilibrium relationship, and this makes economic interpretation extremely complex.

The second expression defines perfectly integrated markets when price changes in one market are reflected by identical changes in the prices of the other markets. In this sense, when the equilibrium relationship is formulated between two markets (see equation (1)) it requires \( \beta_j = 1 \). In the multivariate context, the long-run equilibrium can be expressed as follows:

\[
W_t = B'Y_t = \beta_0 + \sum_{i=1}^{i=k} \beta_i P_{i,t} \tag{4}
\]

and the equivalent hypothesis becomes:

\[
H_0 : \sum_{i=1}^{i=k} \beta_i = 0 \tag{5}
\]

Moreover, in the estimated cointegrating vector, some implicit linkages among pairs of prices can be detected. In this way, some pairs of parameters \( \beta_i \) may have values proportionate to [1,-1]. That is to say, even if only one cointegrating vector exists, stronger linkages among subgroups of prices can be obtained. The hypothesis of perfect integration between markets “\( i \)” and “\( j \)” is formulated as²:

\[
H_0 : \beta_i = -\beta_j \quad (i \neq j) \quad (H_0 : \beta_i + \beta_j = 0) \tag{6}
\]

---

2. If we are interested in testing if markets “\( i \)” and “\( j \)” are perfectly integrated, then a change in \( P_i \) will be reflected by identical change in \( P_j \). This condition can be expressed analytically as follows:

\[
\frac{\partial P_j}{\partial P_i} \frac{P_j}{P_i} = \frac{-\beta_j}{\beta_i} = 1
\]

Hence, perfect integration exists among both markets when \(-\beta_i = \beta_j \) \( (\beta_i + \beta_j = 0) \).
Taking into consideration all what has been mentioned above, it is easy to understand that the existence of a single cointegrating vector can not guarantee perfect integration as it is not possible that only one vector contains all possible bilateral price relationships conditions of perfect integration. Goodwin [1992] suggests that perfect integration requires the number of cointegration vectors equals the number of prices minus one \((r = k-1)\). However, we consider this requirement is not sufficient. The concept of perfectly integrated markets imposes restrictions on \(\beta\) parameters, such that perfect transmission between a pair of prices is identified in each vector. In this way, the fulfillment of condition (5) can be considered as a favorable symptom of market integration but not an absolute test of perfect integration in the long-run.

Nevertheless, some authors (Goodwin et al. [1996]) note that multivariate cointegration tests may lack the power to reveal a large number of cointegrating vectors. Thus, additional tests on the long-run parameters and the study of short-run interdependences will allow us to gain insight into prices interrelationships.

Short-run dynamic linkages among prices are studied from the FEV decomposition\(^3\). This is one of the tools that summarize the information contained in a VAR model. The FEV is obtained from the conversion of the VAR into its equivalent moving-average representation. When there is cointegration, a VAR in levels is misspecified. For this reason, first, the VECM estimated in (3), with the number of cointegration vectors imposed, has to be transformed into its equivalent VAR in levels. The matrices \(A_i\) in (3) are calculated from the \(\Pi\) and \(\Gamma\) matrices in (3):

\[
\begin{align*}
A_1 &= I_k + \Gamma_1 - \Pi \\
A_i &= \Gamma_i - \Gamma_{i-1} & i=2,\ldots, P-1 \\
A_p &= -\Gamma_{p-1}
\end{align*}
\]

The \(h\)-step ahead Forecast Error Variance (FEV) is decomposed into contributions of every variable's innovation in the system. Analysis of FEV provides information about the strength of interrelationships among the vari-

\(^3\) A detailed explanation about FEV calculation can be found in Lutkepohl [1993]
ables. Large proportions attributed to one variable's own innovation indicate that this variable is primarily influenced by its own past structure with limited interaction with the others.

IV. Data

A. Source and Preliminary Analysis

Monthly pig carcass (grade II) prices are obtained from EUROSTAT: "Agricultural Prices". Data cover the period January 1973-December 1993. All prices are measured in European Currency Units (ECU) and transformed into logs. Five countries are considered: Denmark ($P_{Dn}$), Spain ($P_{Sp}$), Germany ($P_{Ge}$), United Kingdom ($P_{UK}$) and Italy ($P_{It}$). The choice is based on data availability and the relative importance of these countries in European pork production and intra EU-trade. The five countries amount to 63% of the total meat production in the EU. Prices series in logs are shown in Figure 1.

Prices evolution is very similar among the five countries considered. However, it is noticeable that more similarities are found, on the one hand, between Germany, United Kingdom and Denmark, and on the other hand, between Spain and Italy. The first three series follow a clear growing trend until 1985, approximately, falling down from that date. This change in trend is less sharp in the two other series, where only a slight change in the slope magnitude is observed.

This break can be explained by the coincidence of at least three factors. First, in 1984 the first CAP reform was performed. Intervention prices reduction and surpluses control measures, such as milk quotas, were approved. This pushed away some milk exploitations, leading to an increase in slaughtering of dairy cows, as well as their transformation into pig farms, with the subsequent increase of pork supply. Second, cereals prices fell in 1984 as a consequence of production surplus and the application of new intervention prices, lower than the preceding ones. Feeding is the main input used in pig production, representing about 75% of total production costs. Lower costs encouraged pig production. Taking into account biological lags, this decision is reflected in the market around 9-11 months later, with an increase in pork
Figure 1

[Graphs showing price series evolution for different countries, including Spain, Italy, and other unspecified countries.]

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supply for slaughtering what could explain falling prices in 1985. Third, by that time, dollar depreciated with respect to the ECU, making the cost of animal feeding cereal substitutes (gluten feed) cheaper, reinforcing the incentive to expand production, while it made exports from the EU, expressed in dollars, more difficult. As a consequence, pig surpluses were generated, pressing prices down.

As far as Spain is concerned, the effects of the accession into the EU in 1986 had two immediate effects: 1) the instantaneous application of the common base price, lower than that prevailing in Spain; and 2) the liberalization of imports coming from the EU, while exports were forbidden due to sanitary regulations. Piglets contingent measures, aids to private stockage and the upward price cycle offset the downward pressure on prices (Buxade [1988] p.182). In 1988, the most relevant producer regions in Spain were declared free of swine fever allowing live animals and meat exports to the EU since then. By this time, prices in Europe had fallen as a result of the herd expansion which pushed domestic prices down.

B. Unit Root Tests

Before applying the methodology exposed in Section III, the non-stationarity of the prices series has to be ascertained. Visual inspection suggests that the prices have suffered a structural change. The distribution of the most common unit roots tests (Dickey-Fuller (DF) and Phillips-Perron (PP)) are altered by the presence of level or trend breaks in the data (Perron [1989]; Rappoport and Reichlin [1989]). Hence, the explicit recognition of structural breaks becomes necessary. The general test equation for the presence of unit roots, corresponding to the mixed break model (change in level and trend), is the following:

---

4. The univariate properties analysis of series began by studying the stochastic seasonal component because of monthly periodicity of data. Franses [1991] method for testing unit roots in seasonal frequencies was used. The null of I(2,1) was tested against the alternatives I(1,1), I(2,0) and I(1,0) for all series. Results indicated that all series were I(1,0) and seasonal dummies were statistically significant. Therefore, seasonality was stationary and seasonal dummies were used to capture this component. Results are available from the authors.
\[ \Delta y_t = \mu + \beta t + \alpha y_{t-1} + d_1 D_t + d_2 D_t + \sum_{i=1}^{p} \gamma_i \Delta y_{t-i} + \varepsilon_t \]  

(8)

where \( t \) is a trend; \( D_t \) is a dummy variable that captures the level shift; \( D_t \) is equal to 1 if \( t > T_B \) and 0 otherwise, where \( T_B \) is the observation in which the break takes place; \( D_t \) is a dummy variable that captures the trend break; \( D_t \) equals \( t - T_B \) if \( t > T_B \) and 0 otherwise.

Following the recommendation by Christiano[1992], the break point has been estimated endogeneously. In particular, the sequential testing strategy has been chosen. Attending to the evolution of the price series in Figure 1, two types of breaks have been considered: trend break (\( D_t = 0 \) in [8]) and mixing break, which allows for a change both in the intercept and in the trend (\( D_t \) and \( D_t \) different from zero). For each \( T_B \) and for each type of break, the \( t \)-ratio on \( \alpha \) is computed. The minimum \( t \)-ratio (\( t_{DF}^{min} \)) indicates that the corresponding \( T_B \) is the estimated break point. \( t_{DF}^{min} \) is then compared with the critical values tabulated in Banerjee et al.[1992] and Montañés[1996]. If \( t_{DF}^{min} \) is greater (in absolute value) than the critical value, then the unit root null hypothesis is rejected in favor of stationarity around a segmented trend (mean). If \( t_{DF}^{min} \) is lower then the unit root null hypothesis can not be rejected. In the trend break model these authors calculate the statistic \( F \) that tests the joint hypothesis \( \alpha = 0 \) and \( d_2 = 0 \) (unit root and significance of trend shift). The maximum \( F \) obtained (\( F_{T}^{max} \)) selects the break point.

The number of lags (\( p \)) included in model (8) was determined by choosing those significative lags that simultaneously canceled out autocorrelation on residuals. Results are showed in Table 1. Both models support the presence of a unit root in all series. Likewise, it is noticeable that the break points selected in the shock model by the statistics \( t_{DF}^{min} \) and \( F_{T}^{max} \) coincide and are consistent with the preliminary analysis.

Summing up, the unit roots tests that fit better the characteristics of prices fail to reject that all the series are integrated of order 1, making necessary the study of cointegration as a preliminary technique to evaluate

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5. Critical values corresponding to the break trend and the mixing break model are in Banerjee et al.[1992] and Montañés[1996], respectively.
Table 1

Results from Sequential Testing for Unit Roots with Structural Change

<table>
<thead>
<tr>
<th>Series</th>
<th>Trend Break</th>
<th>Mixed Break</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>$T_B^{(a)}$</td>
<td>$t_{10}^{(b)}$</td>
</tr>
<tr>
<td>$P_{De}$</td>
<td>Aug. 84</td>
<td>-3.59</td>
</tr>
<tr>
<td>$P_{Sp}$</td>
<td>Dec. 88</td>
<td>-3.98</td>
</tr>
<tr>
<td>$P_{Ger}$</td>
<td>Aug. 84</td>
<td>-3.76</td>
</tr>
<tr>
<td>$P_{UK}$</td>
<td>Aug. 84</td>
<td>-2.25</td>
</tr>
<tr>
<td>$P_{h}$</td>
<td>Aug. 84</td>
<td>-4.17</td>
</tr>
</tbody>
</table>

$^{(a)}$ $T_B$ is the time of break.

$^{(b)}$ Minimum $t$-ratio on $z$ in equation [8]. Critical value for $n=250$ is: -4.39 at the 5% level of significance (Banerjee et al., 1992)

$^{(c)}$ Tests the joint hypothesis $\alpha=0$, $d=0$ in [8]. Critical value for $n=250$ is: 15.94 at the 5% level of significance (Banerjee et al., 1992)

$^{(d)}$ Critical value for $n=250$ is: -5.03 at the 5% level of significance (Montañés, 1996)

market integration in the long-run and to correctly specify the VECM used in the subsequent analysis.

V. Results

A. The VECM Formulation

The first task before testing for cointegration is to appropriately specify the VECM in which such tests are undertaken. More specifically, it is necessary to determine the deterministic components in (2) as well as the optimum lag ($q$). The Akaike Information Criteria test selects a lag order of 20 months$^6$ for each variable. Each equation includes eleven seasonal dummies to account for deterministic and stationary seasonality. A constant is allowed in the cointegration space to consider possible differences among prices in the long-run explained by (stationary) transaction costs. Diagnostic tests for

---

$^6$ This lag coincides with the time needed for herd expansion decision is reflected in market by an increase in supply (3.8 months since the gilt is bred till the pig is born, plus 7-8 months needed for the new gilts to be bred, and an additional period of 10 months is required to have fed pigs ready for slaughter).
Table 2
Diagnostic Tests

<table>
<thead>
<tr>
<th>Order</th>
<th>$P_{Den}$</th>
<th>$P_{Sp}$</th>
<th>$P_{Ger}$</th>
<th>$P_{UK}$</th>
<th>$P_{It}$</th>
<th>$P_{Den}$</th>
<th>$P_{Sp}$</th>
<th>$P_{Ger}$</th>
<th>$P_{UK}$</th>
<th>$P_{It}$</th>
</tr>
</thead>
<tbody>
<tr>
<td>1</td>
<td>0.01</td>
<td>1.59</td>
<td>0.02</td>
<td>0.98</td>
<td>0.07</td>
<td>1.20</td>
<td>0.08</td>
<td>0.00</td>
<td>0.00</td>
<td>0.00</td>
</tr>
<tr>
<td>6</td>
<td>1.90</td>
<td>9.48</td>
<td>3.30</td>
<td>8.40</td>
<td>10.12</td>
<td>3.92</td>
<td>4.91</td>
<td>0.81</td>
<td>2.10</td>
<td>0.96</td>
</tr>
<tr>
<td>12</td>
<td>5.34</td>
<td>13.84</td>
<td>6.00</td>
<td>10.56</td>
<td>21.28</td>
<td>6.40</td>
<td>7.43</td>
<td>11.76</td>
<td>8.84</td>
<td>3.85</td>
</tr>
<tr>
<td>18</td>
<td>19.94</td>
<td>26.75</td>
<td>9.00</td>
<td>12.33</td>
<td>22.81</td>
<td>7.36</td>
<td>10.25</td>
<td>14.30</td>
<td>12.02</td>
<td>5.67</td>
</tr>
</tbody>
</table>

* Both statistics are distributed as a $x^2$ where the number of degrees of freedom correspond to the autocorrelation order tested.

Table 3
Tests on the Cointegration Rank

<table>
<thead>
<tr>
<th>Ho: r</th>
<th>Eigen-Value</th>
<th>Max</th>
<th>CV 10% *</th>
<th>Trace</th>
<th>CV 10%</th>
</tr>
</thead>
<tbody>
<tr>
<td>0</td>
<td>0.1951</td>
<td>50.13*</td>
<td>31.66</td>
<td>96.64 *</td>
<td>71.86</td>
</tr>
<tr>
<td>1</td>
<td>0.0829</td>
<td>20.00</td>
<td>25.56</td>
<td>46.51</td>
<td>49.65</td>
</tr>
<tr>
<td>2</td>
<td>0.0640</td>
<td>15.28</td>
<td>19.77</td>
<td>26.51</td>
<td>32.00</td>
</tr>
<tr>
<td>3</td>
<td>0.0264</td>
<td>6.17</td>
<td>13.75</td>
<td>11.23</td>
<td>17.85</td>
</tr>
<tr>
<td>4</td>
<td>0.0216</td>
<td>5.06</td>
<td>7.52</td>
<td>5.06</td>
<td>7.52</td>
</tr>
</tbody>
</table>

*CV = Critical Value at 10% significance level. Critical values are taken from Osterwald-Lenum [1992]

Table 4
Tests on Perfect Integration among Markets

<table>
<thead>
<tr>
<th>Ho: $\sum_{i=1}^{k} \hat{\beta}<em>{Ger} - \hat{\beta}</em>{UK} - \hat{\beta}_{Den}$</th>
<th>LR test</th>
<th>LR* test</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\sum_{i=1}^{k} \hat{\beta}<em>{Ger} - \hat{\beta}</em>{UK} - \hat{\beta}_{Den}$</td>
<td>1.09</td>
<td>33.7*</td>
</tr>
<tr>
<td>$\sum_{i=1}^{k} \hat{\beta}<em>{Ger} - \hat{\beta}</em>{UK} - \hat{\beta}_{Den}$</td>
<td>20.3*</td>
<td>38.7*</td>
</tr>
<tr>
<td>$\sum_{i=1}^{k} \hat{\beta}<em>{Ger} - \hat{\beta}</em>{UK} - \hat{\beta}_{Den}$</td>
<td>11.5*</td>
<td>24.7*</td>
</tr>
<tr>
<td>$\sum_{i=1}^{k} \hat{\beta}<em>{Ger} - \hat{\beta}</em>{UK} - \hat{\beta}_{Den}$</td>
<td>21.2*</td>
<td>32.6*</td>
</tr>
<tr>
<td>$\sum_{i=1}^{k} \hat{\beta}<em>{Ger} - \hat{\beta}</em>{UK} - \hat{\beta}_{Den}$</td>
<td>6.1*</td>
<td>40.0*</td>
</tr>
<tr>
<td>$\sum_{i=1}^{k} \hat{\beta}<em>{Ger} - \hat{\beta}</em>{UK} - \hat{\beta}_{Den}$</td>
<td>29.6*</td>
<td>40.1*</td>
</tr>
<tr>
<td>$\sum_{i=1}^{k} \hat{\beta}<em>{Ger} - \hat{\beta}</em>{UK} - \hat{\beta}_{Den}$</td>
<td>5.3*</td>
<td>37.7*</td>
</tr>
<tr>
<td>$\sum_{i=1}^{k} \hat{\beta}<em>{Ger} - \hat{\beta}</em>{UK} - \hat{\beta}_{Den}$</td>
<td>2.8</td>
<td>38.2*</td>
</tr>
<tr>
<td>$\sum_{i=1}^{k} \hat{\beta}<em>{Ger} - \hat{\beta}</em>{UK} - \hat{\beta}_{Den}$</td>
<td>23.4*</td>
<td>41.1*</td>
</tr>
<tr>
<td>$\sum_{i=1}^{k} \hat{\beta}<em>{Ger} - \hat{\beta}</em>{UK} - \hat{\beta}_{Den}$</td>
<td>3.3</td>
<td>35.9*</td>
</tr>
</tbody>
</table>

*LR- Likelihood Ratio test. An asterisk indicates rejection of the null hypothesis at the 5%

*No restrictions imposed on the other $\beta$ parameters. The critical value at 5% is $\chi^2_1 = 3.84$

*The other $\beta_i$ parameters restricted to zero. The critical value at 5% is $\chi^2_2 = 9.49$
autocorrelation (Ljung-Box) and hereroskedasticity (ARCH) have been performed. Results are presented in Table 2. Both statistics support the adequate specification of the model.

The two likelihood ratio tests, $\lambda_{max}$ and $\lambda_{trace}$, are used to test for the number of cointegrating vectors. The first one tests the null hypothesis that there are $r$ cointegrating vectors. The second statistic tests the null hypothesis that there are at most $r$ cointegrating vectors. Results are shown in Table 3.

Both statistics suggest that only one cointegrating vector exists among pork prices of the five European countries considered. As stated before, this result supports that markets are integrated but precludes perfect integration in the long-run.

**B. Tests on $\beta$ Parameters**

Results of tests on $\beta$ parameters are displayed in Table 4.

The first hypothesis tested ($\Sigma \beta_i = 0$) can not be rejected at the 5% significance level. This is a favorable indicator on long-run integration among the five markets considered. Analyzing pairwise relationships without imposing any restriction on the non tested parameters, we fail to reject the null of perfect integration between the British and Danish markets, and between the Danish and the Spanish ones. Both relationships could be explained in terms of trade intensity, specially between UK and Denmark. Denmark is the main supplier of UK (in 1990, 48% of British imports came from that country). The relative importance and persistency of such trade flows would have developed tight links among prices, at least in the long-run. Likewise, Denmark is an important provider of Spain. However, Spanish imports coming from that country are not as large as those coming from Germany, France or Italy. It seems logical to think that price transmission between both markets takes place through an intermediary and common trade partner, such as France. However, none of these parity linkages are equilibrium relationships without considering the other prices in the system. That is to say, perfect integration between these pairs of markets only constitutes a stable equilibrium solution within the system defined by the five European prices.
C. Weak Exogeneity Tests

The statistical significance of the $\alpha_i$ parameters means that variable $P_i$ (dependent variable in the $i$-th equation) is not weakly exogeneous with respect to the long-run equilibrium. This implies that, when a transitory disequilibrium occurs ($W_i \neq 0$ in (4)), $P_i$ will react to get the system back to the equilibrium situation. Results from individual tests on $\alpha_i$ parameters are shown in Table 5.

Both German and British prices are weakly exogeneous with respect to the long-run ($\alpha_{Ger}=0$, $\alpha_{UK}=0$). Likewise, the joint null $\alpha_{Ger}=0$, $\alpha_{UK}=0$ can not be rejected. This result does not vary when introducing restrictions on parameters. This result could imply a greater degree of independence in the pricing determination process. However, in order to state if these markets are acting as leaders, more information about the short-run dynamics is required.

D. Short-run Dynamics

The VECM with one cointegrating vector has been transformed into a VAR in levels following the equivalence between parameters showed in (7). The Choleski decomposition has been used to transform the covariance matrix of innovations to an identity matrix which depends on the way variables are ordered. From more to less exogeneous, the prices are ordered as: $P_{Den}$, $P_{Sp}$, $P_{Gen}$, $P_{UK}$, $P_{It}$. This ordering is based upon the degree of self-sufficiency. In this way, the position of greater exogeneity ($P_{Den}$) is assigned to the country with greater coefficient. FEV decomposition for alternative forecast horizons, ranging from 1 to 24 months, are displayed in Table 6.

Danish and German prices can be considered as the most exogeneous in the system. The percentage of these prices' FEV attributed to their own error exceeds 80 and 60 %, respectively, at all reported horizons. Exogeneity involves a relatively greater degree of autonomy in the price determination process, but not independence. Both prices are explained to some extent by innovations in the other prices of the system, although only their reciprocal

---

7. Other orderings were used but did not alter significantly the results.
Table 5
Tests on Weak Exogeneity

<table>
<thead>
<tr>
<th>$H_0$</th>
<th>No Restrictions on $\beta$</th>
<th>$\beta_{uk} = -\beta_{dnk}$</th>
<th>$\beta_{dnk} = -\beta_{sp}$</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>LR</td>
<td>CV 5%</td>
<td>LR</td>
</tr>
<tr>
<td>$\alpha_{dnk} = 0$</td>
<td>10.22 *</td>
<td>$x_1^2 - 3.84$</td>
<td>10.91 *</td>
</tr>
<tr>
<td>$\alpha_{sp} = 0$</td>
<td>11.70 *</td>
<td>14.32 *</td>
<td>20.60</td>
</tr>
<tr>
<td>$\alpha_{dr} = 0$</td>
<td>2.07</td>
<td>5.65</td>
<td>9.48 *</td>
</tr>
<tr>
<td>$\alpha_{uk} = 0$</td>
<td>0.04</td>
<td>2.84</td>
<td>3.68</td>
</tr>
<tr>
<td>$\alpha_{dr} = \alpha_{uk} = 0$</td>
<td>6.05 *</td>
<td>9.48 *</td>
<td>7.60 *</td>
</tr>
<tr>
<td>$\alpha_{dr} = \alpha_{uk} = 0$</td>
<td>2.14</td>
<td>6.38</td>
<td>5.85</td>
</tr>
</tbody>
</table>

*LR = Likelihood Ratio test. An asterisk indicates rejection of the null hypothesis at the 5% significance level

CV = Critical Value

innovations account for more than 10% at some time horizon.

Conversely, the Italian price is the most endogeneous in the system. After one year, only 33% of its own variation can be attributed to its own past. Nevertheless, Italian price is not rapidly influenced by the others. Only after one year its FEV becomes mainly explained by innovations in other prices.

In general, innovations in Danish and German prices are the main factors explaining other prices FEV (apart from their own past). Danish and German markets share a common feature: they are both two of the main producer countries in the EU. Their production represents around 10 and 23%, respectively, of the total EU production. However, while Danish market is one of the largest meat pork exporters in the EU, Germany is the largest importer. Germany imports represent 42% of total pork intra EU trade while exports from Denmark represent 17%. As a result, a distinct pattern of short-run influence in European prices is observed. In shorter horizons (less than one year), the German price innovations exert a greater influence, while the Danish innovations become more important at longer horizons. This supports the leadership of Germany in pricing accordingly to other studies. Bellego[1992] characterizes Germany as a “dominant buyer”.

According to the results about perfect long-run integration mentioned above, a strong dynamic interdependence is also obtained between the Dan-
lish price, and the prices in UK and Spain. In particular, innovations in Danish price explain large proportions of British and Spanish prices' FEV after two years. It doubles the percentage accounted after one year (23 and 14% of British and Spanish FEV, respectively).

At any reported horizon, the Italian price is the least important to explain the evolution of the other prices. Only the German price is affected to some extent (about 4% after six months). Nevertheless, contributions of British and Italian prices are larger as time horizon increases.

In general, a long period is required for interdependence among prices to become strong. All the prices are relatively exogeneous during the first six months and the influence of any of them on the others increases consider-

Table 6
Forecast Error Variance Decomposition

<table>
<thead>
<tr>
<th>Variable</th>
<th>Step</th>
<th>Standard Error</th>
<th>Percent Explained by</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td></td>
<td></td>
<td>$P_{Dem}$</td>
</tr>
<tr>
<td>$P_{Dem}$</td>
<td>1</td>
<td>0.044</td>
<td>100</td>
</tr>
<tr>
<td></td>
<td>6</td>
<td>0.089</td>
<td>86.38</td>
</tr>
<tr>
<td></td>
<td>12</td>
<td>0.118</td>
<td>82.64</td>
</tr>
<tr>
<td></td>
<td>24</td>
<td>0.166</td>
<td>84.70</td>
</tr>
<tr>
<td>$P_{Sp}$</td>
<td>1</td>
<td>0.019</td>
<td>0.56</td>
</tr>
<tr>
<td></td>
<td>6</td>
<td>0.061</td>
<td>6.77</td>
</tr>
<tr>
<td></td>
<td>12</td>
<td>0.089</td>
<td>10.50</td>
</tr>
<tr>
<td></td>
<td>24</td>
<td>0.115</td>
<td>23.35</td>
</tr>
<tr>
<td>$P_{Ger}$</td>
<td>1</td>
<td>0.029</td>
<td>9.52</td>
</tr>
<tr>
<td></td>
<td>6</td>
<td>0.079</td>
<td>10.58</td>
</tr>
<tr>
<td></td>
<td>12</td>
<td>0.111</td>
<td>8.57</td>
</tr>
<tr>
<td></td>
<td>24</td>
<td>0.147</td>
<td>21.92</td>
</tr>
<tr>
<td>$P_{UK}$</td>
<td>1</td>
<td>0.036</td>
<td>4.22</td>
</tr>
<tr>
<td></td>
<td>6</td>
<td>0.079</td>
<td>3.55</td>
</tr>
<tr>
<td></td>
<td>12</td>
<td>0.104</td>
<td>5.20</td>
</tr>
<tr>
<td></td>
<td>24</td>
<td>0.130</td>
<td>14.41</td>
</tr>
<tr>
<td>$P_{It}$</td>
<td>1</td>
<td>0.024</td>
<td>0.26</td>
</tr>
<tr>
<td></td>
<td>6</td>
<td>0.055</td>
<td>4.83</td>
</tr>
<tr>
<td></td>
<td>12</td>
<td>0.080</td>
<td>9.25</td>
</tr>
<tr>
<td></td>
<td>24</td>
<td>0.114</td>
<td>23.86</td>
</tr>
</tbody>
</table>
ably in the long-run. This implies deficiencies in the price transmission process and hence, an insufficient degree of integration from a dynamic perspective.

VI. Concluding Remarks

This paper evaluates spatial market integration for porkmeat in the EU in the period 1973-1993. Five countries are considered (Denmark, Spain, Germany, United Kingdom and Italy) and monthly prices are used. Structural breaks have been found in any of the price series. The CAP reform in 1984, the stimulus to production increases generated by falling cereals prices, and the beginning of dollar depreciation are some factors that can explain the changes in trend experimented by European pork prices by 1985-86. This has been explicitly taken into account when testing for unit roots.

In order to provide a better description of market linkages, cointegration test is complemented in two ways: first, long-run integration is studied in a multivariate framework applying Johansen's procedure and some hypotheses on the long-run parameters are developed; and second, the strength of short-run interdependence among prices is analyzed with the decomposition of the forecast error variance.

Cointegration tests show evidence of long-run spatial integration. European pork prices do not move independently in the long-run. They are tied-up by trade flows generated by efficient arbitragers. However, a single cointegrating vector is not sufficient for all markets to be perfectly integrated in a unique and common European pork market. Nevertheless, a close relationship has been found between Danish and British markets, and Danish and Spanish, in the long-run. The exporting character of Denmark makes this market to hold narrow linkages, not only with its customer markets (Germany, Italy, UK, Spain) but also with its competitors (Netherland, Belgium) in order to save its market share. Both factors make Danish prices sensitive to other markets price changes.

The FEV decompostition show multidirectional interdependence among the set of prices as expected to take place in integrated markets. Linkages are stronger among countries reciprocally involved in trade accordingly to results obtained in the long-run analysis. Weak exogeneity of German prices
together with the large influence of this price in the evolution of the others indicate a relative leadership of the German market in pricing.

Nevertheless, deficiencies in the price transmission process are observed since a long-period is required before significant proportions of the own evolution are explained by the other prices' innovations. This result implies that misallocation of resources and distortion of production and distribution decisions might have occurred in this period. Therefore, a gain in global efficiency could be attained.

The existence of sanitary controls and different protection animal health policies during this period may have affected the free flow of trade. As a result, perfect integration in the long-run has not been achieved and sluggishness in price adjustments in the short-run is observed. Monetary Compensatory Amounts (MCA) have normally been considered as responsible for a low degree of integration among the European agricultural markets (see Zanias [1993]). Nevertheless, it is unlikely that the MCAs have distorted trade and integration linkages in the pork industry. On the one hand, the CAP hardly regulates this sector, implying that purchases by the intervention agencies rarely happen. This means that no incentive exists to intra-EU trade with the aim of taking advantage of institutional prices differentials when converted to national currencies, after exchange rate fluctuations have taken place (Buxade [1988] p.180). On the other hand, since 1986, MCAs are calculated using a constant proportion of intervention price (35%). The tools employed in this paper suggest imperfections in the price transmission mechanism but they are not adequate to identify the causes of these deficiencies.

Results from this study only apply to the markets and sample period considered. The extension to other agricultural products would be of interest in order to check if different production conditions and to what extent public intervention affect the degree of spatial integration. From a methodological point of view, further research is needed in two directions. The first one consists of examining the economic interpretation of more than one cointegrating vector and the formulation of perfect market integration hypotheses in this context. The second direction of research could be addressed to answer the following question: what would happen if all series are not integrated of the same order? or, if cointegration holds only within a subset of

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Trade, Technology and Labor Markets: General Equilibrium Perspectives

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Abstract

This paper summarizes the state of the debate on the effects of “globalization” and spontaneous technical change on wages and, in this context, describes the results from a recent study of the links between trade, technical change and labor market behavior. These new results show that comparatively minor gen-

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