Identification by full adjustment: evidence from the relationship between futures and spot prices

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Summary

This paper proposes a test for orthogonality of the errors in a vector error-correction model (VECM) that focuses on the recursive ordering among the contemporaneously correlated errors. The test is based on the fact that when the frequency of the data is sufficiently low one of the variables in the long-run equilibrium relationship adjusts fully within the same period to its new equilibrium level. An empirical investigation of the relationship between spot and futures prices for commodities traded on the Amsterdam Exchanges and the Chicago Board of Trade reveals that the spot price adjusts fully to its new equilibrium level if the price-discovery function of the futures market works well.

Keywords: cointegration, exogeneity, long-run causality, spot–futures price relationship, price discovery

JEL classification: C32, G10, Q13

1. Introduction

The technique of the vector error-correction model (VECM) has now been widely used in agricultural economics for constructing dynamic systems of equations for policy analysis and forecasting. In practice, however, it often turns out that the covariance matrix of the residuals in a VECM is not diagonal, implying contemporaneous correlation among the errors. Then policy analysis usually conducted by analysis of the evolution of the system caused by an innovation in just one variable may not be appropriate, as this innovation may occur at the same time as another innovation in the system. To solve this simultaneous equation problem, tests focusing on the orthogonality between innovations and conditioning variables, such as the Hausman (1978) specification test, are clearly relevant for identification and parameter estimation. Because many structural models of the errors can be proposed on subjective grounds, Swanson and Granger (1997) and Bessler and Akleman

(1998) discuss data-determined orthogonalisation methods that, in contrast to the method that we will propose in this paper, do not explore the identification restrictions that may be obtained from cointegration.

In this paper, a method is proposed for testing recursive models of the errors as revealed by cointegration and full adjustment. Full-system estimation of models involving the application of cointegrated variables, for example, the Johansen procedure (Johansen, 1995b), yield cointegrating-vector estimates that are super-consistent and free from simultaneity bias. Therefore, cointe-grating vectors can be treated as known (see, e.g. Lütkepohl, 1991: 359; Boswijk and Urbain, 1997: 33). As we will show, a known cointegrating vector allows for testing orthogonality (i.e. recursive-ordering) assumptions about the contemporaneously correlated errors in a VECM under the condition that one of the variables in the long-run equilibrium relationship immediately adjusts to changes in its optimal level. This condition is reasonable when the frequency of the data is not too high.

The empirical illustration of our test procedure concerns a comparative study of the price discovery function for two futures markets: the corn futures contract of the Chicago Board of Trade (CBOT) and the potato futures contract of the Amsterdam Exchanges (AEX).[1] The process of using all available information to formulate prices is often referred to as price discovery (Tomek, 1980). Information on how well the price discovery function is performed by commodity futures markets is essential, because these markets are widely used by firms, engaged in the production, marketing, and processing of commodities, to shift risk, facilitate equity financing, and discover prices (see, e.g. Hudson *et al.*, 1996; Yang and Leatham, 1999; Pennings and Leuthold, 2000). In addition to focusing on the long-run causal relationship (as defined by Bruneau and Jondeau (1999)) between the futures price and the spot price as a usual way to examine the price discovery function of futures markets, we show that the notion of full adjustment, as used in our test procedure, provides a completely new perspective and insight for exploring the price discovery function of futures markets.

The outline of the paper is as follows. In Section 2, we show that cointegration and full adjustment offer restrictions that allow for searching for recursive models of the errors in a VECM if the frequency of the time series is sufficiently low. A test procedure is outlined in this respect. In Section 3, we apply this test procedure to futures–spot price relationships using data from the AEX and the CBOT. In Section 4, conclusions and suggestions for further research are presented.

2. Method

Let x_t be a vector of two variables, y_t and z_t, i.e. $x_t = (y_t, z_t)'$. We assume that the T data observations are generated by a joint probability density function

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1   In 2000, the Amsterdam Exchanges merged with exchanges in Brussels and Paris to become Euronext.

as follows:

$$x_t | X_{t-1} \sim N(\mu_t, \Sigma) \tag{1}$$

where $X_{t-1} = (X_0, x_1, \ldots, x_{t-1})'$ with $X_0 = (x_{-n+1}, \ldots, x_0)$ being the fixed initial values,

$$\mu_t = E(x_t | X_{t-1}, \theta) = \sum_{i=1}^{n} \Pi_i x_{t-i}$$

$\Sigma = E[(x_t - \mu_t)(x_t - \mu_t)' | X_{t-1}, \theta]$, n is the lag-length truncation and the parameters Π_i and Σ are functions of θ. Defining $v_t = x_t - \mu_t$, where v_t is a martingale difference sequence with respect to X_{t-1}, gives the VAR(n) model

$$x_t = \sum_{i=1}^{n} \Pi_i x_{t-i} + v_t \qquad v_t \sim IN(\mathbf{0}, \Sigma). \tag{2}$$

For illustrative purposes, we consider $n = 1$ and express x_t, Π_1, v_t and Σ in terms of their scalar elements

$$y_t = \pi_{11} y_{t-1} + \pi_{12} z_{t-1} + v_{1t} \tag{3a}$$

$$z_t = \pi_{21} y_{t-1} + \pi_{22} z_{t-1} + v_{2t} \tag{3b}$$

and

$$\Sigma = \begin{pmatrix} \sigma_{11} & \sigma_{12} \\ \sigma_{21} & \sigma_{22} \end{pmatrix}.$$

Furthermore, let $x_t \sim CI(1,1)$ so that $e_t = \alpha' x_t \sim I(0)$. For the present, we assume that the cointegrating vector $\alpha = (\alpha_1, \alpha_2)'$ is known.

Model (3a) and (3b) can be rewritten in error-correction form as

$$\Delta y_t = \rho_1 e_{t-1} + v_{1t} \tag{4a}$$

$$\Delta z_t = \rho_2 e_{t-1} + v_{2t} \tag{4b}$$

where ρ_1 and ρ_2 are unknown parameters, one of which may be zero. Without loss of generality and under the assumptions made, model (4a) and (4b) can be decomposed into the conditional model for y_t given z_t:

$$\Delta y_t = \beta \Delta z_t + \lambda_1 e_{t-1} + u_{1t} \tag{5}$$

and the marginal model of z_t given by (4b), where $\lambda_1 = \rho_1 - \beta \rho_2$ and $u_{1t} = v_{1t} - \beta v_{2t}$ with $\beta = \sigma_{12}/\sigma_{22}$ to let $\text{cov}(z_t, u_{1t}) = \text{cov}(v_{2t}, u_{1t}) = 0$. Model (4a) and (4b) can also be decomposed into the conditional model for z_t given y_t:

$$\Delta z_t = \delta \Delta y_t + \lambda_2 e_{t-1} + u_{2t} \tag{6}$$

and the marginal model of y_t given by (4a), where $\lambda_2 = \rho_2 - \delta \rho_1$ and $u_{2t} = v_{2t} - \delta v_{1t}$ with $\delta = \sigma_{12}/\sigma_{11}$ to let $\text{cov}(y_t, u_{2t}) = \text{cov}(v_{1t}, u_{2t}) = 0$.

A variable is predetermined (orthogonal) in a particular equation if it is independent of the contemporaneous and future errors in that equation. Furthermore, z_t (or y_t) is weakly exogenous with respect to the long-run

parameters in (5), i.e. α and λ_1 (or (6), i.e. α and λ_2), if z_t (or y_t) does not display error-correcting behaviour, i.e. $\rho_2 = 0$ in (4b) (or $\rho_1 = 0$ in (4a)). In addition, when z_t (or y_t) is weakly exogenous with respect to the long-run parameters in (5) (or (6)), then z_t (or y_t) is also weakly exogenous with respect to the short-run parameters in (5) (or (6)), i.e. β and $\mathrm{var}(u_{1t})$ (or δ and $\mathrm{var}(u_{2t})$), if Δz_t (or Δy_t) is predetermined in (5) (or (6)) (see, e.g. Boswijk and Urbain, 1997). It should be noted, however, that Δz_t is predetermined in (5) by construction and so is Δy_t in (6). Consequently, the concept of predeterminedness does not provide a testable hypothesis regarding the exogeneity of Δy_t and Δz_t in the conditional model (see Engle *et al.* (1983) and, among others, Ericsson (1992), Johansen (1992a, 1992c) and Hendry (1995b)).

Nevertheless, let us suppose that y_t and z_t are in equilibrium, i.e. $e_t = 0$, and a shock v_{2t} occurs. According to (4b), $\Delta z_t = v_{2t}$, and provided that the disturbance term of the relationship between v_{1t} and v_{2t} is equal to zero, i.e. $u_{1t} = 0$, then (5) shows that $\Delta y_t = \beta \Delta z_t$. If $\beta = -\alpha_2/\alpha_1$, then the v_{2t} shock will not bring y and z out of equilibrium. Consequently, the restriction $\beta = -\alpha_2/\alpha_1$ implies that y_t adjusts immediately, within the same time period, to its new equilibrium level. Adjustment costs (Lucas, 1967), for example, may prohibit y_t from adjusting immediately; gradual adjustment requires β to be smaller than $|-\alpha_2/\alpha_1|$. However, if $\beta = -\alpha_2/\alpha_1$, deviations from equilibrium are generated by u_{1t} shocks—being uncorrelated with Δz_t—and the error-correction term in (5) ensures that y and z are brought back into equilibrium.

Similarly, z_t may respond to a change in y_t without violating their equilibrium. This requires the restriction $\delta = -\alpha_1/\alpha_2$. Deviations from equilibrium originate from u_{2t} shocks and gradually disappear because of the error-correction term in (6).

If a VAR of order two or higher is selected, it is possible that, in spite of the immediate adjustment in the beginning, there is a subsequent deviation from the equilibrium, unless the parameters of the lagged y and z variables comply with restrictions that transform these variables into lagged e variables (i.e. an AR(n) model with $n \geq 1$ applies to e leading to $n - 1$ COMFAC restrictions in the short-run dynamics; see Hendry, 1995a). Immediate adjustment followed by a transitory deviation from the equilibrium occurs, for example, when an agent must meet his or her contractual commitments first, before he or she can make adjustments that will ultimately lead to a new equilibrium between y and z.

Our test for orthogonality in case of immediate adjustment consists in testing the restriction $\beta = -\alpha_2/\alpha_1$ in (5) and testing $\delta = -\alpha_2/\alpha_1$ in (6). This is done by performing a t-test of $\varphi_1 = 0$ in the OLS regression

$$\Delta e_t/\alpha_1 = \varphi_1 \Delta z_t + \lambda_1 e_{t-1} + u_{1t} \tag{7}$$

and by performing a t-test of $\varphi_2 = 0$ in the OLS regression

$$\Delta e_t/\alpha_2 = \varphi_2 \Delta y_t + \lambda_2 e_{t-1} + u_{2t}. \tag{8}$$

Two of the possible outcomes are of interest for the recursive ordering of the VECM residuals. First, if $\varphi_1 = 0$ is not rejected whereas $\varphi_2 = 0$ is rejected, then, conditional on the assumption of immediate adjustment, it is concluded that Δz_t is orthogonal. Second, if $\varphi_1 = 0$ is rejected and $\varphi_2 = 0$ is not rejected, then orthogonality of Δy_t is supported by the data, again provided there is immediate full adjustment.

In practice, α is usually unknown. Recently developed procedures (see, e.g. Johansen, 1988, 1991, 1995a, 1995b; Phillips, 1995) allow for full-system estimation of models involving cointegrated variables, such that, in contrast to the widely used Engle–Granger estimation method (Engle and Granger, 1987), the dynamics of the system are taken into account and the choice of normalisation of the cointegrating vector is allowed to be arbitrary. In these procedures, the estimator of α is super-consistent, converging to α at $O_p(T^{-1})$ rather than the usual $O_p(T^{-1/2})$ without suffering from serious finite sample biases. Therefore, if we replace α_1 and α_2 by their full-system estimates and then estimate the other parameters in the test equations (7) and (8) using OLS, estimates are obtained that are asymptotically the same as they would be if α were known (see, e.g. Lütkepohl, 1991: 359; Boswijk and Urbain, 1997: 33).

Furthermore, we note that the test equations (7) and (8) do not require the choice of instrumental variables as in orthogonality tests of the Hausman type (as in, e.g. von Cramon-Taubadel (1998)). Urbain (1992) shows that the outcomes of these tests are sensitive to the choice of instruments used. The choice of instrumental variable implies imposing the exclusion restriction that this variable is not included in the conditional model. Unfortunately, such an exclusion restriction is often rather arbitrary. However, thanks to our assumption of immediate adjustment in the same time period, an identifying restriction is obtained that allows us to apply orthogonality tests without the need to impose exclusion restrictions.

In the remainder of this section we provide some analytical derivations and simulations to illuminate the relationship between contemporaneous adjustment and data frequency. For this purpose, we return to matrix notation, set $n = 1$ and rewrite (2) in vector error-correction form

$$\Delta x_t = \rho \alpha' x_{t-1} + v_t \tag{9}$$

where $\rho \alpha' = -(\mathbf{I} - \Pi_1)$ and $\rho = (\rho_1, \rho_2)'$ (see (4a) and (4b)). If we lower the frequency to be considered by one period, we obtain

$$\Delta_2 x_t = \rho \alpha'(x_{t-1} + x_{t-2}) + v_t + v_{t-1} \tag{10}$$

where $\Delta_2 x_t = x_t - x_{t-2}$. To substitute $\rho \alpha' x_{t-1}$ out of (10), we pre-multiply (9) by α' and obtain after some rewriting

$$\alpha' x_t = (1 + \alpha' \rho) \alpha' x_{t-1} + \alpha' v_t \tag{11}$$

with stationarity condition $-2 < \alpha' \rho < 0$. If we first multiply (11) by the lag operator, L, and then insert its right-hand side for $\alpha' x_{t-1}$ in (10),

we obtain

$$\Delta_2 x_t = \rho[1 + (1 + \alpha'\rho)]\alpha' x_{t-2} + v_t + (\mathbf{I} + \rho\alpha')v_{t-1}. \qquad (12)$$

The covariance matrix of $v_t + (\mathbf{I} + \rho\alpha')v_{t-1}$ is given by

$$\operatorname{cov}(\Delta_2 x_t | \alpha' x_{t-2}) = \Sigma + (\mathbf{I} + \rho\alpha')\Sigma(\mathbf{I} + \rho\alpha')' \qquad (13)$$

and can be used to derive the parameter of $\Delta_2 z_t$ in the conditional model for $\Delta_2 y_t$ and the parameter of $\Delta_2 y_t$ in the conditional model for $\Delta_2 z_t$.

To generalise, for a $\Delta_f x_t$ model ($f = 2, 3, \ldots$) the residual covariance matrix can be derived by the recursions

$$\operatorname{cov}(\Delta_f x_t | \alpha' x_{t-f}) = \operatorname{cov}(\Delta_{f-1} x_t | \alpha' x_{t-f}) + \left(I + \rho \left[\sum_{j=0}^{f} (1 + \alpha'\rho)^j \right] \alpha' \right)$$

$$\times \Sigma \left(I + \rho \left[\sum_{j=0}^{f} (1 + \alpha'\rho)^j \right] \alpha' \right)'. \qquad (14)$$

Using (14) one may compute the OLS estimates of the coefficient of $\Delta_f z_t$ (or $\Delta_f y_t$) in the conditional model for $\Delta_f y_t$ (or $\Delta_f z_t$) for several chosen values of α, ρ and Σ to check that these estimates converge to the cointegrating parameter in the normalised long-run equilibrium relationship. This result is the underpinning of our conjecture that full adjustment within the same time period is a plausible assumption at lower frequencies. Our identifying restriction will be the assumption that at a certain value of f one of the variables in the cointegrating relationship adjusts immediately to its new equilibrium level.

By way of an example, Table 1 presents the computed parameters of $\Delta_f z_t$ (or $\Delta_f y_t$) in the conditional model for $\Delta_f y_t$ (or $\Delta_f z_t$) when choosing the values[2]

$$\alpha = \begin{pmatrix} 1 \\ -1 \end{pmatrix}, \quad \rho = \begin{pmatrix} -0.15 \\ 0.30 \end{pmatrix} \text{ and } \Sigma = \begin{pmatrix} 1 \\ -1 \end{pmatrix}$$

which have been confirmed as leading to reasonable values for R^2 in the VECM equations in a Monte Carlo simulation with 5000 replications. Table 1 shows that the computed parameter of $\Delta_3 z_t$ in the conditional model for $\Delta_3 y_t$, which is equal to 0.977, is close to the parameter of z_t in the normalised cointegrating relationship, which is equal to unity. Consequently, if we consider $f = 3$ to represent a sufficiently low frequency to justify the assumption that one of the variables in the cointegrating relationship adjusts immediately to its new equilibrium level, then we may apply the test equations (7) and (8), which can be generalised for a bivariate VAR(n) ($n = 1, 2, \ldots$) with respect to $\Delta_f y_t$ and $\Delta_f z_t$

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2   All computations in this paper are performed in the software packages GAUSS and EViews.

Table 1. Computed coefficients and estimated coefficients of $\Delta_f z_t$ (or $\Delta_f y_t$) in the conditional model for $\Delta_f y_t$ (or $\Delta_f z_t$)

f	Computed coefficient of $\Delta_f z_t$ in $\Delta_f y_t$ model[a]	Computed coefficient of $\Delta_f y_t$ in $\Delta_f z_t$ model[a]	Estimated coefficient of $\Delta_f z_t$ in $\Delta_f y_t$ model[a,b]	Estimated coefficient of $\Delta_f y_t$ in $\Delta_f z_t$ model[a,b]
1	0.450	0.150	0.452	0.150
			(0.306, 0.597)	(0.102, 0.198)
2	0.825	0.303	0.826	0.305
			(0.701, 0.950)	(0.258, 0.353)
3	0.977	0.417	0.976	0.419
			(0.873, 1.076)	(0.372, 0.467)
4	1.028	0.503	1.025	0.505
			(0.939, 1.109)	(0.458, 0.552)
5	1.041	0.569	1.039	0.571
			(0.965, 1.110)	(0.526, 0.616)
6	1.042	0.621	1.040	0.623
			(0.975, 1.103)	(0.580, 0.666)
7	1.039	0.662	1.037	0.664
			(0.979, 1.093)	(0.623, 0.705)
8	1.036	0.696	1.034	0.698
			(0.982, 1.085)	(0.659, 0.737)
9	1.032	0.724	1.030	0.726
			(0.982, 1.078)	(0.688, 0.763)
10	1.029	0.747	1.027	0.749
			(0.982, 1.072)	(0.713, 0.784)
15	1.018	0.823	1.017	0.824
			(0.980, 1.054)	(0.792, 0.855)
20	1.013	0.864	1.013	0.865
			(0.979, 1.047)	(0.835, 0.894)

[a] The coefficient in the cointegating relationship is equal to one.
[b] The two-sided lower and upper 95 per cent confidence limits are given in parentheses.

$(f = 1, 2, \ldots)$ to become

$$\Delta_f e_t / \alpha_1 = \varphi_{1f} \Delta_f z_t + \lambda_{1f} e_{t-f} + \sum_{i=0}^{n-1} (\gamma_{11f} \Delta y_{t-f-i} + \gamma_{12f} \Delta z_{t-f-i}) + u_{1ft}$$

(15)

$$\Delta_f e_t / \alpha_2 = \varphi_{2f} \Delta_f y_t + \lambda_{2f} e_{t-f} + \sum_{i=0}^{n-1} (\gamma_{21f} \Delta y_{t-f-i} + \gamma_{22f} \Delta z_{t-f-i}) + u_{2ft}$$

(16)

for $f = 3$ to test $\varphi_{if} = 0$ $(i = 1, 2)$ using, as $f \geq 2$, the Newey–West (1987) heteroscedasticity and autocorrelation consistent standard error for the OLS estimator. By way of approximation, one may also estimate the coefficients and their standard errors in (15) (or (16)) while correcting for the autocorrelation in the residuals, for which an AR(1) process may already be sufficient in many empirical applications.

Clearly, Table 1 shows that for increasing values of f the computed coefficients of $\Delta_f z_t$ in the conditional model for $\Delta_f y_t$ converge much faster to the coefficient of z_t in the cointegrating relationship normalised to y_t than the computed coefficients of $\Delta_f y_t$ in the conditional model for $\Delta_f z_t$ converge to the coefficient of y_t in the cointegrating relationship normalised to z_t. This observation is confirmed by the estimated coefficients (sample means) in Table 1 and their 95 per cent confidence intervals obtained from the Monte Carlo simulation. At $f = 3$ the cointegrating parameter of z_t (y_t) is (is not) in the 95 per cent confidence interval of the coefficient of $\Delta_f z_t$ $(\Delta_f y_t)$ in the conditional model for $\Delta_f y_t$ $(\Delta_f z_t)$. In fact, this test is equivalent to testing $\varphi_{i3} = 0$ $(i = 1, 2)$ in (15) and (16), respectively. Hence, based on the test result and conditional on the assumption that for $f = 3$ one of the variables in the cointegrating relationship fully adjusts immediately, $\Delta_3 z_t$ is concluded to be orthogonal in the conditional model for $\Delta_3 y_t$.

Now that we know on which variable we can condition at $f = 3$, we also know that we can condition on this variable at $f < 3$. This shows that we can use the results of our test for full adjustment for orthogonalisation of the residuals at higher frequencies of interest.

3. Empirical illustration: futures and spot prices

In this section, the spot and futures prices of potatoes in the Netherlands and the spot and futures prices of corn in the USA are investigated for cointegration, causality and exogeneity. In the Netherlands, potato futures are traded on the Amsterdam Exchanges (AEX) and in the USA corn futures are traded on the Chicago Board of Trade (CBOT). An interesting feature of these two empirical cases is that the futures market in the Netherlands is rather thin, whereas the CBOT futures market is deep (i.e. very liquid) (see, e.g. Pennings *et al.*, 1998; Pennings and Leuthold, 2001). One may, therefore, expect that the CBOT futures market contributes significantly more to the discovery of spot market prices than the AEX. Consequently, we expect that unlike potato spot markets in the Netherlands, the spot prices in corn in the USA closely follow the futures price changes according to the long-run equilibrium relationship between the futures price and the spot price.

3.1. The Dutch case: Amsterdam Exchanges

In the Netherlands, there is daily trading on a large number of small, regional potato markets. Once a week, the Rotterdam Potato Cash Market quotes a representative spot price. We use this weekly price quotation in our study, because traders in the Netherlands see it as the reference spot price for

potatoes. The futures price data concern the futures contract for delivery in April (the contract month that accounts for the most volume on the AEX, which is the largest potato futures market in Europe). We consider the closing futures price on the day the Rotterdam market announces the quotation. The data have been obtained from the AEX Clearing Corporation (NLKKAS).

The dataset covers the period October–April for 1989–1990, 1990–1991 and 1991–1992. The October–April period reflects the so-called potato marketing year, as most trade is conducted in this period. Hence, our sample consists of three discontinuous sub-samples, from each of which one loses one observation when taking first differences. Using weekly prices makes our assumption of immediate adjustment plausible, because both markets are very transparent, ensuring quick and efficient information dissemination. Moreover, the futures market is linked to the spot market through the delivery option at maturity of the futures contract. Hence, a long-run equilibrium relationship between spot and futures prices should exist (see, e.g. Brenner *et al.*, 1989; Balabanoff, 1995).

In a strict arbitrage setting, the long-run equilibrium relationship between the spot and futures prices is given by

$$PS_{sw} = (1 + i_s)^{-\tau_{sw}} PF_{sw} \tag{17}$$

where PS_{sw} is the spot price in week w of season s ($s = 1, 2, 3$), i_s is the weekly nominal interest rate in season s and is considered to be constant within season s (all other storage costs are assumed to be negligible), τ_{sw} represents the number of weeks in season s between week w and the week of delivery in April, and PF_{sw} denotes the futures price in week w in season s. Taking natural logarithms, (17) becomes

$$ps_{sw} = -\tau_{sw} \ln(1 + i_s) + pf_{sw} \tag{18}$$

where $ps_{sw} = \ln(PS_{sw})$ and $pf_{sw} = \ln(PF_{sw})$. First, the futures price must be corrected for the basis (defined as spot price minus futures price). To do this, we first regress the differential ($ps_{sw} - pf_{sw}$) on τ_{sw} (without an intercept term) for each s, and then add pf_{sw} to the fitted value of the regression to obtain the futures price corrected for the basis. This corrected futures price will be henceforth denoted as pf_t, with $t = 1, \ldots, T$.

A bivariate VAR of ps_t and pf_t in levels is estimated. Using five as the upper limit for the lag length, the AIC selected a VAR of order one. Next, the prices are tested for their order of integration and cointegration using the Johansen procedure (see, e.g. Johansen and Juselius, 1990; Johansen, 1995b). The results are presented in Table 2. Let p, r and ($p - r - h$) denote the dimension of the VAR, the number of cointegrating vectors and the number of $I(2)$ components, respectively. Here, $p = 2$. Along the lines of Johansen (1992b) we could simplify the deterministic part of the $I(1)$ model to only include a constant. However, conditional on the restriction suggested by the trace and λ-max statistics that $r = 1$, we must reject the restriction that the constant only enters the cointegrating space (p value $= 0.01$). Furthermore, both cointegrating parameters significantly differ from zero. Consequently, ps_t and pf_t

Table 2. Testing for the order of integration and cointegration: the Dutch case[a]

$p - r$	Trace and λ-max statistic for the $I(1)$ model					
	trace	p value	trace 95%	λ-max	p value	λ-max 95%
1	4.19	0.27	8.19	4.19	0.27	8.19
2	48.19	0.00	18.11	44.00	0.00	15.02

$p - r - h$	Trace and λ-max statistics for the $I(2)$ model ($r = 1$ and $\alpha = (1; -1)'$)		
	trace $= \lambda$-max	p value	95% critical value
1	42.82	0.00	8.19

[a] p values and critical values are obtained from MacKinnon *et al.* (1999), case III with $k = 0$.

are cointegrated. Given $r = 1$ and an unrestricted constant, we cannot reject the restriction that the cointegrating parameters of the prices are equal but have opposite signs (p value $= 0.48$). This result supports our assumption that we may ignore the storage costs that would be modelled by including an additive term in (17). Conditional on this cointegrating vector we test for $I(2)$ components in the data using a VAR(2). The statistic in Table 2 strongly rejects the restriction $h = 0$. Hence, there are no $I(2)$ components.

Because $\alpha = (1; -1)'$ is the estimate of the cointegrating vector, the error-correction term, denoted by e_t, is given by

$$e_t = ps_t - pf_t. \tag{19}$$

On the basis of this result, we obtain the following parameter estimates in the marginal error-correction models (t values in parentheses, * indicates significance at the 5 per cent level):

$$\Delta ps_t = 0.023 - 0.529 e_{t-1} + \hat{v}_t^{ps} \tag{20a}$$
$$ (4.50^*) \quad (-4.17^*)$$

$T = 73$; $R^2 = 0.20$; $\hat{\sigma} = 0.04$; $DW = 1.39$; $ARCH$1-1 $F(1, 70) = 0.63$

$$\Delta pf_t = 0.017 + 0.092 e_{t-1} + \hat{v}_t^{ps} \tag{20b}$$
$$ (2.80^*) \quad (0.15)$$

$T = 73$; $R^2 = 0.00$; $\hat{\sigma} = 0.05$; $DW = 1.76$; $ARCH$1-1 $F(1, 70) = 1.58$.

In (20a) and (20b) we test for the absence of a seasonally varying intercept term. The null hypothesis is not rejected. Furthermore, as already indicated by the DW statistic, there is some autocorrelation left in the residuals of (20a). Nevertheless, after omitting the last five observations of the 1989–1990 season where the residuals are relatively large and volatile, no auto-correlation remains whereas the other regression results are not significantly different from those presented in (20a).

The estimates show that e_{t-1} significantly enters (20a), but can be deleted from (20b). Consequently, we conclude that the futures price is weakly exo-genous for the long-run parameters. Because we have a bivariate VAR(1),

weak exogeneity coincides with Granger non-causality and hence, long-run non-causality (see Bruneau and Jondeau, 1999: 548). This result implies that the futures price drives the spot price in the long run. Such an outcome supports the view that the futures price is an important price indicator in the spot market. However, does this also imply that the futures price is ortho-gonal in the conditional model for the spot price and hence, weakly exogenous with respect to the short-run parameters? To answer this question, we con-sider the results of our test procedure:

$$\Delta e_t = 0.006 + 0.029\Delta ps_t - 0.536e_{t-1} + \hat{u}_t^{ps} \qquad (21)$$
$$\quad\quad (1.45) \qquad (0.38) \qquad (-5.72^*)$$

$T = 73; \quad R^2 = 0.38; \quad F(2, 70) = 21.61^*; \quad \hat{\sigma} = 0.03; \quad DW = 1.94; \quad ARCH1\text{-}1$
$F(1, 70) = 0.57$

$$\Delta e_t = 0.011 - 0.293\Delta pf_t - 0.545e_{t-1} + \hat{u}_t^{ps} \qquad (22)$$
$$\quad\quad (3.67^*) \quad (-5.13^*) \qquad (-7.60^*)$$

$T = 73; \quad R^2 = 0.55; \quad F(2, 70) = 42.71^*; \quad \hat{\sigma} = 0.02; \quad DW = 1.57; \quad ARCH1\text{-}1$
$F(1, 70) = 0.17.$

In (21) we can delete Δps_t at the 5 per cent significance level, but in (22) we cannot omit Δpf_t (p value $\ll 0.001$). Hence, the test for weak exogeneity with respect to the short-run parameters conditional on the full adjustment hypothesis shows what we are unable to observe testing for the absence of long-run causality. In other words, in contrast to the spot price, the futures price is correlated with the innovations in e_t, which is in line with a thin futures market. Such a market is characterised by large price changes as a result, in part, of lack of liquidity. These futures price changes, therefore, are not followed fully by spot traders in so far as they consider these price changes to be unrelated to market fundamentals. We know this because the estimated coefficient of Δpf_t in (22) is -0.293 (i.e. smaller than zero), and $\Delta e_t = \Delta ps_t - \Delta pf_t$ so adding Δpf_t to both sides of (22) leads to $\Delta ps_t = 0.707\Delta pf_t + \dots$. This indicates partial adjustment on the part of the spot traders and, in our model, is a source of normalisation bias. In contrast, according to (21) the futures price fully adjusts to changes in the spot price, which complies with the transparency of both markets when followed on a weekly basis.

3.2. The US case: Chicago Board of Trade

We subsequently analyse the price data for corn on the CBOT. The futures contract relates to delivery in March 1998, which was the largest contract among all months of delivery in 1998. The dataset contains 390 daily obser-vations on spot and futures closing prices and covers the period 23 September 1996 to 20 March 1998. The behaviour of the daily data confirms the proof by Samuelson (1965) that today's closing price is the best predictor of tomor-row's closing price. This result does not leave any room for statistical evidence on cointegration, as cointegration implies error correction and, in turn, error

Table 3. Testing for the order of integration and cointegration: the CBOT case[a]

Monday ($i = 1$)	Trace and λ-max statistic for the $I(1)$ model					
$p - r$	trace	p value	trace 95%	λ-max	p value	λ-max 95%
1	0.23	0.69	4.13	0.23	0.69	4.13
2	19.25	0.00	12.32	19.02	0.00	11.23

	Trace and λ-max statistics for the $I(2)$ model ($r = 1$ and $\alpha = (1; -1)'$)		
$p - r - h$	trace $= \lambda$-max	p value	95% critical value
1	39.98	0.00	4.13

Tuesday ($i = 2$)	Trace and λ-max statistic for the $I(1)$ model					
$p - r$	trace	p value	trace 95%	λ-max	p value	λ-max 95%
1	0.03	0.89	4.13	0.03	0.89	4.13
2	11.59	0.07	12.32	11.56	0.04	11.23

	Trace and λ-max statistics for the $I(2)$ model ($r = 1$ and $\alpha = (1; -1)'$)		
$p - r - h$	trace $= \lambda$-max	p value	95% critical value
1	50.08	0.00	4.13

Wednesday ($i = 3$)	Trace and λ-max statistic for the $I(1)$ model					
$p - r$	trace	p value	trace 95%	λ-max	p value	λ-max 95%
1	0.01	0.93	4.13	0.01	0.93	4.13
2	14.26	0.02	12.32	14.26	0.01	11.23

	Trace and λ-max statistics for the $I(2)$ model ($r = 1$ and $\alpha = (1; -1)'$)		
$p - r - h$	trace $= \lambda$-max	p value	95% critical value
1	38.40	0.00	4.13

Thursday ($i = 4$)	Trace and λ-max statistic for the $I(1)$ model					
$p - r$	trace	p value	trace 95%	λ-max	p value	λ-max 95%
1	0.06	0.84	4.13	0.06	0.84	4.13
2	13.79	0.03	12.32	13.73	0.02	11.23

	Trace and λ-max statistics for the $I(2)$ model ($r = 1$ and $\alpha = (1; -1)'$)		
$p - r - h$	trace $= \lambda$-max	p value	95% critical value
1	46.49	0.00	4.13

Table 3. (Continued)

Friday ($i = 5$) $p - r$	Trace and λ-max statistic for the $I(1)$ model					
	trace	p value	trace 95%	λ-max	p value	λ-max 95%
1	0.00	0.98	4.13	0.00	0.98	4.13
2	10.65	0.09	12.32	10.65	0.06	11.23

$p - r - h$	Trace and λ-max statistics for the $I(2)$ model ($r = 1$ and $\alpha = (1; -1)'$)		
	trace $= \lambda$-max	p value	95% critical value
1	50.46	0.00	4.13

[a] p values and critical values are obtained from MacKinnon *et al.* (1999), case I with $k = 0$.

correction implies predictability of tomorrow's closing price more than know-ing today's closing price (see, e.g. Copeland, 1991; Yang and Leatham, 1998). Nonetheless, if we lower the frequency of our data, for example, by consider-ing the weekly quotations of a specific day in the week, then the predictability implied by cointegration concerns predictions with a lower accuracy than the predictions offered by today's closing price. Using weekly data of a specific day in the week reduces the information set of the rational agents and there-fore, predictability as a consequence of cointegration in these weekly data is not in contradiction with market efficiency.

To apply our orthogonality test, we need cointegration and hence we con-sider weekly data from each day in the week. This leads to weekly price data with respect to Monday, weekly price data with respect to Tuesday, up to and including weekly price data with respect to Friday. For each i ($i = 1, 2, 3, 4$ or 5, representing Monday, Tuesday, Wednesday, Thursday and Friday, respec-tively) a VAR of ps_{it} and pf_{it} in levels is considered, where pf_{it} is corrected for the basis in the same way as we corrected the futures price for the basis in the Dutch case. Using five as the upper limit for the lag length, the AIC selected a VAR of order one for all i. Next, the VARs are rewritten as reduced-form error-correction models and the prices are tested for their order of integration and cointegration using the Johansen procedure. The results are presented in Table 3. Along the lines of Johansen (1992b) and using the 10 per cent significance level for $i = 2$ and $i = 5$ instead of the 5 per cent significance level, we select the VAR without deterministic terms and according to the trace and λ-max statistics for the $I(1)$ model, we can impose the restriction $r = 1$ for each i, whereas the restriction $r = 0$ is rejected. Furthermore, for each i both cointegrating parameters differ significantly from zero. Conse-quently, ps_{it} and pf_{it} are cointegrated. Given $r = 1$, we test the restriction that the cointegrating parameters of the prices are equal but have opposite signs. The results of the test statistic are in favour of the restriction (p values are 0.61, 0.84, 0.56, 0.85 and 0.95 for $i = 1, 2, 3, 4$ and 5, respectively). Conditional on such a cointegrating vector we test for $I(2)$ components in the

Table 4. Coefficient estimates in the marginal error-correction models: the CBOT case[a]

Dependent variable	i	Coefficients independent variables		T	R^2	$\hat{\sigma}$	DW	p value ARCH1-1
		intercept	$e_{i,t-1}$					
Δps_{1t}	1	−0.009	−0.481	67	0.22	0.03	2.24	0.05
		(−2.06*)	(−4.28*)					
Δpf_{1t}	1	−0.004	−0.103	67	0.02	0.02	2.15	0.20
		(−1.15)	(−1.28)					
Δps_{2t}	2	−0.006	−0.301	72	0.11	0.03	2.19	0.92
		(−1.52)	(−2.88*)					
Δpf_{2t}	2	−0.003	−0.125	72	0.03	0.03	2.23	0.92
		(−0.86)	(−1.49)					
Δps_{3t}	3	−0.006	−0.309	72	0.14	0.03	1.80	0.58
		(−1.81)	(−3.32*)					
Δpf_{3t}	3	−0.002	−0.122	72	0.03	0.02	1.92	0.50
		(−0.80)	(−1.59)					
Δps_{4t}	4	−0.006	−0.342	66	0.13	0.03	2.32	0.50
		(−1.55)	(−3.05*)					
Δpf_{4t}	4	−0.003	−0.182	66	0.05	0.02	2.47	0.44
		(−0.75)	(−1.74)					
Δps_{5t}	5	−0.006	−0.282	69	0.09	0.03	2.22	0.63
		(−1.54)	(−2.65*)					
Δpf_{5t}	5	−0.003	−0.125	69	0.02	0.03	2.42	0.42
		(−0.72)	(−1.30)					

[a] t values in parentheses.
*Significant at the 5 per cent level.

data using a VAR(2). The statistics in Table 3 strongly reject the restriction $h = 0$. Hence, there are no $I(2)$ components.

The coefficient estimates in the marginal error-correction models are presented in Table 4, where e_{it} $(= ps_{it} - pf_{it})$ are the deviations from the equilibrium as given by the cointegrating relationships. For all i it appears that the spot price is driven by the futures price, as $e_{i,t-1}$ significantly enters the equation of Δps_{it}, but not of Δpf_{it}. Consequently, we conclude that the futures price is weakly exogenous for the long-run parameters.

Next, conditional on the immediate adjustment assumption, we test for orthogonality by applying the test regressions presented in Table 5. For all i, Δps_{it} cannot be deleted from the equation at the 5 per cent significance level, whereas the coefficient of Δpf_{it} is never significant. Consequently, conditional on the immediate adjustment hypothesis we conclude that the futures price is orthogonal in the conditional model for the spot price, and, because

Table 5. Coefficient estimates in the test regressions: the CBOT case[a]

Dependent variable	i	Coefficients independent variables				T	R^2	$\hat{\sigma}$	DW	p value ARCH1-1
		intercept	Δps_{it}	Δpf_{it}	$e_{i,t-1}$					
Δe_{1t}	1	−0.001	0.579		−0.099	67	0.60	0.02	2.40	0.00[b]
		(−0.24)	(7.91*)		(−1.32)					
Δe_{1t}	1	−0.007		−0.190	−0.397	67	0.23	0.03	2.49	0.03[b]
		(−1.81)		(−1.35)	(−4.27*)					
Δe_{2t}	2	−0.001	0.317		−0.081	72	0.44	0.01	2.25	0.68
		(−0.14)	(6.25*)		(−1.72)					
Δe_{2t}	2	−0.004		0.061	−0.168	72	0.13	0.02	2.21	0.12
		(−1.77)		(0.78)	(−3.00*)					
Δe_{3t}	3	−0.002	0.333		−0.084	72	0.42	0.01	1.83	0.87
		(−1.09)	(5.73*)		(−1.73)					
Δe_{3t}	3	−0.004		−0.016	−0.189	72	0.14	0.02	1.70	0.81
		(−1.94)		(−0.18)	(−3.38*)					
Δe_{4t}	4	−0.002	0.171		−0.101	66	0.25	0.01	1.98	0.96
		(−1.37)	(3.21*)		(−1.99)					
Δe_{4t}	4	−0.004		−0.043	−0.168	66	0.14	0.01	1.83	0.90
		(−1.93)		(−0.70)	(−3.20*)					
Δe_{5t}	5	−0.002	0.231		−0.092	69	0.27	0.02	2.00	0.17
		(−1.09)	(3.93*)		(−1.71)					
Δe_{5t}	5	−0.004		−0.062	−0.165	69	0.11	0.02	1.79	0.05
		(−1.74)		(−0.89)	(−2.88*)					

[a] t values in parentheses.
* Significant at the 5 per cent level.
[b] No longer significant after omitting some outliers that do not have an impact on the other regression results.

the futures price has already been confirmed as weakly exogenous for the long-run parameters, it is now also weakly exogenous with respect to the short-run parameters. A joint test of weak exogeneity of the futures price for both the long- and short-run parameters (see Boswijk and Urbain, 1997) leads to non-rejection for each i as well, confirming the results of the two separate tests for each i. So, in contrast to the AEX where spot traders acknowledge the long-run price discovery function of futures prices without fully incorporating the substantial short-run futures price changes into spot prices, we find that the CBOT performs its price discovery function in such a way that spot traders immediately transmit the futures price changes along the lines of the long-run equilibrium relationship between the futures and the spot prices. This result complies with the high liquidity of the CBOT corn futures contract, delivery March, as opposed to the potato futures contract traded in Amsterdam (see, e.g. Pennings *et al.*, 1998). This

result is confirmed by the much higher interest rates (i_s) that we estimate in the Dutch case (>50 per cent) when correcting the futures price for the basis as compared with the interest rate estimates of about 2.7 per cent per year that we find for the CBOT.

4. Conclusions and suggestions for further research

In this paper, we have introduced a procedure to test for orthogonality of the errors in VECMs. The test procedure is based upon the fact that, when the frequency of the data is chosen to be sufficiently low, one of the variables in the long-run equilibrium relationship adjusts fully within the same time period to its new equilibrium level. Applying our test to the futures and spot potato price relationship in the Netherlands shows that the spot price is orthogonal in the conditional model for the futures price, notwithstanding the result that the test for weak exogeneity for the long-run parameters shows that the spot price is driven by the futures price in the long run. This result of the orthogonality test might be explained by the lack of liquidity on the AEX, leading to futures price changes that are so extreme that spot traders refuse to transmit them fully into the spot price.

In contrast to the results for the Netherlands, we find that the price of the CBOT futures contract for corn, delivery March 1998, is weakly exogenous for both long- and short-run parameters. Consequently, the spot price is not just driven by the futures price, but rather fully adjusts to the changes in the futures price. Hence, spot traders are able to fully transmit the futures price changes into the spot price. Therefore, the CBOT futures price is not only the reference price in the long run, but it is also the representative price for changes over a time interval that is large enough (we show this for a weekly interval) to capture the delays caused by the adjustment costs faced by spot traders when following futures price changes.

As we outline our test procedure in a bivariate setting, several extensions are of interest. Applying our method to a higher-dimensional system with one cointegrating relationship is straightforward, but more than one cointegrating vector requires the identification of the long-run structure along the lines of Johansen and Juselius (1994) (see also Johansen, 1995a, 1995b; Boswijk, 1995). Nevertheless, our test procedure can still be used to determine which variables adjust immediately to their new equilibrium levels. From the vector error-correction model we can derive test equations as in (15) and (16) for each cointegrating relationship. We consider as many test equations as there are variables in the cointegrating relationship with respect to each cointegrating relationship. All test equations are similar, except that the right-hand side of each test equation includes the unlagged first difference of another variable in the cointegrating relation. In each test equation, the unlagged difference term is tested for its absence. If more than one of the unlagged difference terms do not significantly enter their respective test equations, then their joint absence can simply be tested by including all of them in one test equation.

Orthogonality tests for testing recursive models of the errors in VECMs require identifying restrictions. Our test procedure provides a way to search for these without the use of instrumental variables.

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