

# Examining long-run relationships between Australian beef prices<sup>†</sup>

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Cointegration and impulse response analyses are used to investigate the short-run and long-run dynamics of the Australian beef market. The aim of this study is to determine whether long-run relationships existed between Australian beef prices at the farm, wholesale and retail levels. Based on monthly data from 1971 to 1994, the results show that all three prices considered are cointegrated. Furthermore, the wholesale price is found to be weakly exogenous. The latter result might be an indication of market inefficiency due in part to price levelling often practised in the beef marketing system.

## 1. Introduction

The price transmission process in the food marketing system has been of considerable interest to economists because of the implications for market efficiency (Williams and Bewley 1993). Market efficiency implies that in a competitive market with perfect information, arbitrage will ensure that price differentials in related spatial, temporal or product transformation markets reflect the costs of providing marketing services. For example, price differentials in spatially related markets should reflect the cost of transportation while in temporally related markets, the cost of storage (Faminow and Benson 1990; Goodwin and Schroeder 1991). Furthermore, a price change in one market will be followed by similar price changes in related markets. Therefore, this comovement of related price series implied by market efficiency suggests the existence of long-run relationships between them. This means, in econometric terminology, that these price series should be cointegrated. In cases where a market is not efficient as a result of market

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imperfection, the price links are weak and the related price series may be found not to be cointegrated.

Conventionally, efficiency in price transmission processes is examined using standard econometric methods. However, evolving developments in time series econometrics cast doubt on the use of standard econometric methods for estimating commodity models because of the nonstationarity often found in time series data (Myers 1994). Moreover, the standard econometric studies have tended to focus on the short-run dynamics of price relationships with little reference to long-run equilibrium conditions. Possible inconsistencies between short-run and long-run behaviour of prices is of interest because of their implications for the distribution of the benefits from R&D or deregulation (Wohlgenant 1989).

In a recent commentary, it is argued that 'while live cattle prices have slumped by as much as 50 per cent over the year, wholesale prices have only dropped 7 per cent, and worst of all retail prices have remained unchanged' (Williamson 1997). The author went on to imply that the beef marketing system is not efficient. Although the conclusion is premature given the simplicity of the analysis, the question deserves further examination.

The overall purpose of this study therefore is to re-examine market efficiency issues with an improved methodology, as suggested by Myers (1994). The specific objective is to determine whether long-run relationships exist between Australian beef prices at the farm, wholesale and retail levels using cointegration techniques. Monthly data from 1971 to 1994 are used for the analysis. Results show that all three prices considered are integrated of order one, and that they are cointegrated. Furthermore, the wholesale price is weakly exogenous. The latter result is an indication of market inefficiency due in part to price levelling often practised in the beef marketing system.

The article begins with a brief introduction of the main concepts involved in the econometric time series literature, particularly the relationship between vector autoregression, error correction models and cointegration. In the next section, an error correction model is used to test for cointegration. The results based on the Johansen procedure and impulse response analysis are then reported. Discussion of the results is then provided, followed by areas for further research and concluding remarks.

## 2. Cointegration

Vector autoregressive (VAR) models, error correction models (ECMs) and cointegration are related concepts in time series analysis used to characterise the relationships between the series being studied. In essence, it can be shown that, with re-parameterisation, the ECM is a standard VAR in first differences augmented by error correction terms. Moreover, according to the

Granger Representation theorem, an ECM representation for a set of variables which are integrated of order one implies cointegration among the variables and vice versa (Engle and Granger 1987). These relationships are described below.

A standard VAR with lag length  $p$ , VAR( $p$ ), can be written as:

$$x_t = A_0 + A_1x_{t-1} + \dots + A_px_{t-p} + BD_t + CS_t + v_t, \quad t = 1, \dots, T, \quad (1)$$

where:

- $p$  = lag length;
- $x_t$  = an  $(n \times 1)$  vector of endogenous variables;
- $A$ 's =  $(n \times n)$  matrices of unknown parameters;
- $x_{t-j}$  = an  $(n \times 1)$  vector of the  $j$ th lagged value of  $x_t$ ;
- $D_t$  = a set of centred seasonal dummies;
- $S_t$  = a set of dummy variables representing structural changes; and
- $v_t$  = white-noise disturbance terms which may be contemporaneously correlated.

An ECM with lag length  $p$ , ECM( $p$ ), can be derived from the above VAR( $p$ ) by re-parameterisation. That is, simply by term manipulation, an ECM of the following form can be obtained:

$$\begin{aligned} \Delta x_t &= \Pi_0 + \Pi_1\Delta x_{t-1} + \dots + \Pi_{p-1}\Delta x_{t-(p-1)} + \Pi x_{t-p} + BD_t + CS_t + v_t \\ &= \Pi_0 + \sum_{i=1}^{p-1} \Pi_i \Delta x_{t-i} + \Pi x_{t-p} + BD_t + CS_t + v_t, \end{aligned} \quad (2)$$

where:

$$\begin{aligned} \Pi_0 &= A_0, \\ \Pi_j &= -\left(I - \sum_{i=1}^j A_i\right), \quad j = 1, 2, \dots, p-1; \\ \Pi &= -\left(I - \sum_{i=1}^p A_i\right); \quad \text{and} \end{aligned}$$

$$\Delta x_{t-j} = \text{an } (n \times 1) \text{ vector of } x_{t-j} \text{ in first differences, } j = 1, 2, \dots, p-1.$$

Other variables are as previously defined. Therefore, without any loss of information, the ECM( $p$ ) in equation 2 is a transformation of the VAR( $p$ ) in equation 1 expressed in first differences augmented by the error correction term,  $\Pi x_{t-p}$ . For detailed derivations, see Enders (1995, pp. 389–90).

The  $\Pi$  matrix in equation 2, which is termed the long-run impact matrix of the ECM, is of primary importance. First, the rank of  $\Pi$  provides the basis for determining the existence of cointegration or the long-run relationship between the variables. According to Johansen (1988), there are three possibilities with regard to the rank of  $\Pi$ :

- Case 1 If  $\text{rank}(\Pi)$  is zero, then the variables are not cointegrated and the model is equivalent to a VAR in first differences;
- Case 2 If  $0 < \text{rank}(\Pi) < n$ , then the variables are cointegrated; and
- Case 3 If  $\text{rank}(\Pi) = n$ , then the variables are stationary and the model is equivalent to a VAR in levels.

Second, the  $\Pi$  matrix can be decomposed into the product of matrices  $\alpha$  and  $\beta$ , i.e.  $\Pi = \alpha\beta'$ .  $\alpha$  is the matrix of speed of adjustment coefficients, which characterises the long-run dynamics of the system, while  $\beta$  is the matrix representing the cointegrating relations in which  $\beta'x_t$  (the disequilibrium error) is stationary (Johansen and Juselius 1990). A large (small) value of  $\alpha$  means that the system will respond to a deviation from the long-run equilibrium with a rapid (slow) adjustment. On the other hand, if the  $\alpha$ s are zero for some equations, it implies that the corresponding variables do not respond to the disequilibrium error and, hence, may be weakly exogenous.

In summary, cointegration of a set of time series implies that long-run stationary relationships exist among the component non-stationary series. Because these series are linked by common stochastic trends, they do not move independently of each other and there are systematic co-movements among the series. Furthermore, the time paths of cointegrated variables are influenced by deviations from the long-run equilibrium. In addition to the ECM, the dynamics of the cointegrated system can be investigated via the impulse response or dynamic multiplier analysis. The impulse response functions trace out the responses of endogenous variables in the system to shocks in the error terms. Both the ECM and the impulse response analysis are used to analyse Australian beef prices in the following sections.

### 3. Data

Monthly beef price data in New South Wales (NSW) for the period January 1971 to September 1994, a total of 285 observations, are used for the current analysis. The data set includes prices for beef, suitable for the domestic market, at three market levels: farm, wholesale and retail. The price variables are defined as follows:

**Farm price.** Monthly auction price (cents/kg of estimated dressed carcass weight) of a composite beef carcass, adjusted for by-products and shrinkage to achieve retail equivalent weight. The composite carcass is a weighted average of the various cattle types suitable for the domestic market. By-product values include hide, meat meal, tallow and edible offal.

**Wholesale price.** Monthly wholesale price (cents/kg) of a composite beef carcass, delivered to butchers and supermarkets, adjusted for shrinkage to achieve retail equivalent weight.

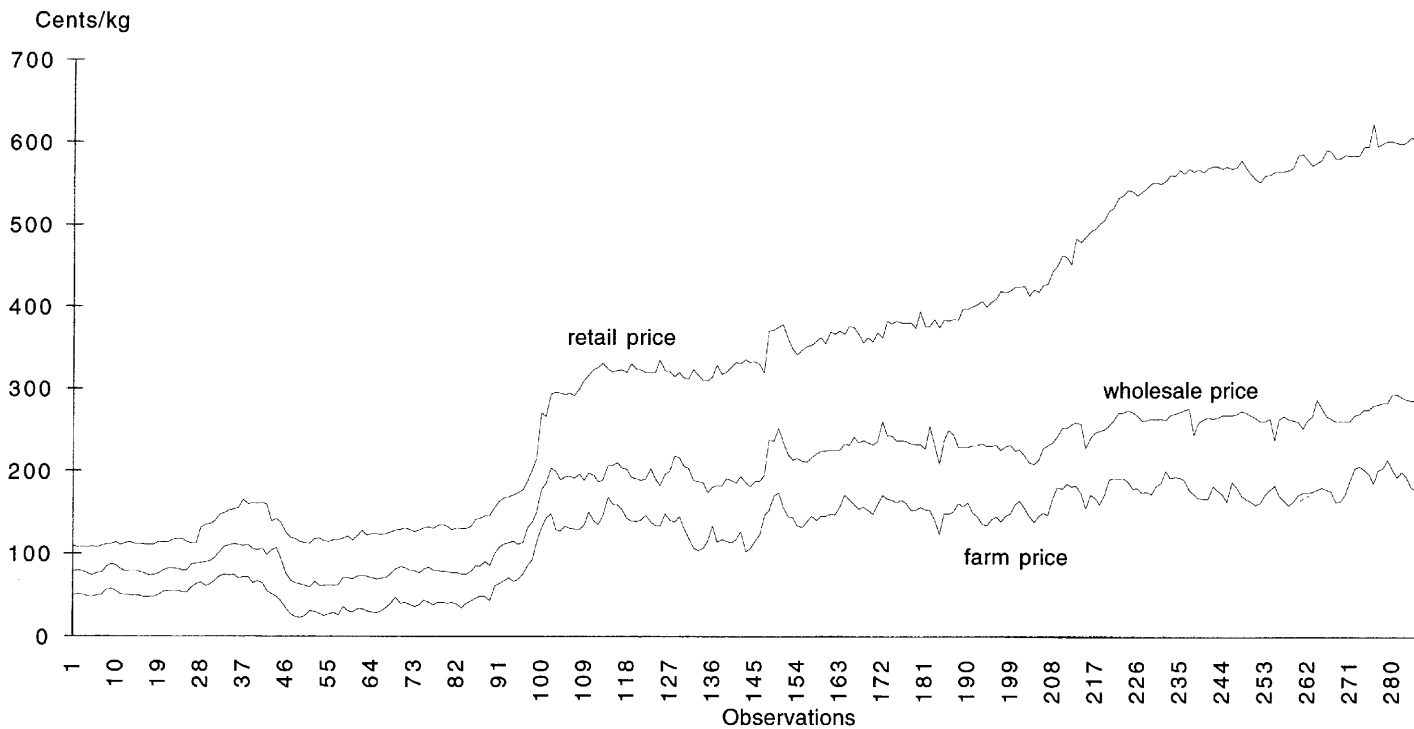
**Retail price.** Monthly retail price (cents/kg) of a composite beef carcass. The prices of various cuts are weighted by carcass proportions determined by cutout tests.

These price series are obtained from the records of the Economic Services Unit of NSW Agriculture. They have been used in various studies of meat price relationships (for example, Griffith, Green and Duff 1991).

These three price series are presented in figure 1. The co-movements in these price series suggest that there may be potential cointegration among the variables. Several other observations can also be made. First, there appears to be a structural break in the 1970s. Second, it appears that prices are more volatile in the 1970s than the later period. Third, there are differences in price variability among prices at different market levels.

In terms of the structural break, it can be seen that prices are relatively low during the period from 1974 to 1979. The price slump is a result of changes that occurred in the marketing environment for Australian beef. Of particular relevance are the formation of the European Community in 1972, the closure of the North Asian markets in late 1974 and the two oil price shocks in the 1970s.

In terms of price variability over time and across markets, it can be seen from table 1 that during the period 1971–79, the farm price is more volatile than the wholesale price, which is, in turn, more volatile than the retail price. However, during the second period the retail price is most volatile, followed by the farm price and the wholesale price. Another observation is that, the farm, wholesale and retail prices are more volatile during the period 1971–79 than the period 1980–94. The contributing factors to these changes in price variability include: the emphasis on diversification of export markets, changes in the supply side of the industry in response to the price slump of the mid-1970s and the increasing importance of feedlots and direct marketing between producers and large retailers. The following analysis based on the ECM will examine these structural issues.



**Figure 1** Monthly beef prices in the Australian domestic market, 1971.1–1994.9

**Table 1** Price variability of original beef price series (cents/kg), 1971–94

	Mean	Standard Deviation	COV <sup>a</sup> (%)
<i>Farm price</i>			
1971–1994	120.94	56.45	46.68
1971–1979	56.50	28.34	50.15
1980–1994	160.26	23.50	14.66
<i>Wholesale price</i>			
1971–1994	183.81	76.95	41.86
1971–1979	94.38	34.15	36.18
1980–1994	238.38	30.68	12.87
<i>Retail price</i>			
1971–1994	332.63	171.14	51.45
1971–1979	145.00	48.54	33.48
1980–1994	447.11	105.10	23.50

Note: <sup>a</sup> COV = (standard deviation/mean) \* 100.

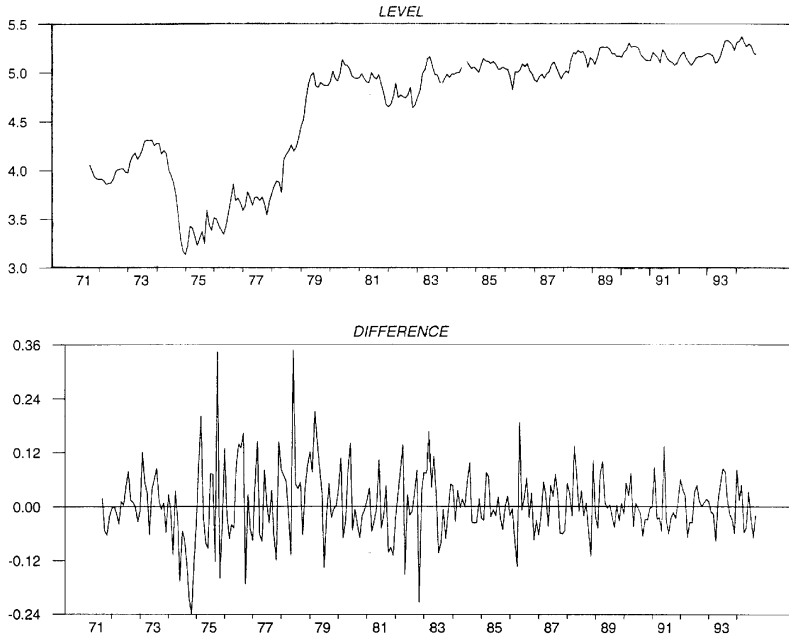
#### 4. The error correction model and estimated results

In this section, an ECM is developed to test the hypothesis that long-run relationships exist between Australian beef prices based on the Johansen procedure. The Johansen procedure, as suggested in Enders (1995, pp. 396–400), includes the following four steps:

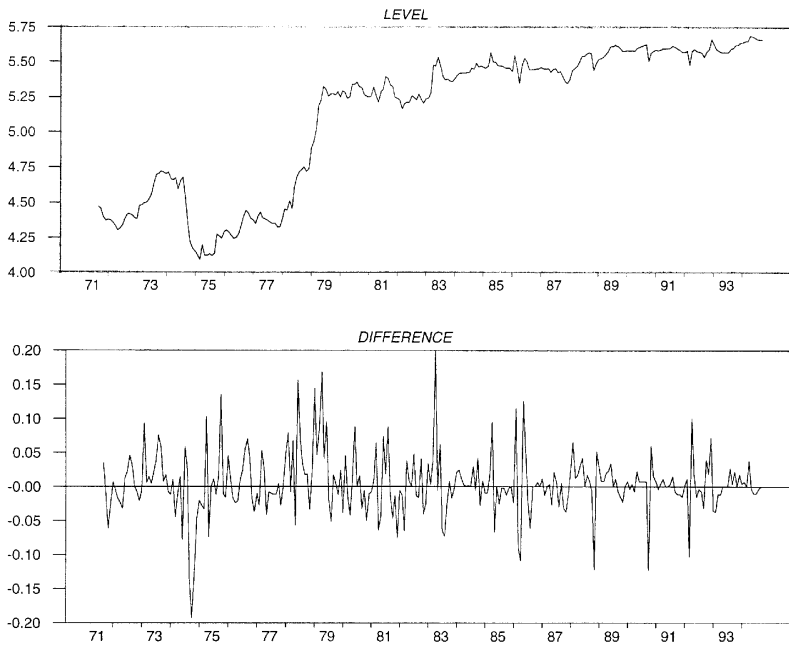
- Step 1 Pre-test the order of integration and determine the lag length for the ECM based on a standard VAR.
- Step 2 Estimate the ECM and determine the rank of  $\Pi$ .
- Step 3 Analyse the cointegrating vector(s) and the speed of adjustment coefficients.
- Step 4 Perform innovation accounting and causality tests on the ECM.

In Step 1, the order of integration of beef prices (hereafter expressed in logarithms) at each of the farm, wholesale and retail levels is tested based on the augmented Dickey-Fuller (DF) tests. Although the Johansen test can detect differing orders of integration, it is important not to mix variables with different orders of integration. As such, the pre-test employed here is to ensure that all variables are of the same order of integration.

The unit root tests are performed using SHAZAM (Version 7, 1993). The results indicated that all three log beef prices in levels have unit roots. Further unit root tests on the series in first differences indicate absence of unit roots. These results confirm that the three log price series are indeed integrated of order one (I(1)). The price series, both in first differences and in levels, are shown in figures 2–4.

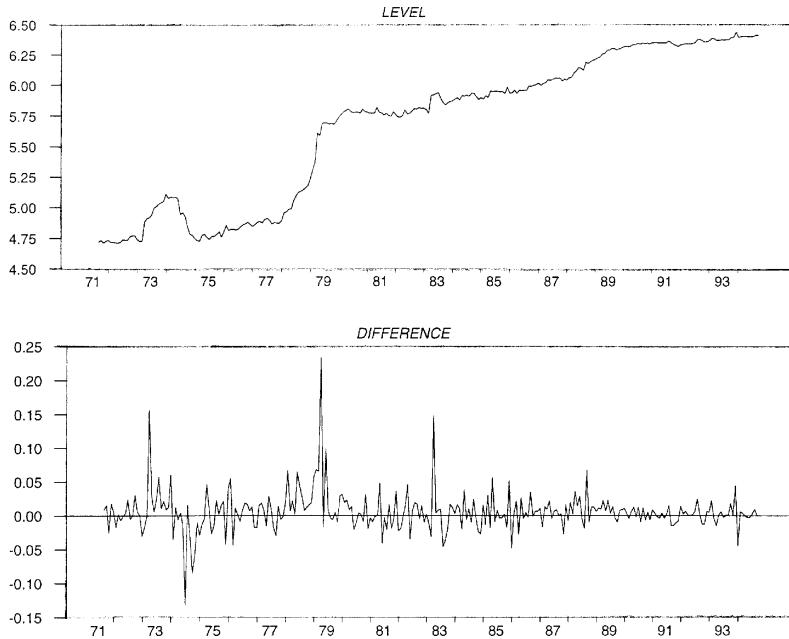


**Figure 2** Monthly log farm prices in levels and first differences, 1971.1–1994.9



**Figure 3** Monthly log wholesale prices in levels and first differences, 1971.1–1994.9





**Figure 4** Monthly log retail prices in levels and first differences, 1971.1–1994.9

After confirming that the price series under consideration are  $I(1)$ , the next task is to determine the proper lag length for the ECM. The testing procedures involve making pair-wise comparisons between two standard VARs, each having a different lag length. The tests can be done based on the likelihood ratio (LR) tests or the Akaike information criterion (AIC) (Enders 1995, pp. 312–15). Applying the beef data to the standard VAR( $p$ ) specified in equation 1, the variables are defined as follows:

$x_t = [FP_t, WP_t, RP_t]'$  = a vector of farm, wholesale and retail prices, expressed in logarithms;

$x_{t-i} = [FP_{t-i}, WP_{t-i}, RP_{t-i}]'$  for  $i = 1, 2, 3, \dots, p$ ;

$S_t = 1$  for the period between 1974.1 to 1979.12; and  $S_t = 0$ , otherwise;

$D_t$  = monthly seasonal dummies using December as the base period; and

$A$ 's,  $B$  and  $C$  = unknown parameters to be estimated.

Other variables are as previously defined.  $S_t$  is included in the VAR model to reflect a range of changes in the Australian beef market which occurred during the 1970s, as discussed above.

Using RATS (Version 4, 1995), paired comparisons, each with different lag lengths, are made to determine the appropriate lag length for the ECM. The LR test results, which are presented in table 2, suggest that  $p = 8$ .

**Table 2** Summary of test statistics for determining the lag length in standard VARs

Lag length			
Unrestricted model	Restricted model	Calculated $\chi^2(df)^a$	Test results
12	8	29.73(36)	fail to reject $H_0$
8	7	16.74(9)	reject $H_0$

Note: <sup>a</sup> Critical  $\chi^2$  values at the 10 per cent significance level are 14.68 and 40.25 for 9 and 30 degrees of freedom, respectively.

With the lag length of the ECM now determined, the next task is to set up an ECM(8) for the beef data. Based on the results obtained from Step 1, the ECM for the log beef prices is specified as:

$$\Delta x_t = \Pi_0 + \sum_{i=1}^7 \Pi_i \Delta x_{t-i} + \Pi x_{t-8} + BD_t + CS_t + v_t, \quad (3)$$

where:

$\Delta x_t = [\Delta FP_t, \Delta WP_t, \Delta RP_t]'$  = log farm, wholesale and retail prices in first differences;

$\Delta x_{t-i} = [\Delta FP_{t-i}, \Delta WP_{t-i}, \Delta RP_{t-i}]'$  for  $i = 1, 2, \dots$ , and 7;

$x_{t-8} = [FP_{t-8}, WP_{t-8}, RP_{t-8}]'$ ;

$\Pi_i s = (3 \times 3)$  matrices of unknown parameters representing the short-run dynamics;

$\Pi = \alpha\beta'$  = the matrix of unknown parameters representing the long-run dynamics; and

$v_t$  = white noise disturbance terms which may be contemporaneously correlated.

Other variables are as previously defined. Two versions of the ECM(8), Models A and B, are estimated based on the Johansen procedure using CATS in RATS (Doan 1996). Model A is the unrestricted model where the constant term is incorporated in the equation. Model B is the restricted model where the constant term is incorporated in the cointegrating vectors. The estimated results regarding the rank of  $\Pi$ ,  $r$ , for both versions are presented in table 3.

As can be seen in table 3, the maximal eigenvalue statistics ( $\lambda_{\max}$ ) and the trace statistics ( $\lambda_{\text{trace}}$ ) indicated that the rank of  $\Pi$  is one for both Model A and Model B. This meant that the three beef prices are cointegrated with one cointegrating relation. To discriminate between Models A (with a trend drift) and B (with a constant term in the cointegrating vector), the test statistic  $LR_\lambda$  which is suggested in Enders (1995, pp. 393), is used. The  $LR_\lambda$  is defined as:

**Table 3** Summary of rank tests on matrix  $\Pi$  of the error correction model

	Estimated $\lambda$	$\lambda_{\max}$	$\lambda_{\text{trace}}$
Model A. Incorporating a trend drift in the error correction model, $r = 1$			
$\lambda_1$	0.101	29.54 (20.78) <sup>a</sup>	39.14 (29.51)
$\lambda_2$	0.032	8.93 (14.04)	9.61 (15.20)
$\lambda_3$	0.002	0.67 (3.96)	0.67 (3.96)
Model B. Incorporating a constant in the cointegrating vector of the error correction model, $r = 1$			
$\lambda_1^*$	0.113	33.30 (21.89)	46.77 (35.07)
$\lambda_2^*$	0.036	10.25 (15.75)	13.47 (20.17)
$\lambda_3^*$	0.012	3.22 (9.09)	3.22 (9.09)
$LR_\lambda = 3.97 < 5.99$ ( $df = 2; \alpha = 5\%$ )			

Note: <sup>a</sup> The figures in parentheses are critical values for  $\lambda_{\max}$  and  $\lambda_{\text{trace}}$  statistics at the 5 per cent significance level.

$$LR_\lambda = -T \sum_{i=r+1}^n [\ln(1 - \lambda_i^*) - \ln(1 - \lambda_i)], \quad (4)$$

where  $\lambda_i^*$  and  $\lambda_i$  are estimated eigenvalues of the matrix  $\Pi$  for the restricted (Model B) and unrestricted (Model A) models, respectively;  $r$  is the number of cointegrating vectors in the unrestricted model; and  $n$  is the number of endogenous variables.

Given that  $n = 3$  and  $r = 1$ , the computed value for  $LR_\lambda$  is 3.97, which is smaller than the critical value of 5.99 with two degrees of freedom at the 5 per cent significance level (bottom of table 3). Therefore, Model B (the restricted version) is not rejected and it is concluded that the ECM could be specified with a constant term in the cointegrating relation.

Based on the results from Model B, it is found that the residuals of the ECM are free from any form of autocorrelation but are not normal. Standard asymptotic chi-squared tests are, however, still valid when the ECM is estimated by the Johansen procedures (Gonzalo 1994).

As indicated in table 4, the estimated cointegrating relation or long-run equilibrium relationship, normalised by the  $\beta$  associated with the farm price, can be written as:

$$0.573 + FP - 0.879WP - 0.129RP = 0. \quad (5)$$

**Table 4** Estimated long-run parameters,  $\alpha$ s and  $\beta$ s,  $r = 1$ 

	<i>FPt</i>	<i>WPt</i>	<i>RPt</i>	Constant
Estimated $\beta$ s	1.000 (—) <sup>a</sup>	−0.879 (—)	−0.129 (—)	0.573 (—)
	− <i>FPt</i>	− <i>WPt</i>	− <i>RPt</i>	
Estimated $\alpha$ s	−0.199 (−3.94) <sup>b</sup>	0.008 (0.28)	0.042 (2.20)	

Notes: <sup>a</sup>  $t$ -ratios for  $\beta$  coefficients are not calculated.

<sup>b</sup> Figures in parentheses are  $t$ -values.

It can be seen that the estimated coefficients associated with *WP* and *RP* just about sum to one, implying constant returns in prices. That is, a 1 per cent increase in farm prices is associated with a 1 per cent increase in wholesale and retail prices. Moreover, since the three series are shown to be cointegrated, the system can be expected to respond to exogenous shocks and return to equilibrium after being perturbed. The speed of adjustment to the disequilibrium errors, which is indicated by the magnitude of the adjustment coefficients, is shown in the bottom half of table 4. It can be seen that the  $\alpha$  coefficient associated with the wholesale beef price is statistically insignificant. This result suggests that the wholesale price is weakly exogenous to the system.

The results from the ECM showed that the majority (about two-thirds) of estimated individual short-run coefficients (coefficients which are associated with lagged prices in first differences) are statistically insignificant at the 5 per cent level. However, joint significance tests show that lagged coefficients associated with a particular price variable when considered as a group are statistically significant, with three exceptions. The exceptions are that, as a group, the lagged farm prices in the farm price equation, the lagged retail prices in the wholesale price equation and the lagged wholesale prices in the retail price equation are not statistically significant. From the ECM, there are also indications of seasonality and structural breaks in the Australian beef market. The full results from the ECM, which are not reported here to save space, are available from the authors for interested readers.

## 5. Impulse response analysis

Impulse response or dynamic multiplier analysis is commonly used to investigate the interrelationships between variables in dynamic models. It is therefore a valuable tool in cointegrated systems (Lütkepohl and Reimers

1992). Since in such systems the deviations from equilibrium are stationary, any shock to the system will generate time paths which eventually settle down to a new equilibrium, provided no further shocks occur. These time paths of the variables provide insights into the short-run and long-run relations between the variables. The tool is particularly useful when there are two or more cointegrating relations, which may be difficult to interpret on economic grounds. In this section, impulse response and forecast error variance decompositions pertaining to the Australian beef prices are investigated.

The simulated time paths, or the impulse functions, are found by imposing a recursive structure on the moving average representation of the VAR model. As such, the order of the variables is of importance. The ordering can be determined based on *a priori* economic reasoning and/or some statistical properties of the system such as the correlation among residuals and possible exogeneity of the variables (Enders 1995, p. 309; Doan 1996, pp. 8–14). Because the wholesale price is found to be weakly exogenous to the cointegration analysis, the order *FP, RP, WP* is used. This ordering implies that shocks in the farm and retail prices will have an instantaneous effect on the wholesale price, while the wholesale price can only have a lagged impact on the other two prices.

Following Williams and Bewley (1993), the impulse response analysis is modelled in an unrestricted VAR based on the first differences of the log price series and the disequilibrium error in the previous period, i.e.  $ee_{t-1}$  calculated based on equation 5. The first variable in the sequence is thus the change in the log of the farm price, followed by the changes in the log of the retail and the wholesale prices. Therefore, the VAR used for the impulse response analysis is specified as:

$$x_t = A_1x_{t-1} + A_2x_{t-2} + \dots + A_7x_{t-7} + A_8ee_{t-1} + A_9D_t + A_{10}S_t + e_t \quad (6)$$

where:

$$x_t = [\Delta FP_t, \Delta RP_t, \Delta WP_t]';$$

$$x_{t-i} = [\Delta FP_{t-i}, \Delta RP_{t-i}, \Delta WP_{t-i}]' \text{ for } i = 1, 2, \dots, \text{ and } 7;$$

$$ee_{t-1} = 0.573 + FP_{t-1} - 0.879WP_{t-1} - 0.129RP_{t-1};$$

$A_s$  = parameters to be estimated; and

$e_t$  = white noise disturbance terms which may be contemporaneously correlated.

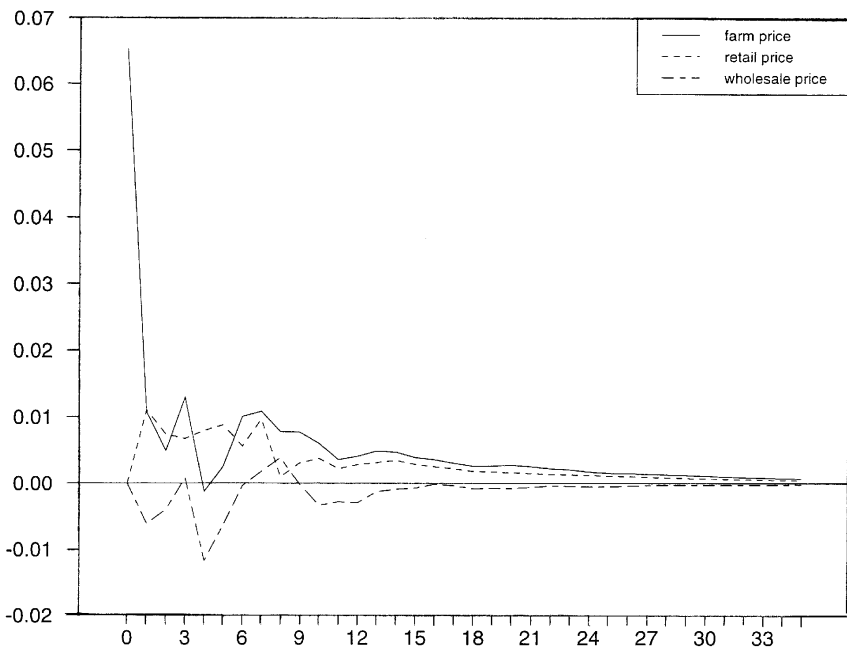
All other variables are as previously defined in equation 4.

Because the impulse response functions generated from equation 6 are from a cointegrated system, a shock in any one price is expected to have a

permanent effect on the system, which eventually settles down to a new equilibrium. This is evident from the impulse response functions presented in figures 5, 6 and 7. The simulation is based on 36 lags to allow for sufficient time for responses.

In figure 5, it is shown that a one unit (a one standard deviation) shock in the change in the farm price brought about an initial response of a 0.065 unit increase in its own price, but no change in either the retail or wholesale prices. Further, a one unit shock in the change in the wholesale price brought about an initial response of 0.03 unit, 0.02 unit and 0.004 unit increases in the wholesale, farm and retail prices, respectively (figure 6). Finally, a one unit shock in the change in the retail price brought about an initial response of a 0.023 unit increase in its own price and a 0.008 unit increase in the farm price (figure 7). It had no impact on the wholesale price, however. Despite the differing responses to shocks, the changes in all price series stabilise after fifteen months and settle down to approximately zero eventually, which is to be expected from a cointegrated system.

The basic results from the impulse response functions can also be seen from the forecast error variance decompositions. With a 36-month forecasting horizon used, the variance decompositions are reported in table 5. As expected, each price series explained the highest proportion of its own



**Figure 5** Impulse responses to an innovation in the farm price

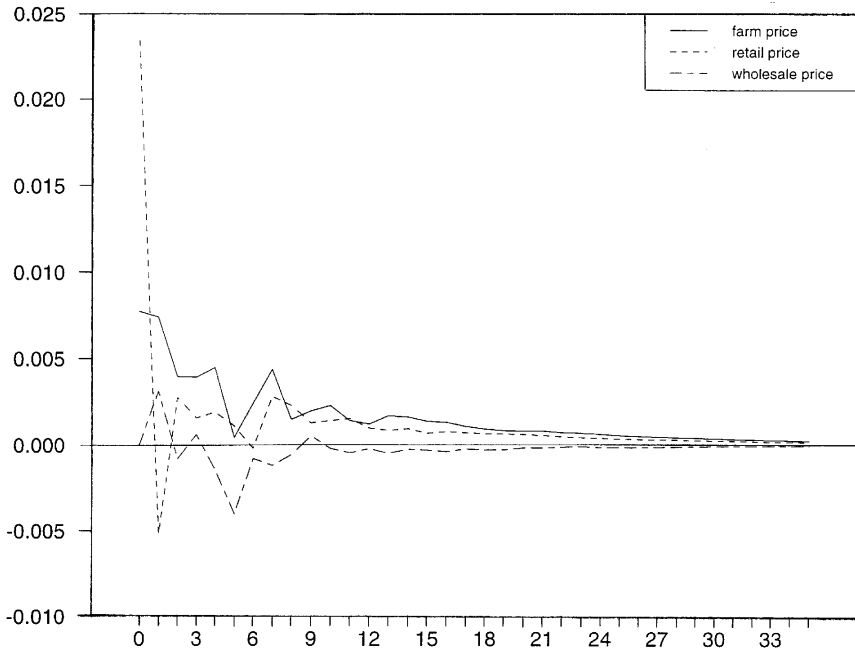


Figure 6 Impulse responses to an innovation in the retail price

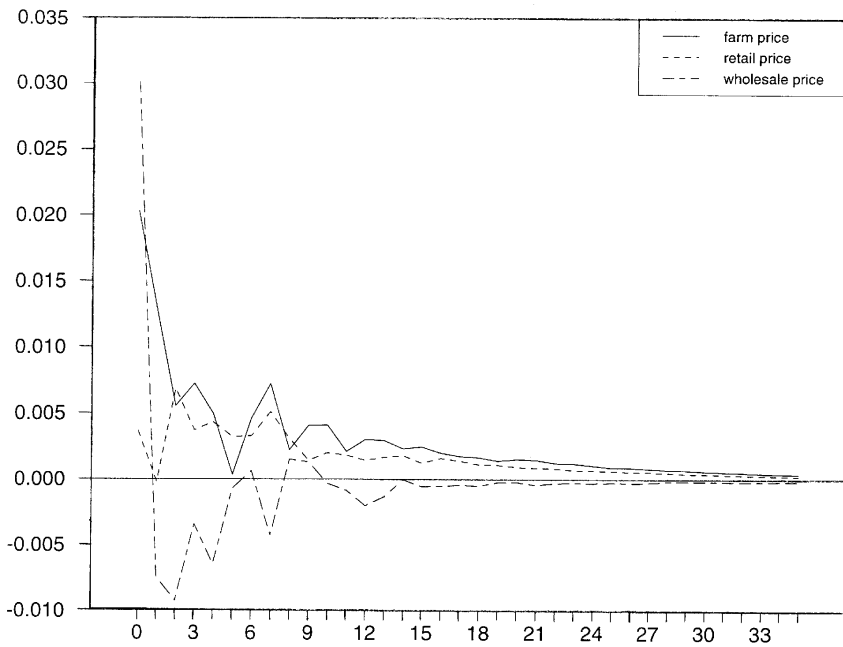


Figure 7 Impulse responses to an innovation in the wholesale price

**Table 5** Variance decomposition percentage of one-month and 36-month forecast error variance

	Percentage of forecast error variance explained by shocks in:					
	Farm price		Retail price		Wholesale price	
	one-month	36-month	one-month	36-month	one-month	36-month
Farm price	100	85.46	0	9.78	0	4.76
Retail price	9.81	25.37	90.19	70.81	0	3.82
Wholesale price	30.96	39.45	1.0	8.32	68.0	52.24

past values. The farm price, the retail price and the wholesale price explained 100, 90 and 68 per cent, respectively, of their own forecast error variance in the first month. The corresponding figures for 36 months ahead are 85.46, 70.81 and 52.24 per cent, respectively. The cross-price effects also show that the shock in the farm price can explain around 25 per cent and 40 per cent of variations in the retail and wholesale prices, respectively. By contrast, the shock in the retail price explains less than 10 per cent of variations in the farm and wholesale prices, while the shock in the wholesale price explains less than 5 per cent of variations in the farm and retail prices. These results suggest that shocks to the Australian beef market originate mostly from the farm level and that the farm price is most affected followed by the wholesale price and the retail price.

## 6. Discussion of results

The main results from the cointegration test suggest that Australian beef prices are cointegrated and that the wholesale price is weakly exogenous. This means that although a long-run relationship exists between the price variables, the wholesale beef price does not adjust to deviations from the long-run equilibrium. As such, the long-run equilibrium in the Australian beef market after a shock is restored by adjustments made at the retail and farm levels. Although the wholesale price adjusts only through the short-run dynamics in the system, it still plays a role in the price transmission process despite being found to be weakly exogenous.

The result that the wholesale price is weakly exogenous may be explained by price levelling practised by some marketers in the Australian beef market, which is found by Naughtin and Quilkey (1979) and by Griffith *et al.* (1991). Price levelling refers to the practice of retailers and wholesalers holding their selling prices relatively stable in the face of rising or falling input procurement costs. As a result, the impact of fluctuating input prices on prices charged to customers is smaller than it would otherwise be (Griffith



and Piggott 1994). According to Parish (1967), price levelling destabilises farm level prices while stabilising wholesale and retail prices. As a result, not only are farm prices more volatile than they would otherwise be, but the changes are not passed on to the higher levels of the market and do not influence the final demand for beef. Therefore, pricing efficiency is lessened.

Although price levelling and averaging are short-run phenomena and cannot persist in the long run, it is likely that such practices over a longer time period balance out any profit or loss which may exist in the short run. Indeed, Griffith and Piggott (1994) show that there are nearly the same number of price increases as price decreases in the beef price data over the period 1971–91. What this means is that if the frequency of price increases is the same as price decreases, then holding prices relatively constant would not result in persistent gains or losses so long as the changes in either direction are of similar magnitude. Search costs and preferences for one-stop shopping on the part of consumers could also contribute to market imperfection of this sort.

Using Canadian data, Larue (1991) and Larue and Babula (1994) find that the farm output price and the retail food price are weakly exogenous. They explained the findings by demand-pull and cost-push tendencies which dictate the causal relations between input and output prices. Changes in demand structure and increased concentration in the retail market have also been suggested as possible factors influencing the results. Since the main purpose of the current analysis is to determine whether long-run relationships exist between the Australian beef prices, factors which contribute to a structural break in the 1970s and to weak exogeneity of the wholesale price are not fully investigated here. Instead, they are areas for further research. As such, this article should be considered a first attempt at understanding the long-run dynamics in the Australian meat prices based on cointegration theory, as opposed to the traditional mark-up models which focus on short-run dynamics. Because beef is the largest component of domestic meat consumption, developments in the beef market are expected to influence prices in related meat markets. Therefore, the techniques employed here can be used to analyse the relationships between beef and other meat prices.

## 7. Conclusion

Long-run relationships between Australian beef prices at the farm, wholesale and retail levels are expected to exist because they are linked through derived demand. This price link means that, in a competitive market, not only do they tend to move together over time but respond to the same shocks in the system, although to varying degrees. Moreover, the price differentials at

different levels should reflect the cost of providing marketing services when the product is moved from the farm to the final consumers. The existence of a long-run equilibrium relationship, in econometric terms, means that these price series are cointegrated. That is, there exists a linear combination of non-stationary variables that is stationary and they respond to deviations from the equilibrium.

Using Johansen's procedure, the cointegration results show that all three prices considered are integrated of order one and they are cointegrated. Therefore, it is concluded that a long-run relationship does exist between Australian beef prices at the farm, wholesale and retail levels. There is also evidence that the price series are influenced by seasonality, and in addition, by a structural break in the 1970s. Furthermore, the wholesale price is found to be weakly exogenous. The latter result may be an indication of market inefficiency due in part to price levelling behaviour being observed at the wholesale level.

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