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The Long Run, Market Power and Retail Pricing

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Abstract

The paper uses the Johansen cointegration approach to analyse long-run pricing strategies of pork and chicken retailers in Austria. Long-run retail pricing strategy is found to be dependent on market share and price elasticity of demand for product. A combination of mark-up pricing strategy for pork and a competitive pricing strategy for chicken is considered by retailers to yield maximum profit. Long-run price adjustment reveals linkages to pricing strategy. The versatility of the Johansen cointegration technique as a tool capable of analysing both competitive and imperfect market situations is also revealed. The paper recommends meat policy to be product specific rather than holistic.

Keywords

Market power, markup pricing, cointegration, long run

JEL Classifications

C32, D43, L11, Q13

Comments

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Contents

Introduction 1

2. The Meat Market 2

2.1 Policy Spectrum 2

2.2 Market Structure 3

3. Methodology 5

3.1 Price Determination 5

3.2 The Model 7

4. Results and Discussions 9

4.1 Data 9

4.2 Cointegration Analysis 10

4.3 Structural Tests 13

5. Summary and Conclusion 14

References 15

Introduction

Studies on long-run relationships relate economic variables in situations where any adjustments of the variables to positions of disequilibrium are assumed to have been completed (Charemza and Deadman, 1992, p. 57). It is, however, important to distinguish between the two types of long-run relationships – *static equilibrium* where the variables are assumed to be unchanging between periods, and *stable equilibrium* where all variables are changing at some constant rate. Most previous studies concerning long-run relationships between producer and retail prices are based on static equilibrium models of firm behaviour (e.g., Gardner 1975, Wohlgenant 1989, Griffith and Moore, 1991; McCorrison and Rayner, 1998). These studies assume equality of supply and demand in the respective producer, retail and marketing-input markets. Due to the existence of short-run disequilibrium in markets and the fact that time is required for markets to clear, Heien (1980) points out that the static equilibrium model is not always appropriate since it has little to say about the time path from one equilibrium point to another. Especially, in oligopolistic markets, where a change in the behaviour of one firm is likely to lead to unspecified responses of rival firms, the entire future time path of choice variables changes as current decisions are made. For this reason, such markets are expected to lend themselves more easily to dynamic analysis. Beginning with Larue (1991), several recent studies have taken a stable (dynamic) equilibrium approach based on cointegration techniques. The advantage of the cointegration framework over other approaches is that it permits the joint modelling of short-run economic reactions, trends, long-run equilibria as well as the speed of adjustment.

Usually, the application of cointegration techniques to studies involving relationships between farm (producer) and retail prices is implicitly based on perfectly competitive market assumptions (e.g., Larue, op. cit.; Palaskas, 1995; 1996). Related studies in which imperfect market assumptions are invoked have employed other approaches such as structural methods (e.g., Azzam and Pagoulatos, 1990; Holloway, 1991; Hyde and Perloff, 1998) and reduced-form techniques (e.g., Panzar and Rose, 1987; Hall, 1988; Zhoa et al., 1996) – among others. Barrett (1996) argues that the implicit perfectly competitive market assumption is flawed in that even if price differences exactly equal transfer costs, one cannot reasonably presume perfect competition, since this is equally consistent with monopolistic limit pricing, with collusive pricing by a spatial oligopoly (Faminow and Benson, 1990) or with Pareto inferior trade (Newbery and Stiglitz, 1984). Using linear-quadratic models, however, Karp and Perloff (1989; 1993) have shown that the dynamic behaviour of oligopolistic markets is relatively more competitive than collusive. Theoretical support for this evidence includes Eichner (1973), who suggests three possible causes for this behaviour: larger elastic demand in the medium to longer term due to substitution effects; the probability of new firms invading the market; the fear of some sort of detrimental government intervention.

The current paper is motivated by the theoretical arguments of Barrett (op. cit.) and the empirical evidence provided by Karp and Perloff (op. cit.). The paper examines the existence of market power in the respective pork and poultry¹ (“white” meat) meat markets in Austria by means of the Johansen cointegration technique. Verification of the market power parameters is made possible by imposing identifying restrictions on the cointegration space related to the respective price equations and testing whether restrictions implied by competition can be rejected. Presumably for dietary reasons, the demand preferences for meat in Austria have been shifting increasingly from beef and veal to pork and poultry since the 1970’s (Handschr, 1991). This observation provides some rationale for studying the pricing strategies of pork and poultry retailers and their implications for policy. At the heart of the problem is the fact that the Austrian food retail sector is one of the most concentrated in the OECD countries – currently placing third in Europe, after Finland and Switzerland. Thus, it can be envisaged that meat retailers exercise a high degree of market power. Also, the long run is considered because under imperfect competition, retailers are more likely to establish a price with long-run rather than short-run profits in mind and to maintain that price for a considerable period of time.

The organization of the paper is as follows: section 2 describes the market characteristics; section 3 expounds the model employed in the study and summarises the Johansen maximum-likelihood (ML) technique; section 4 interprets the empirical results and section 5 concludes.

2. The Meat Market

2.1 Policy Spectrum

In Austria, livestock accounts for about 63% of the final output of agriculture. Meat also accounts for about 20% of the household budget for food. Policy makers have, therefore, increasingly shown concern for the price of meat paid by the consumer and that received at other stages in the marketing chain. In particular, policy makers worry that the incidence of imperfect competition, especially at the retail level, might prevent prompt and accurate adjustment in supply and demand conditions from one market level to another. Both regional and national policies have been engaged to alleviate these worries. For example, during Austria’s pre-EU membership negotiations, it was expected that two of the main elements of EU membership – participation in the Common Agricultural Policy and trade liberalization with the EU – would bring about lower producer and retail prices as well as a more accurate transmission of prices from one market level to another. Failure to realize these expectations would imply the existence of price signal distortion in the Austrian meat market and departure from EU membership objective. In 1998, the government demonstrated its

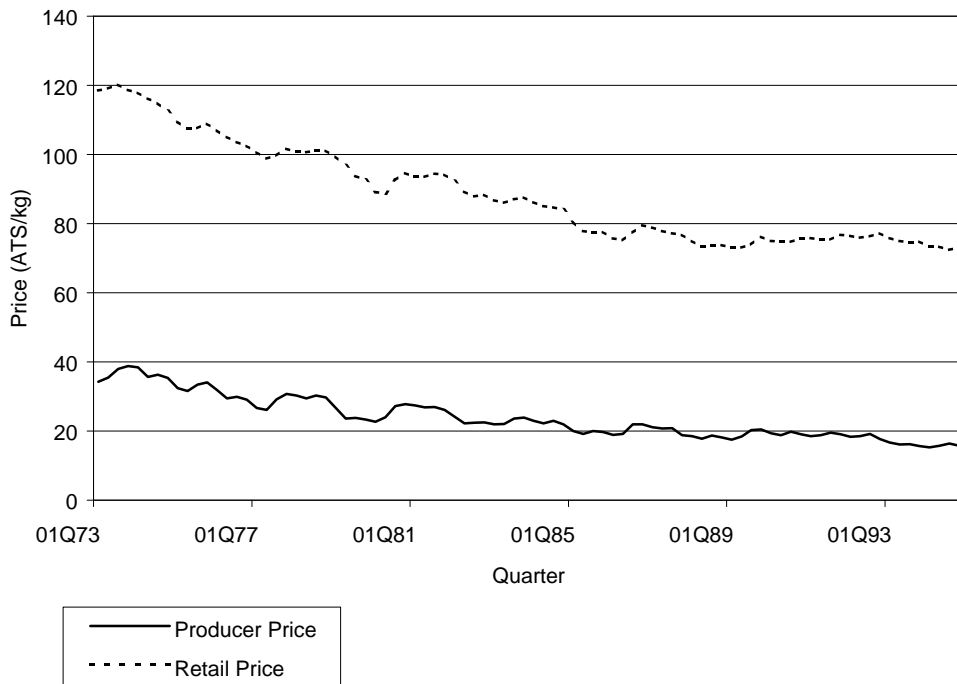
¹ The term poultry as used in the paper comprises poultry, geese, duck, and turkey meat.

persistent effort towards achieving a more competitive domestic retail environment by preventing a take over bid of the Julius Meinl food retail chain (6.9% market share, 1997) by the German group Rewe which already owns Billa-Merkur-Mondo retail outlets (32.5% market share, 1997).

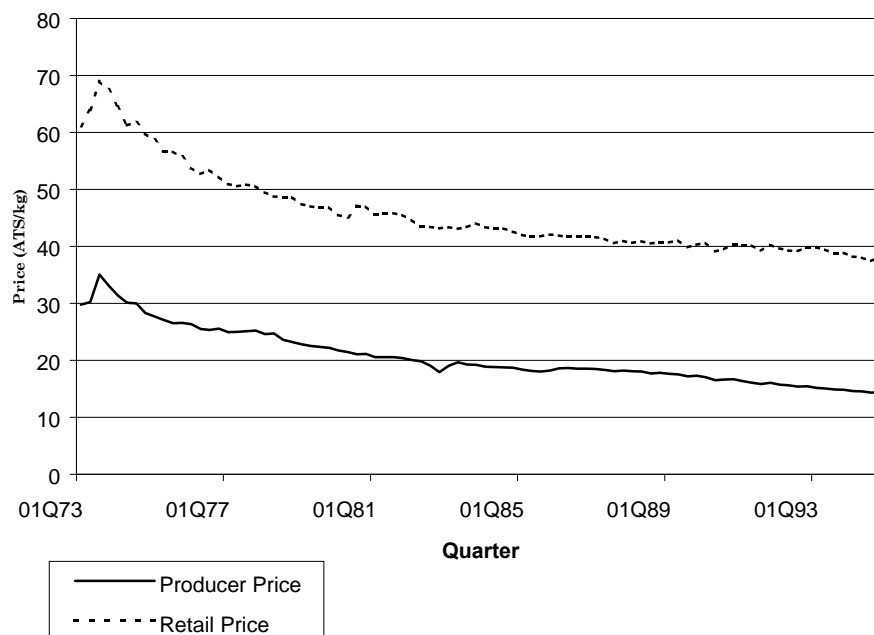
2.2 Market Structure

Figures 1 and 2 show that the trend of real² farm and retail prices of meat in Austria have been falling gradually since the early eighties. These downward price movements may be explained by improvements in the productivity of breeding stocks and in the performance of feeder livestock (Jumah and Kunst, 1996). Other factors that might have influenced these price trends are improvements in labour productivity and the efficiency of marketing services.

Figure 1: Real Producer and Retail Prices for Pork in ATS/kg (1973:1-1994:4)

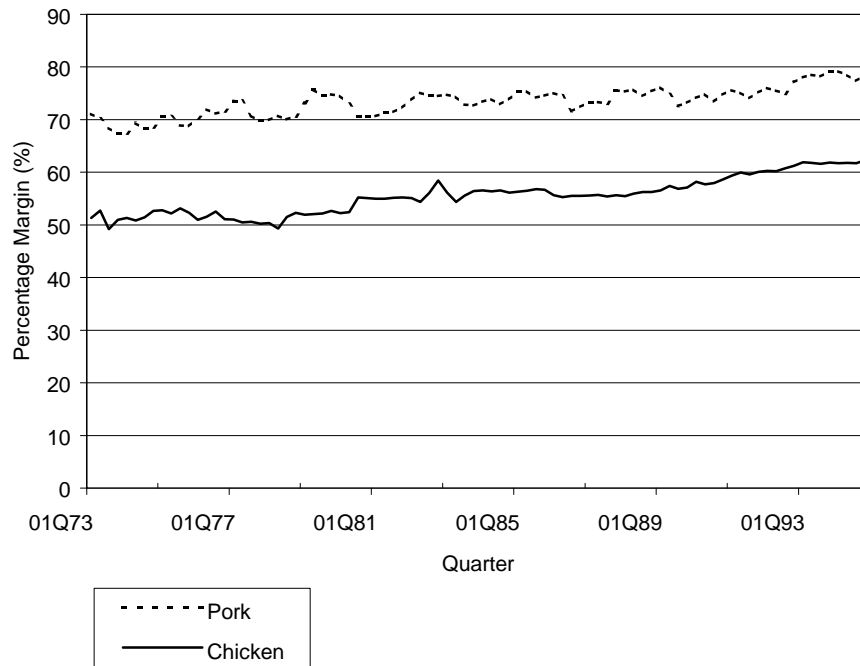


² Real prices are prices deflated by the consumer price index (CPI) using 1986 as base year.

Figure 2: Real Producer and Retail Prices for Poultry in ATS/kg (1973:1-1994:4)

As can be seen in Figure 3, percentage margins³ in the Austrian meat market are high. Arguably, a fraction of these margins may be attributed to the exercise of market power as afforded by the high concentration in the meat retail sector. In a sample survey of consumers by Scheikl (1989, p. 52) on the Austrian meat market, 44% of the respondents were found to purchase meat from supermarkets. In 1994 large chain supermarkets accounted for about 66% of total meat sales in Austria. Also, the proportion of the Austrian market occupied – in terms of market sales – by the top four food chains or supermarkets increased from 75.2% in 1987 to 80.7% in 1994, indicating that a high degree of concentration exists in the Austrian meat retail sector. Digby (1989), argues that with some degree of market power, marketing chains and supermarkets may adopt a ‘wait and see approach’ in the short run and retain all or part of the benefits of cost decreases as added profits while passing on higher input costs immediately to consumers and producers. Similarly, when retailers have market power and are faced with increasing demand for a particular product or group of products, they will be able to reap super normal profits in the long run.

³ Defined as the ratio of the marketing margin to the retail price expressed in percentages.

Figure 3: Percentage Margin for Pork and Poultry (1973:1-1994:4)

The wholesale sector exhibits a structure analogous to that of the retail sector. For example, Pretterhofer et al. (1996) have shown that concentration in the meat wholesale (slaughterhouse) sector is quite high, with 6% of the firms accounting for about 50% of total wholesale sales per week in 1995. Similar to the observation of Hall et al. (1979) on the US beef market, however, the slaughterhouses face both intra-industry competitive pressures in obtaining livestock and retail competitive pressures due to the existence of substitute supply outlets, such that most of the firms are unable to operate at maximum capacity. In contrast to the retail sector, therefore, the wholesale sector is expected to operate on a very small margin.

3. Methodology

3.1 Price Determination

When dealing with prices at different marketing levels such as farm and retail prices where M is the marketing margin, while P_f represents the price of the farm component of a retail good priced at P_r , then the relationship between the three variables can be represented linearly as:

$$P_r = M + P_f \quad (1)$$

Following Tomek and Robinson (1995 p. 44), the margin M is generally specified as the linear combination of a constant absolute amount c and a percentage (markup) a of the retail price such that:

$$M = c + aP_r \quad (2)$$

where $0 \leq c$ and $0 \leq a < 1$.

It follows from equations (1) and (2) that

$$P_r = c + aP_r + P_f \quad (3)$$

$$\Rightarrow P_r = \frac{1}{1-a}c + \frac{1}{1-a}P_f \quad (4)$$

If $a = 0$, equation (4) reduces to the case of the constant absolute margin:

$$M = c = P_r - P_f \quad (5)$$

Equation (4) can be rewritten in the reduced form as:

$$P_r = c^* + a^*P_f \quad (6)$$

where $c^* = \frac{1}{1-a}c$ and $a^* = \frac{1}{1-a}$.

When retailers have market power, they influence prices strategically. As a result, they are able to hold prices above marginal costs by charging retail markups. Thus, under a situation of market power, it is expected that a is strictly greater than zero (i.e., $0 < a < 1$) so that $a^* > 1$ in equation (6). On the other hand, competitive firms bid prices down to marginal cost levels. Consequently, equation (5) will correspond to the perfectly competitive case where $a = 0$.

Also, in Figure 3, percentage margin defined as $\frac{M}{P_r}$ becomes

$\frac{c + aP_r}{P_r} = \frac{c}{P_r} + a$, for retailers with market power and $\frac{c}{P_r}$, for retailers facing perfect competition.

In case P_r and P_f are nonstationary in the regression equation:

$$P_{r,t} = c_t^* + a^* P_{f,t} + e_t \quad (7)$$

e_t must be stationary if equation (7) is true in the long run. Given that both producer and retail prices follow non-stationary processes as confirmed by the results in Tables 1 and 2, equation (7) may be estimated using cointegration techniques that properly account for non-stationarity properties.

Equation (7) is a representation of the first step of the Engle-Granger (EG) (1987) cointegration technique. The EG technique examines bivariate relationships and a subsequent test for stationarity of the residuals by means of ordinary least squares. In a multivariate system, the EG approach leads to a partial analysis. The limitations of the EG approach have been addressed by Hendry and Mizon (1992) and Banerjee et al. (1993) – among others. Hendry and Mizon (op. cit.) emphasize that under certain conditions, cointegration analysis within a partial system might be inefficient. Johansen (1988, 1991, 1992) provides a more suitable way of testing for cointegration within a multivariate (also bivariate) framework by means of ML techniques. Moreover, Johansen and Juselius (1992) have shown that the ML framework is well suited to the issue of price convergence.

3.2 The Model

Cointegration relationships between the respective farm and retail prices of pork and poultry (P_{pf} and P_{cf} ; P_{pr} and P_{cr}) in Austria are estimated using the Johansen ML procedure taking the form of a vector error correction (VEC) representation:

$$\Delta p_t = \sum_{i=1}^k \Gamma_i \Delta p_{t-i} + \Pi \tilde{p}'_{t-k-1} + \Phi D_t + e_t \quad (8)$$

where the reduced rank of the $n \times n$ matrix of Π equals the number of cointegration vectors $r \leq n - 1$ in the system and n equals the number of series in question. Thus, Π can be written as $\Pi = \mathbf{a} \begin{pmatrix} \mathbf{b} \\ \mathbf{m} \end{pmatrix}'$, where \mathbf{a} and \mathbf{b} are both of the dimension $n \times r$ and rank r . The constant \mathbf{m} is a scalar restricted to the cointegration space and $\tilde{p}'_{t-k-1} = (p'_{t-k-1}, 1)$. This specification assumes that there are no linear trends in the levels of the data as roughly depicted by Figures 1 and 2. The effect of \mathbf{m} is to capture the constant absolute portion of the long-run marketing margin. The matrix \mathbf{b} contains the cointegrating vectors \mathbf{b}_i , i.e., $\mathbf{b} = (\mathbf{b}_1, \dots, \mathbf{b}_r)$ while the matrix of the adjustment coefficients \mathbf{a} describes the speed of adjustment of the particular series $\Delta p_{i,t}$ to deviations from the cointegration relationships.

Choice of the appropriate lag-length k was based on model selection criteria and also on the issue of whether the residuals are Gaussian, i.e., $\mathbf{e}_t \sim N(0, \Lambda)$, conditional on a set of $I(0)$ variables, D_t that are weakly exogenous. Swine production exhibits seasonal patterns – the so-called hog cycle. Turkey consumption also rises at Christmas. One would therefore expect some seasonal features in the data. In the analysis, D_t consists of centered seasonal dummies which are used to account for any deterministic seasonal behaviour that may be present in the price data and a point dummy variable (in the case of poultry) to account for an outlier for the period 1973:3. Centered seasonal dummies sum to zero over time and thus, they do not affect the asymptotic distributions upon which statistical tests depend.

Having established the number of cointegrating vectors, tests of structural hypotheses (to verify the long-run market behaviour or retail pricing strategy) on the cointegrating vectors are performed. Following Johansen and Juselius (op. cit.), the combined hypotheses $H_b: \mathbf{b} = H_i \Phi_i$ for $i = 1, \dots, r$ are tested, with fixed $n \times s_i$ matrices H_i indicating the linear economic hypothesis to be tested on each of the r cointegration relations and the $s_i \times 1$ vectors Φ_i of freely estimated parameters:

$$\mathbf{b} = (\mathbf{b}_1, \mathbf{b}_2, \dots, \mathbf{b}_r) = (H_1 \Phi_1, H_2 \Phi_2, \dots, H_r \Phi_r)$$

The Likelihood Ratio (LR)-statistics for the combined hypotheses, given the number of cointegrating vectors, is asymptotically standard χ^2_m -distributed, where m denotes the number of over-identifying restrictions on \mathbf{b} . Beginning with a just-identifying restriction (i.e., normalizing on the coefficient relating to the retail price) to achieve identification along the lines of equation (6), an over-identifying restriction (i.e., homogeneity restriction: $\mathbf{b}_{retail} = -\mathbf{b}_{farm}$ or $a^* = 1$ in equation (6)) is imposed on the a^* coefficient in order to verify whether $a = 0$ and that the pricing behaviour of retail firms corresponds to the perfectly competitive case as exemplified by equation (5).

4. Results and Discussions

4.1 Data

Quarterly producer and retail prices of pork and poultry for the period 1973:1 to 1994:4 were used. 1973:1 represented the period in which prices net of value added tax began to be recorded, and 1994:4 represented the last quarter prior to Austria's entry into the EU and as such, the end of extreme border protection, trade management and possibly oligopolistic behaviour of meat retailers. The retail values of carcass were calculated from price data on a range of cuts, using their respective weights in the consumer price index⁴ and scaling these weights such that they add up to unity. This should make the resulting retail prices comparable to producer prices.⁵ All data are in real values in order to isolate the analysis from inflationary factors. The data were kindly provided by Karl M. Ortner of the Federal Institute of Agricultural Economics, Vienna⁶ and are described in the ALFIS data bank. The data were collected by Statistik Österreich as part of the Austrian Consumer Price Index survey.

Following Engle, Granger and Hallman's (1989) suggestion with respect to the effects on integration and cointegration of using series containing strong seasonal components, the series in question were subjected to seasonal unit root tests. Referring to Table 1, pork prices were found to have unit roots at the zero frequency, whilst poultry prices featured unit roots at the semiannual frequency. None of the price series, however, featured unit roots at the annual frequency.

Table 1: Results of seasonal unit root tests for price series

Series	Constant	$t(\mathbf{g}_1)$	$t(\mathbf{g}_2)$	$t(\mathbf{g}_3)$	$t(\mathbf{g}_4)$	$F(\mathbf{g}_3 \cap \mathbf{g}_4)$
Pork farm price	-1.847	-0.270	4.311**	4.367**	1.029	18.972**
Pork retail price	-2.968	-0.650	4.507**	5.003**	1.297	30.286**
Poultry farm price	-3.806	6.577**	-0.140	6.051**	-4.527**	25.492**
Poultry retail price	-7.183	7.317**	-1.423	3.656**	-5.456**	17.732**

** indicates rejection of null hypothesis at the 5% level.

Null hypothesis: $\mathbf{g}_1 = 0, \mathbf{g}_2 = 0, \mathbf{g}_3 \cap \mathbf{g}_4 = 0$

⁴ The factors used were 0.253 for the flank, 0.305 for ham, 0.217 for the shoulder butt and 0.299 for the loin – in the case of pork – according to the weight of these cuts in the consumer price index.

⁵ To compute exactly comparable prices, one would need the retail prices for all cuts which are made from carcasses. The cuts for which prices are available were assumed to be representative of those, but by this assumption retail prices are likely to have been overestimated.

⁶ Bundesanstalt für Agrarwirtschaft.

Engle and Granger (op. cit.) argue that, if two series of interest contain unit roots at the zero frequency and if the existence of seasonal unit roots cannot be rejected in at least one of the series, the possible presence of seasonal unit roots means that the standard EG cointegration procedure lacks consistency. In order to make the analysis for poultry comparable on the same frequency as that of pork, the unit root components at the semi-annual frequency were filtered out with the following filter suggested by Hylleberg et al. (1990):

$$x_t = S(B)_2 P_{c,t}$$

where $S(B)_2 = (1 - B^4) / (1 - B)(1 + B^2) = (1 + B)$ and x_t is the filtered series.

Table 2 presents the results of the seasonal unit root tests performed on the filtered farm and retail prices of poultry. Both filtered price series have unit roots at the zero frequency but no unit roots at the semi-annual and annual frequencies.

Table 2: Results of seasonal unit root tests for filtered poultry price series

Series	Constant	$t(\mathbf{g}_1)$	$t(\mathbf{g}_2)$	$t(\mathbf{g}_3)$	$t(\mathbf{g}_4)$	$F(\mathbf{g}_3 \cap \mathbf{g}_4)$
Poultry farm price	-2.510	-0.354	3.418**	5.173**	0.795	18.952**
Poultry retail price	-5.215	-1.783	5.314**	3.092**	2.232**	11.008**

** indicate rejection of null hypothesis at the 5% level.

Null hypothesis: $\mathbf{g}_1 = 0, \mathbf{g}_2 = 0, \mathbf{g}_3 \cap \mathbf{g}_4 = 0$.

4.2 Cointegration Analysis

Results of the cointegration analysis are presented in Tables 3 and 4. Cheung and Lai (1993) suggest that the trace test shows more robustness to both skewness and excess kurtosis in (the residuals) than the maximum eigenvalue (I_{\max}) test. Following this proposition, the choice of the rank r was based on the trace test. Choice of the lag length was based on Schwarz Bayesian criterion (SBC). According to the SBC, a lag length of one was found to be appropriate for both cointegration analyses. Also, there was only one cointegration vector at the 1% level of significance in both analyses.

In the case of poultry, the Ljung-Box (LB) test for residual autocorrelation which is based on the estimated auto- and crosscorrelations of the first $(T/4)^7$ lags, rejected null hypothesis of no autocorrelation. The null hypothesis that the residuals are normally distributed was also

⁷ T is the length of the series.

rejected. All this could have been due to the heterogeneity of this category (see also Jumah and Kunst, op. cit.). The reliability of the estimates for poultry prices is therefore based on the significance of the Lagrange Multiplier (LM)-type tests for first and fourth order autocorrelation. Although the estimation results would be more reliable if the overall model conformed to the assumptions that the disturbances are serially independent and normally distributed, the first point is the more critical one.

Residual autocorrelation was strong in the pork price model. The presence of large residual autocorrelation could imply that there are other unobserved variables that could explain the model. It would be expected that conditioning on the labour variable would correct residual misspecification, since labour costs influence the marketing chain. However, autocorrelation in the residuals persisted after conditioning on the labour variable.

Johansen (1992) has shown that, if in equation (8), for example, the farm price of pork ($\Delta P_{pf,t}$) does not respond to deviations from the long-run equilibrium $\begin{pmatrix} \mathbf{b} \\ \mathbf{m} \end{pmatrix}' \tilde{p}_{t-k-1}$, i.e., if the corresponding rows of the short-run adjustment coefficients \mathbf{a} are zero, then $\Delta P_{pf,t}$ is weakly exogenous with respect to \mathbf{a} and \mathbf{b} (see also Engle *et al.* 1983). Having found $\Delta P_{pf,t}$ to be weakly exogenous to the system as indicated in Table 4, equation (8) was modified by handling $\Delta P_{pf,t}$ as a conditioning variable⁸. Characteristically, equation (8) is split into the partial model:

$$\Delta P_{pr,t} = \Gamma_0 \Delta P_{pf,t} + \sum_{i=1}^k \Gamma_i^{pr} \Delta p_{t-i} + \mathbf{a}^{pr} \begin{pmatrix} \mathbf{b} \\ \mathbf{m} \end{pmatrix}' \tilde{p}_{t-k-1} + \Phi^{pr} D_t + \mathbf{e}_t^{pr} \quad (9)$$

for pork retail price $\Delta P_{pr,t}$ only – where the dynamics of the latter is conditioned on $\Delta P_{pf,t}$ – and the marginal model:

$$\Delta P_{pf,t} = \sum_{i=1}^k \Gamma_i^{Pf} \Delta p_{t-i} + \Phi^{Pf} D_t + \mathbf{e}_t^{Pf} \quad (10)$$

for the farm price of pork $\Delta P_{pf,t}$. Under such circumstances, cointegration analysis might be equivalently based on the partial model for $\Delta P_{pr,t}$. Urbain (1995) has shown that estimating equation (9) might result in considerable efficiency gains (especially, in large systems) as compared to equation (8) and it has the additional advantage that the estimation does not

⁸ Conditioning in this sense means that $\Delta P_{pf,t}$ enters the cointegration space, i.e., it appears in the long-run model.

rely on the correct specification of the process for the conditioning variable $\Delta P_{pf,t}$. Equation (9) was therefore used instead of equation (8) in the case of pork.

Table 3: Results of zero frequency cointegration tests

Item	Lag length	Trace statistic [#]	β	Constant	Q(21)	LM(1)	LM(4)	Normality test	
								c^2	P-value
Pork	1	50.94 4.55*	2.113	35.463	0	0	0	$c^2(4) = 9.015$	0.06
Pork [§]	1	–	2.131	35.257	0.06	0.04	0.11	$c^2(2) = 2.923$	0.23
Poultry	1	65.31 3.91*	0.873	48.926	0	0.36	0.10	$c^2(4) = 72.35$	0.00

[§] Partial model for pork.

[#] For critical values of trace statistics see Osterwald-Lenum (1992)

* significant at the 1%.

Table 4: Results of tests for weak exogeneity

Item	Test	LR statistic	P-value
Pork	$\mathbf{a}_{rp} = 0$	$c^2(1) = 31.34$	0.00
Pork	$\mathbf{a}_{fp} = 0$	$c^2(1) = 1.51$	0.22
Poultry	$\mathbf{a}_{rc} = 0$	$c^2(1) = 57.04$	0.00
Poultry	$\mathbf{a}_{fc} = 0$	$c^2(1) = 36.83$	0.00

The cointegration vectors are presented in normalised form, by setting the first element (i.e., the retail price) of \mathbf{b}' equal to unity. Based on these normalised cointegration vectors, equation (6) can be derived for pork as:

$$P_{pr,t} = 35.257 + 2.131P_{pf,t} \tag{11}$$

and that for poultry as:

$$P_{cr,t}^* = 48.926 + 0.873P_{cf,t}^* \tag{12}$$

where $P_{cr,t}^* = P_{cr,t} + P_{cr,t-1}$ and $P_{cf,t}^* = P_{cf,t} + P_{cf,t-1}$, i.e., the filtered price series.

4.3 Structural Tests

Table 5 presents the results of the structural test as to whether the restriction implied by competition can be rejected. The test which is performed by imposing over-identifying restrictions on the cointegration space related to equations (8) and (9) – based on the hypothesis that $\mathbf{b}_{retail} = -\mathbf{b}_{farm}$ in equations (8) and (9) or $a^* = 1$ in equation (6) – is implemented using:

$$H:\mathbf{b} = (H,\mathbf{j}) \tag{13}$$

(see, Johansen and Juselius, 1992), where H is $(n \times r)$ with known elements, and \mathbf{j} is the $(s \times r)$ matrix of unknown parameters. s represents the number of unrestricted parameters. The H matrix for the hypothesis involving one restriction is given by:

$$H = \begin{bmatrix} 1 & 0 \\ -1 & 0 \\ 0 & 1 \end{bmatrix} \tag{14}$$

The resulting LR test is based on the χ^2_m – distribution, where m denotes the number of over-identifying restrictions. The results indicate that the null hypothesis is rejected in the case of pork but accepted in the case of poultry.

Table 5: Results of the tests for overidentifying restrictions

Item	Test	P_{retail}	P_{farm}	Constant	LR statistic	P-value
Pork	$H = \beta$	1	-1	-54.325	$\chi^2(1) = 21.34$	0.00
Poultry	$H = \beta$	1	-1	-47.555	$\chi^2(1) = 0.30$	0.59

This implies that equation (2) is valid for pork; it also means that $\frac{1}{1-a} = 2.131$ and $a = 0.531$ in the case of pork. Having solved for a , c can be derived from $c^* = 35.257$, which implies that $c = 16.536$. Thus, the long-run margin equation for pork in the reference period was:

$$M = 16.536 + 0.531P_r \tag{15}$$

Retailers charged a mark up of about half the retail price for pork plus an absolute constant margin of 16.54 ATS in the long run.

For poultry, the results indicate that the price relationship followed equation (5) – the constant absolute margin. Based on the estimated equations, the long-run price equation for poultry was:

$$P_r \approx 24.46 + P_f \quad (16)$$

5. Summary and Conclusion

The study uses the Johansen cointegration approach to analyse long-run pricing strategies of pork and poultry retailers in their isolated markets in Austria. Retailers charged pork consumers a markup of about twice the farm price or half the retail price of pork in addition to a constant margin. On the other hand, mark-up pricing was virtually absent in the poultry market. This is translated as the existence of market power in the pork retail market and competitive behaviour in the poultry retail market.

It is, however, unclear what the driving forces behind the pricing strategies are. Presumably, retailers base their pricing strategies on relative demand shares for product. Pork accounted for about 57% of the total meat consumption in Austria whilst poultry meat accounted for 14% in the reference period (Handshur, *op. cit.*).

A comparison of the results to a former study in Austria (Jumah, 1996) and similar studies in Australia (McDonald and Spindler, 1987; Griffith et al., 1991; Hyde and Perloff, *op. cit.*) showed that analysing several markets simultaneously rather than examining markets in isolation is likely to reveal evidence of competitiveness when the alternative evidence for the existence of market power is true.

The paper also revealed the versatility of the Johansen cointegration technique as a tool capable of analysing both competitive and imperfect market situations. The results indicate that meat policy should be product specific rather than holistic.

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