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Buyer Market Power in UK Food Retailing

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Abstract
The potential existence of buyer market power in UK food retailing has attracted the scrutiny of the UK's anti-trust authorities, culminating in the decision to launch the second of two comprehensive regulatory inquiries in recent years. Throughout, detection of buyer power has been dogged by the paucity of reliable evidence of its existence. In this paper we present a simple theoretical model of oligopsony which delivers quasi-reduced form retailer-producer pricing equations in which the presence of market power can be detected using readily available market data. Using a cointegrated vector autoregression, we find empirical results that are consistent with the presence of oligopsony power in all six food products investigated.

Key words: Buyer power, Cointegrated VARs, UK food industry

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Introduction

In common with many national retail food markets in Europe, the rising degree of market concentration in the UK food sector has been a cause of concern to both consumer groups and food producers in recent years. By 2005, the four leading food retailers in the UK had a combined share of the grocery market of around 75 per cent, with the largest of these accounting for around one-third of all food sales (Office of Fair Trading, 2006). The issue has also aroused the attention of the UK's principal anti-trust authority, the Competition Commission, which published a report of its first statutory inquiry in 2000. A key issue highlighted in that report was the extent to which retailers can exert buyer power over their suppliers and the impact this has on consumer choice and competition in the food chain (Competition Commission 2000). It concluded that while there was only limited potential for abuse of market power with respect to consumers, there were grounds for significant concern regarding food retailers' relationships with suppliers, highlighting 27 oligopsonistic practices that specifically gave cause for concern. Despite the subsequent imposition of a Supermarket Code of Practice in 2002 effectively outlawing such practices, concerns over buyer power remain in the Office of Fair Trading's recent decision to refer the supermarkets to a further Competition Commission inquiry (Office of Fair Trading, 2006).

These concerns were most cogently illustrated by the nature of trading between retailers and suppliers of “fresh” food products in that 'Generally, suppliers of fresh produce appear to be most dependent on their largest main party customers [big supermarkets] for their sales' (Competition Commission 2000 11.15, p232) and '. . . most suppliers of fresh fruit and vegetables meat and poultry . . . appear to concentrate on trade with a limited number of suppliers (often four or less)' (Competition Commission 2000 11.8 p.231). Indicative figures from the food industry underline this reliance with some 75%1 of total UK output of apples and 80%2 of total UK fresh potato output being sold to the supermarkets. Around 65% of liquid milk sales are accounted for by the main food retailers (KPMG, 2002). With respect to meat products, the data are more indirect in that they relate to consumption of meat via the retail sector as a whole rather than the supermarkets alone, though given their share of consumer markets, the figures are informative of the likely dominance in the procurement market. With this caveat in mind, the data show that 85% of beef is consumed via the retail sector, with the corresponding figures for pork and lamb being 81% and 90% respectively.

Establishing detailed empirical evidence of the existence of buyer power, however, is problematic. Indicative measures often rely on anecdotal accounts, small-scale surveys of the parties involved or at a more representative level, summary measures of concentration. Relating simple measures of concentration to the existence of selling power has long been recognised as of limited value and the same is true for buying power (Clarke *et al*, 2002). For example, the high levels of concentration evident in the UK food retailing sector, coupled with the high profits they report, is not necessarily indicative of the exploitation of market power. Similarly, there is a spectrum of econometric approaches that may be employed to detect market power. Where estimation is based upon price data alone, such as in orthodox price transmission studies (e.g. London Economics, 2004) the veracity of antitrust inference is undermined by the reduced form nature of the price regressions employed (Hoehn *et al.* 1999, p.113). Although structural econometric models address this issue of 'measurement without theory' directly, they are often confounded by data limitations and methodological shortcomings relating to market definition and the validity of the behavioural assumptions employed (Baker and Bresnahan, 1992). In these circumstances, a simple and reliable test derived from economic theory detecting the existence of market power offers some appeal, and it is in this regard that this paper seeks to make a contribution. Specifically, we provide such a test by devising a simple quasi-reduced form model of price formation at retailer and supplier levels in which the hypothesis of buyer power can be readily tested using widely available market-level data. While the approach does not aim to derive an explicit measure of market power, it does provide a test for its existence, emphasising the test's 'path-finder' role alongside extant sources of indicative evidence.

In terms of the academic literature, the test proposed here lies between two related fields in the industrial organisation literature. At one end is the estimation of structural models in the context of the new empirical industrial organisation literature. Bresnahan (1989) provides an overview. The key feature of this methodology is the use of exogenous shocks (such as exogenous shifts in the demand or supply functions) in order to identify the presence of market power. From this one can retrieve a measure of the aggregate conjectures representing the degree of market power in a specific market. In the approach followed here, we also employ exogenous shocks as a means to detect the potential for market power. At the other end is the empirical literature on the incidence of policy changes (such as tax changes) or other shocks since the incidence of taxes may differ in the presence of market power. Fuerstein (2002) and Delipalla and O'Donnell (2001) would be recent examples. The approach followed here relates to these empirical strategies in that we exploit the presence of exogenous shocks in order to identify the presence of market power based on a theoretical model of the incidence of shocks on both upstream and downstream prices. As we explain below, the detection of market power simply depends on how these shocks affect *both* sets of prices. While the simplicity of the approach does not allow us to retrieve an empirical estimate of the degree of market power the trade-off does circumvent some of the obstacles inherent in the estimation of structural econometric modelling and the difficulties associated with the interpretation of estimated conjectures.

More specifically, in the framework we present, the difference (or spread) between prices at different marketing levels can be attributed solely to marketing costs under competitive conditions. In other words, shocks impact on prices at each marketing level equally. If market power exists then the spread between retail and producer supply prices behaves differently since price setting by the sector with market power will be reflected in the mark down that the firms can earn, and so affects the spread. Hence, as we show in section 2, where buyer power exists, market shocks have a differential impact at each stage in the marketing chain and thus determine the behaviour of the spread between prices at different stages in addition to marketing costs. In effect, shocks to the underlying supply and demand functions are mediated through market power parameters and thus give rise to predictable effects on the spread. In the absence of market power, the effect of shocks is common at all market levels so that the spread is simply determined by marketing costs.

In what follows, we develop a model of price transmission in a two-stage vertical market that explicitly allows for shocks in both the demand and supply functions for the product. Moreover, given that the impact of shocks appears with definite sign in the theoretical model of the spread, the basis for reliable inference is strengthened accordingly. Our approach is applied to data from six food groups in the UK food industry. For each product, the empirical test rejects the null of perfect competition at conventional levels of significance. Furthermore, coefficients on the exogenous shifters are signed according to the predictions in the theoretical model in all cases. The paper is structured as follows. In Section 1 we outline the theoretical model that underpins our conceptualisation of a vertically related market. The econometric techniques employed are discussed in Section 2 while Section 3 describes the data. The results of the testing procedure are outlined in Section 4 and we offer some concluding thoughts and caveats in Section 5.

1. Theoretical Model

In this section, we outline a simple framework that delivers a formal test of market versus perfect competition that we use to motivate the empirical analysis. The demand function for the processed product is given by:

$$Q = h(R, D) \qquad (1)$$

where R is the retail price of the good under consideration and D is a general demand shifter. The supply function of the agricultural raw material is given by (in inverse form):

$$P = k(A, S) \qquad (2)$$

where A is the quantity of the agricultural raw material and S is the exogenous shifter in the farm supply equation.

In accordance with the findings of the Competition Commission (op. cit.) the source of market power in the food chain is given to be at the retail level. For a representative retail firm, the profit function is given by:

$$\pi_i = R(Q)Q_i - P(A)A_i - C_i(Q_i) \qquad (3)$$

where C_i is other costs and, assuming a fixed proportions technology, $Q_i = A_i / a$ where a is the input-output coefficient which is assumed to equal 1. This assumption corresponds closely to the construction of the data in the vertical market chain used in the empirical analysis that follows. Constant returns to scale are assumed. The first-order condition for profit maximisation is given by:

$$R + Q_i \frac{\partial R}{\partial Q}\frac{\partial Q}{\partial Q_i} = \frac{\partial C_i}{\partial Q_i} + aP + aA_i \frac{\partial P}{\partial A}\frac{\partial A}{\partial A_i} \qquad (4)$$

In order to get an explicit solution, consider linear functional forms for equations (1) and (2) and assume *a*=1 (which is consistent with the construction of the data series): *a=1*

$$Q = h - bR + cD \tag{1'}$$

$$P = k + gA \tag{2'}$$

with domestic supply being given by:

$$A = Q + S$$

where S is the exogenous supply shifter. From this we can rewrite (4) as:

$$R = M + P + \mu g Q \tag{4'}$$

where μ is the aggregate input conjectural elasticity , such that with n firms in the retail sector, $\mu = (\Sigma_i\,[\partial A/\partial A_i][A_i/A])/n$. This parameter can be interpreted as an index of buyer market power with μ=0 representing competitive behaviour and 1=μ representing collusive behaviour. While μ is the measure of buyer power, as noted above, we do not aim to derive an explicit value for this parameter, but test only for its existence. M is a composite variable that represents all other costs that affect the retail-farm price margin.

To allow for changes in costs, we assume a linear marketing cost function of the form:

$$M = y + zE \tag{5}$$

where y is a constant and represents the costs of inputs from the marketing sector (for example, wages). Using (1'), (2'), (4') and (5), we can derive an explicit solution for the endogenous variables:

$$Q = \frac{(h - by) + cD - bzE - bgS)}{1 + bg(1 + \mu)} \tag{6}$$

$$R = \frac{h + [1 + bg(1 + \mu)][(1 - b)(y + k + gS) + (1 - bz)E + cD)]}{1 + bg(1 + \mu)} \tag{7}$$

$$P = \frac{g[h - by + cD - bzE] - g[b - (1 + bg(1 + \mu))(k + S)]}{1 + bg(1 + \mu)} \tag{8}$$

To derive the spread between retail and producer prices, use (7) and (8) to give:

$$R - P = \frac{hg\mu + (1+bg)(y+zE) + g\mu cD - bg\mu(k+gS)}{1+bg(1+\mu)} \quad (9)$$

Note that if oligopsony power does not matter in determining the retail-producer price spread (i.e. μ=0), then equation (9) reduces to:

$$R - P = y + zE = M \quad (10)$$

i.e. the source of the retail-producer price margin in a perfectly competitive industry is due to changes in marketing costs only. In this case, the exogenous shifters relating to the retail and agricultural supply functions play no role in determining the spread. This is not to say that they do not affect each price individually, but in a perfectly competitive industry they play no role in determining the relative gap between the prices at each stage of the food chain. Correspondingly, if oligopsony power in the food sector is important, each shifter affects the two prices differentially and thus the margin between the prices changes.

Equations (7)-(9) form the basis of our econometric modelling. Consider first of all equation (9) that relates to the retail-producer spread. Note that if buyer market power does characterise the UK food sector, then the supply and demand shifters should enter our econometric model of the margin between retail and producer prices. Writing the margin equation in unrestricted form (i.e. in terms of prices) gives an empirical testing equation,

$$R = \beta_0 + \beta_1 P + \beta_2 M + \beta_3 X + \beta_4 N \quad (11)$$

From (9) and (10), $\beta_1>0$, and $\beta_2>0$ irrespective of the degree of retail competition. The test for the existence of buyer power is whether the coefficients on these variables in the retail-producer spread equation are statistically significant. Specifically, rejection of the null hypothesis,

$$H_0 : \beta_3 = \beta_4 = 0$$

implies the existence of buyer market power. Furthermore, equation (9) unambiguously signs the effect of the shifters in the presence of market power. Whereas shocks to the demand shifter widen the margin, supply-side shocks narrow it, hence if market the shifters are significant in the margin equation, theory predicts that $\beta_3>0$ and $\beta_4<0$ in (11). In the empirical section, we test these propositions using data for six product groups.

2. Empirical Method

To allow for the possibility that retail and producer prices of each product group are non-stationary and cointegrated, we couch the empirical analysis in a vector autoregressive (VAR) framework. For each of the six product groups it is assumed that the data may be approximated by a VAR(*p*) model,

$$x_t = \Phi_1 x_{t-1} + \Phi_2 x_{t-2} + ... + \Phi_p x_{t-p} + \Psi D_t + \varepsilon_t \qquad (12)$$

where x_t is a (k*1) vector of jointly determined I(1) variables, D_t is is a (d*1) vector of constants and centered seasonals and each $\Phi_i (i = 1,..., p)$ and Ψ are (k*k) and (k*d) matrices of coefficients to be estimated using a (t = 1, . . .T) sample of data. ε_t is is a (k*1) vector of n.i.d. disturbances with zero mean and non-diagonal covariance matrix, Σ

Equation (12) represents an unrestricted reduced form representation of the variables in x_t comprising retail and producer prices, a measure of marketing costs and the supply

and demand shifters. Given the monthly frequency of the data, lag length (p) of the VAR is determined for each product group in step-wise fashion (p=13,12,…,1) using standard information criteria and vector-based diagnostics. The preferred lag length is thus the most parsimonious model that is free of residual correlation at the 5% significance level.

The presence of cointegration is detected by estimating (12) in its error correction representation using Johansens's (1988) maximum likelihood procedure,

$$\Delta x_t \alpha\beta' x_{t-p} + \sum_{i=1}^{p-1}\Gamma_i \Delta x_{t-i} + \Psi D_t + \varepsilon_t \qquad (13)$$

Attention focuses on the (*k*r*) matrix of co-integrating vectors, comprising β, that quantify the 'long-run' (or equilibrium) relationships between the variables in the system and the (*k*r*) matrix of error correction coefficients, α, the elements of which load deviations from equilibrium (*i.e.* $\beta' x_{t-k}$) into $\Delta \mathbf{x}_t$, for correction. The Γi coefficients in (13) estimate the short-run effect of shocks on $\Delta \mathbf{x}_t$, and thereby allow the short and long-run responses to differ. The number of cointegrating relations, corresponding to the rank of β in (12), is evaluated by Johansen's Trace (η_r) and Maximal Eigenvalue (ξ_r) test statistics (Johansen, 1988). The η_r statistic tests the null that there are at least *r* cointegrating relationships ($0 \leq r < n$) and the ξ_r evaluates the null that there are *r* against the alternative that there are at most $r+1$ such relationships. While the η_r test is generally preferable because it is robust to residual non-normality and delivers a sequentially consistent test procedure, it is standard practice to report both test statistics. In the empirical analysis that follows we also report both asymptotic and the degree-of-freedom-adjusted test statistics of Cheung and Lai (1993).

Where a single cointegrating relationship is detected, formal testing is undertaken to investigate whether buyer power is implied. Following from section 2, if the vertical market for a product is perfectly competitive, retail and producer prices may be expected to form a cointegrated relationship with at most marketing costs. Where retailers exert buying power, the shifters also enter the pricing relationship. This then gives rise to a null hypothesis of perfect competition which can be evaluated empirically by a standard likelihood ratio test of the exclusion restrictions on the shifters in the cointegrating relation. In addition, given that the theoretical model signs the parameters in the pricing relation we can offer some additional evidence on market power by comparing the estimated signs of the shifters in the cointegrating relation with that predicted by the theoretical model.

3. Data From The Food Industry

In this paper we focus on six products to explore the presence of market power. We use fresh products as these are subject to the smallest degrees of processing by the post-farm gate chain prior to the food reaching the retail shelves, and thus potentially provide a clearer correspondence between theory and data. Further, given this limited processing, it is more readily acceptable to envisage production along the supply chain as being characterised by fixed proportions technology, as in the theoretical model. Finally, as the introduction highlighted, the fresh food sector is more likely to reveal areas of asymmetry in bargaining since this is where small suppliers predominate and thus where evidence of market power is most likely to be found.

We analyse the nominal monthly prices of six UK food products, namely: apples (A); beef (B); chicken (C); lamb (L); milk (M) and potatoes (Pt) at retail (R) and producer (P) levels. In addition, each price model includes three industry-level 'shifters' representing proxies for marketing costs and shocks to the demand and supply functions. As discussed above, retail and producer product prices are expressed in prices per standard unit (pence/kg of carcass weight for all meats; pence/pint for liquid milk, pence/lb for potatoes, and apples are an index [1987=100] of prices in pence/lb). The price series are illustrated in Figure 1.

[Figure 1 about here]

As Figure 1 illustrates, there is considerable variation in the price series between products and across marketing levels, although a tendency to diverge over time is a common feature While growth in the price spread is not in itself indicative of market power (marketing costs may account for it), it is necessary given the strong trend-like behaviour of the shifters, which are plotted in Figure 2.

[Figure 2 about here]

Referring to Figure 2, it is evident that all shifters display the tendency to grow over time. As noted in section 1, measures of product-specific marketing costs are not available in the UK and thus we use an index of unit wage cost index for manufacturing industries (M), given the labour intensity of the food retailing sector. In order to incorporate the impact of farm-level production costs, the supply shifter (S) represents a price index of all goods and services purchased on UK farms. Demand-side shocks are proxied by two measures. For meat products, we have a direct measure reflecting health scares with respect to the consumption of meat products. Specifically, we use the (natural logarithm of the) cumulative count of articles regarding the health and safety of food published in four broadsheet newspapers (D1), on the basis that such articles principally relate to or affect the demand for meat, rather than non-meat products. For non-meat products, no obvious direct demand shifter was available. For this case, we therefore use the food retail price index (D2) on the basis that this represents a general demand shifter affecting the food retailing sector as a whole. The application of the Augmented Dickey-Fuller test indicates that all prices and shifters are integrated of order one in levels and stationary in first differences. ADF test statistics are reported in the Appendix.

4. Results

Having established the non-stationarity of the data, equation (13) is estimated for each of the six product groups sequentially for k = 13 to 1. Since there is no consensus on the best criterion to use to determine lag length, three commonly applied measures are used here, namely the information criteria developed by Shartwz, Hannan-Quinn and Akaike (SBC, HQC and AIC respectively) and vector diagnostic tests for residual autocorrelation, heteroscedasticity and normality. The SBC tends to select the most parsimonious model and the AIC the least with the HQC selecting a lag length that is generally common to one of the other two, in roughly equal measure. In only one case (milk) is the lag length selected unanimously by the three information criteria. The vector test for residual autocorrelation tends to select models with longer lag lengths and hence concur with the AIC in most cases. To determine the preferred lag length, a consensus view is taken, although this usually conforms to the most parsimonious model in which the null of no residual correlation cannot be rejected at the 5% level. In many cases, test statistics reject the null of (residual) normality emphasizing that care should be exercised in interpreting results. The selected models are unrestricted reduced forms and represent the baseline models against which all subsequent parameter restrictions are evaluated.

Having established lag length, the cointegrating rank is evaluated in the selected specification for each product group. Table 1 reports the results from the cointegration analysis using the Trace (η_r) and maximal Eigenvalue (ξ_r) tests in asymptotic (∞) and finite sample (*T-mp*) forms (Cheung and Lai, 1993). Overall, the evidence points to the presence of a single cointegrating vector in all product groups. Evaluating hypotheses at the 5% significance level, the null of no cointegration is rejected in 11 out of 12 tests using asymptotic critical values and on 9 out of 12 occasions using degree-of-freedom-adjusted critical values. Confining inference to the more stringent (degree-of-freedom-adjusted) tests, every product has at least one statistic rejecting the null of no cointegration at the 5% level. In the three cases where the null cannot be rejected at 5%, only one is above 10%. No finite sample statistics reject the null of multiple cointegrating vectors at the 5% level of significance.

Table 1: Asymptotic (∞) and Finite Sample Test Statistics for Cointegration

Product	Rank	Trace	Maximal Eigenvalue	Trace	Maxi,al Eigenvalue
Apples	0	83.77[0.002]**	36.88[0.018]*	77.87[0.009]**	34.28[0.041]**
	1	46.89[0.060]	23.28[0.165]	43.59[0.118]	21.64[0.247]
	2	23.62[0.224]	16.36[0.213]	21.95[0.311]	15.20[0.286]
	3	7.26[[0.554]	4.66[0.782]	6.75[0.613]	4.33[0.819]
	4	2.60[0.107]	2.60[0.107]	2.42[0.120]	2.42[0.120]
Beef	0	78.75[0.007]**	40.89[0.004]**	71.18[0.037]*	36.96[0.017]*
	1	37.86[0.312]	23.90[0.140]	34.22[0.495]	21.60[0.250]
	2	13.96[0.843]	7.29[0.932]	12.62[0.905]	6.59[0.959]
	3	6.67[0.622]	5.11[0.729]	6.03[0.695]	4.62[0.787]
	4	1.56[0.211]	1.56[0.211]	1.41[0.235]	1.41[0.235]
Chicken	0	85.85[0.001]**	35.92[0.024]*	76.84[0.011]*	32.15[0.077]
	1	49.93[0.030]*	26.21[0.072]	44.69[0.095]	23.46[0.158]
	2	23.72[0.219]	14.84[0.313]	21.24[0.353]	13.28[0.441]
	3	8.89[0.383]	6.24[0.590]	7.96[0.477]	5.59[0.671]
	4	2.65[0.104]	2.65[0.104]	2.37[0.124]	2.37[0.124]
Lamb	0	82.11[0.003]**	34.23[0.042]*	75.15[0.016]*	31.32[0.097]
	1	47.88[0.048]*	25.79[0.082]	43.83[0.113]	23.61[0.152]
	2	22.09[0.303]	15.68[0.254]	20.22[0.419]	14.35[0.351]
	3	6.41[0.651]	5.25[0.712]	5.87[0.713]	4.81[0.765]
	4	1.16[0.281]	1.16[0.281]	1.06[0.302]	1.06[0.302]
Milk	0	103.04[0.000]**	61.83[0.000]**	96.83[0.000]**	58.11[0.000]**
	1	41.20[0.183]	20.87[0.294]	38.72[0.275]	19.61[0.381]
	2	20.33[0.411]	11.38[0.619]	19.11[0.496]	10.70[0.684]
	3	8.95[0.377]	8.53[0.335]	8.41[0.430]	8.02[0.385]
	4	0.42[0.517]	0.42[0.517]	0.39[0.530]	0.39[0.530]
Potatoes	0	67.89[0.069]	39.08[0.008]*	60.67[0.216]	34.92[0.033]
	1	28.81[0.777]	13.35[0.857]	25.75[0.894]	11.93[0.925]
	2	15.47[0.754]	10.93[0.662]	13.82[0.850]	9.77[0.767]
	3	4.53[0.852]	3.43[0.904]	4.05[0.893]	3.07[0932]
	4	1.10[0.295]	1.10[0.295]	0.98[0.322]	0.98[0.322]

**denotes significance at 1%; * denotes significance at 5%, and p=values are in parenthesis. Asymptotic (∞) are those of those of the Osterwald Lenum (1992) and finite sample (degree of freedom) adjusted test statistics are those of Cheung and Lai (1993) where the correction is (*T-mp*) where *T* is sample size and m is number of endogenous variables and *p* is the lag length in the VAR.

On the basis of these results we proceed on the assumption that a single cointegrating vector is present for each product group. Normalising each vector on retail prices, the long-run coefficients and associated standard errors are reported in Table 2. Recall that the theoretical model presented in section 2 signs these coefficients such that, $\beta_1>0$ and $\beta_2>0$; and where market power exists, $\beta_3>0$ and $\beta_4<0$.

Referring to the table a number of points seem noteworthy. First, price transmission coefficients (β_1) are positive in all cases and significantly so in all but potatoes. Second, marketing costs, as proxied by labour costs in manufacturing, (β_2) are positive in three cases, significantly so in two. Third, the demand shifter coefficient (β_3) is significantly positive in the cointegrating relations of four out of six products; and fourth, the coefficient on the supply shifter is significantly negative for all six products.

Table 2: The Cointegrating Vectors (normalised on retail prices)

Product	Producer prices (β_1)	Marketing costs (β_2)	Demand shifter (β_3)	Supply shifter (β_4)
Apples	1.94** (0.23)	-6.42** (2.2)	8.07* * (2.21)	-3.73** (1.33)
Beef	2.02* *(0.23)	6.15* *(1.44)	18.5* (7.39)	-3.19** (0.88)
Chicken	10.38** (1.55)	12.31** (3.04)	30.3 (16.24)	-11.79** (1.93)
Lamb	3.95** (0.62)	-7.19 (6.55)	148.03* *(42.12)	-29.01** (5.73)
Milk	0.55* *(0.13)	0.06 (0.08)	0.10 (0.10)	-0.13* (0.05)
Potatoes	0.49 (0.32)	-2.02* *(0.54)	3.24* *(0.32)	-1.67* *(0.53)

Figures in bracket are asymptotic standard errors; ** denotes significance at the 1% and *denotes significance at the 5% level.

Of key interest are the last two results which indicate that the shifters play an important role in the long run determination of prices, and enter the cointegrating relations with signs that are consistent with the use of retail market power. To investigate this issue more formally, we perform a second set of likelihood ratio tests to evaluate these exclusion restrictions, results from which are contained in Table 3. The first two columns test the individual significance of each shifter in each cointegrating vector and thus perform the same role as the standard errors in Table 2. Results are similar, with the statistical significance of the shifters being confirmed in 10 out of 12 occasions at the 5% level. The final column of Table 3 evaluates the null hypothesis that both shifters are jointly zero. This corresponds to perfect competition in the theoretical model and is rejected in all six products studied. Overall, the behaviour of prices in the products considered here are consistent with the use of buyer power.

Table 3: Tests for Market Power

Product	$H_0=\beta_3=0$	$H_0=\beta_4=0$	$H_0=\beta_3=\beta_4$
Apple	6.38 [0.01]*	3.86 [0.05]*	7.04 [0.03]*
Beef	4.06 [0.04]*	10.76 [0.00]**	11.12 [0.00]**
Chicken	4.12 [0.04]*	0.47 [0.49]	26.48 [0.00]**
Lamb	4.69 [0.03]*	8.34 [0.00]**	15.50 [0.00]**
Milk	0.66 [0.42]	5.83 [0.02]*	7.71 [0.02]*
Potatoes	18.14 [0.00]**	16.1 [0.00]**	18.30 [0.00]**

Figures in bracket are asymptotic *p*-values; ** denotes significance at the1% and *denotes significance at the 5% level

Returning to the results presented in Table 2, there are two caveats to note. First, while the theoretical model additionally implies that $\beta_1=1$, this condition is not borne out in the empirical setting. This may be due to heterogeneity within product groups and other practical factors such as wastage and product specification that interfere with the strict one-to-one correspondence of the theoretical model. Second, with respect to marketing costs, the significantly negative marketing cost coefficients in the models for apples and potatoes are at odds with the theoretical model outlined above. This is likely to reflect the inadequacy of a general marketing cost variable in these cases and/or that we are not picking-up specific trends in marketing technology or costs in these two sectors. Notwithstanding these two caveats, the overall results give considerable support for the exercise of oligopsony power in the food sector.

5. Concluding comments

In this paper, we have devised a simple yet robust means of testing for the presence of buyer power in vertically-related markets such as those characterising the food chain. By constructing a quasi-reduced form model of the retailer-supplier pricing equations, the null of perfect competition can be rejected if the shifters from the supply and demand equations are significant and correctly signed. In principle, the approach sits between other methods of evaluation, to which it is complementary. In particular, we are able to move away from naïve concentration-based indicators of market power and the practical limitations of structural econometric modelling. The approach is simple and transparent yet delivers a statistical test derived from a theoretically consistent basis. Furthermore, the test demands relatively little in terms of data and is executed using standard techniques of modern time-series analysis.

The technique is most applicable where products undergo relatively little transformation between marketing levels and is thus particularly well-suited to the relatively unprocessed products of the food chain. In the UK at least, these are also products over which concerns of buyer power abuse have been most acute. Drawing on data from a basket of six basic products of the UK food industry, we show that in all cases, the hypothesis of perfect competition can be firmly rejected at conventional levels of significance, implying that for these food products at least, the market is characterised by buyer power. As such, our findings corroborate the findings of Competition

Commission (2000) and lend support to the recent request by the Office of Trading for further detailed scrutiny of the UK food chain by the UK's competition authorities.

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Appendix Table 1: ADF Test Statistics

Variable	Levels		First-differences	
	ADF	Lag	ADF	Lag
RA	-2.67	0	-10.88**	0
PA	-2.38	4	-6.94**	9
RB	-1.88	0	-12.70**	0
PB	-2.49	1	-8.41**	0
RBr	-2.74	0	-7.77**	1
PBr	-2.91	1	-9.49**	1
RC	-1.52	3	-11.10**	2
PC	-2.38	4	-4.13**	3
RL	-1.83	6	-7.22**	6
PL	-1.50	6	-8.27**	5
RP	-1.67	0	-11.20**	0
PP	-2.24	8	-6.68**	5
RM	-1.14	3	-7.78**	2
PM	-2.11	13	-7.29**	1
RPt	-2.12	0	-11.30**	0
PPt	-2.61	2	-8.16**	1
RE	-2.75	1	-9.46**	1
PE	-2.86	5	-2.97*	4
S	-2.61	12	-3.06*	10
D1	-1.93	3	-3.18*	3
D2	-2.37	0	-12.02**	0
M	-1.26	9	-3.92**	7

Lag length is selected on basis of the information criteria (see main text for details). Regressions include constant, trend and seasonals (if appropriate) in the levels; constant (and seasonals) only in first differences. 95% (*) and 99% (**) critical values are -3.45 and -2.88 respectively.