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STRUCTURAL CHANGES IN KOREAN LIVESTOCK PRICES

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Key words: structural changes, var, multivariate gradual switching regressions, transition function, grid search, livestock prices

ABSTRACT

A vector autoregressive (VAR) model with grid search that allows parameters to vary with time was applied to investigate the structural break points and the speed of adjustment for Korean livestock prices. Our results suggest that VAR for Korean beef prices experienced a rather rapid structural shift in the mid 1987 while pork prices showed gradual structural break in the early 1991. Chicken prices experienced a structural change centered on late 1984. Beef, pork, and chicken prices experienced a structural change in the early 1980s.

I. Introduction

In econometric jargon, a structural change or a structural break means that the relationship between the dependent and independent variables changes (Ramanathan 1995) or there is a change in parameter constancy from one regime to another (Johnston and Dinardo 1997). It is a common problem in empirical econometric modeling that the parameter estimates are non-constant (Kennedy 1998; Heinesen 1997). If the parameters are not constant for allsample observations, the statistical model is changed and the

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estimates may lose their significance.

A vector autoregression (VAR) model is widely used for forecasting economic variables and for furnishing insights into dynamic relationships including changes in interrelationships among price series.[1] The assumption of structural stability in the unknown parameters of the underlying economic model is implicit in VAR models (Sims 1980). In the event of a structural change in the underlying economic relationships, standard VAR models may produce biased forecasts and inaccurate inferences regarding dynamic relationships among the economic variables (Goodwin 1992).

The objectives of this paper are twofold. First, exact structural break points are investigated for pairs of VAR models for Korean livestock prices. Second, the nature of structural change is also analyzed. Livestock prices examined in this study are retail, wholesale, and farm prices of beef, pork, and chicken. Monthly Data from 1977.1 to 1999.12 are used.

II. Elements of Structural Changes in Livestock Prices

Recent many empirical researches in the livestock industry notably undertaken in the United States suggest several reasons to suspect a structural change in economic relationships determining livestock prices. As Boehlje (1999) observed, for example, the U.S. food production and distribution industry is in the midst of structural change - change in production characteristics, in worldwide production and consumption, in technology, in size of operation, and in geographic location. On the demand side, consumption of

[1] VAR models differ from standard econometric analysis of structural relationships in that they do not apply the usual exclusion restrictions to specify a priori which variables appears in which equations (Goodwin 1992). Instead, a set of distributed lag equations is used to model each variable as a function of other variables in the system. Such an approach could reduce a priori restrictions on the dynamic relationships (Sims 1980).

poultry products has continuously increased, while beef consumption has steadily declined. Many have found that structural change in demand for food in the United States has occurred since the mid 1970s, partly due to the increasing health concerns of consumers regarding red meat consumption (e.g., Eales and Unnevehr 1988; Moschini and Meike 1989; Choi and Kim 1990).[2]

In the case of the United States, significant changes in the structure of the cattle industry have been examined. Paul (1987) has noted that the declining importance of terminal markets relative to direct markets has had significant effects on pricing relationships in the U.S. cattle industry. The expansion of electronic marketing systems in the 1980s (Bailey et al. 1991) and increased use of cattle future markets in the 1970s and 1980s (Paul 1987) may also have altered cattle pricing relationships. Finally, considerable changes have occurred in the structure of the livestock slaughter industry representing increased concentration of the meatpacking industry through the 1970s and the 1980s. The four-firm concentration ratio (CR 4) that is often used as a measure of the degree of consolidation shows that four firms accounted for 78 percent of purchases in 1997, up from 36 percent in 1980. In contrast, the average CR 4 is 40 percent across all U.S. manufacturing industries (Mathews, Jr., et al., 1999). Hog slaughter is less concentrated-the top four hog packers handled 54 percent of slaughter in 1997. Ward (1987) has noted that changes in marketing patterns and industry concentration ratios may have had important structural repercussions on the market performance of the beef market. The effect of increased market concentration on the speed of price adjustment has received considerable attention in recent years (Goodwin 1992). A number of papers including Weaver, Chattin, and Banerjee (1989), and Brorsen, Chavas and Grant (1991), and

[2] Further, in Reeds recent work (2000), evidence is provided that consumers' responses to changes in relative retail prices affect retail and farm prices of related products is provided. All these changes in the structure of production and demand may affect pricing behavior of the industry under consideration and interrelationships among prices.

Martin (1993) dealt with the relationship between the degree of market concentration and the speed of price adjustment. Some confirmed that increased concentration of an industry causes faster price adjustment, while the others have found a negative relationship between market concentration and the speed of price adjustment. In light of the conflicting conclusions offered by previous research, the effect of increased concentration of the meat market on cattle price dynamics is uncertain (Goodwin 1992).

So far, the major aspects of structural change in the U.S. livestock industry have been examined. In the case of the Korean livestock industry, an increase in national income and trade liberalization have had a significant influence on Korean agriculture leading to structural adjustment (Kim, Y.T 1997). The livestock sector was not exempted. In particular, during the past few decades, especially since the mid-1980s, changes in consumers' diet patterns from grain-dominated to one consuming more fruit, vegetable and meat have brought about not only a significant increase in meat consumption but also changes in the structure of meat demand (KREI 1999). For instance, per capita consumption of cereals decreased from 216 kg in 1970 to 156 kg in 1999, while per capita consumption of livestock products increased from 8.4 kg in 1970 to 30.5 kg in 1999. Of livestock products, per capita beef consumption has increased the most significantly. During the last two decades, it has increased more than three times. These changes in meat consumption raise strong suspicions that there have been changes in Korean livestock price relations.

The next important factor that may have contributed to changes in Korean livestock prices is the effect of import liberalization. The conclusion of the Uruguay Round negotiations brought about the general reform of the agricultural sector and accelerated agricultural market opening including livestock products (KREI 1999). Since most Korean agricultural prices are higher than those of imported products, an increase in imported agricultural products may lower domestic agricultural prices and may lead to a change in the structure of meat demand and

supply, and in turn, of livestock prices.

Unlike the case of the U.S. livestock industry, the possible effects of increased concentration and vertical integration on Korean livestock prices seem relatively small, especially in the beef sector, although it will become increasingly important in the near future. The poultry industry in Korea is a leading sector in which vertical integration is rapidly proceeding. It was not until the mid 1980s that widespread concerns about the advantage of integrated management for the poultry industry arose, and that vertically integrated firms emerged. Most of vertical integration is a form of partial integration, rather than full-range integration covering both production and marketing activities. As of 1998, poultry production by vertically integrated firms accounted for 45 percent of total poultry production, an increase from 15 percent in 1989 (Jung 1999). In the pig industry, vertical integration led mainly by large scale firms has been actively progressing. The pork market share of vertically integrated firms increased from 1.3 percent in 1991 percent to 26.1 percent in 1998 (Jung 1999). The Korean beef industry is still operated by a large number of small-scale individual producers and is the lagging sector in terms of the degree of vertical integration (KREI 1999). The average number of cattle head per household in 2000 is lower than 6, proving that cattle production has not fully developed into an efficient business (KREI 1999).

It needs to be also mentioned that most Korean livestock products are still marketed by a large number of scattered small-scale marketing agents. As of 1997, the total number of slaughtering firms was 109, down from 179 in 1980. Most slaughtering firms are small scale as per firm average slaughtering capacity per day was less than 100 head in 1997 for both beef and pork slaughter (Jun, et al. 1999). The number of meat retailers in 1999 was 48,176. About 83 percent of them had less than 10 pyeong(approximately 0.4 square yards) of shop area (Jun, et al. 1999). This may imply that in Korea, consolidation or market power in meat marketing, slaughtering, and retailing is not a major factor affecting the structure of meat price relationships.

III. Models for Investigating Structural Change

In general, four types of tests have been particularly popular for testing structural change: the Chow test, the predictive failure test, tests based on recursive residuals (CUSUM), and tests based on recursive parameter estimates (Kennedy 1998). Among them, the Chow test and CUSUM tests for structural change have been widely used but are subject to several weaknesses (Swamy, et al. 1980). For example, those test techniques assume an instantaneous abrupt structural break, which may be inappropriate in the real world. Changes in economic relationships may be gradual. Suppose there is a policy change at time t. In actual practice, however, the switch to the new regime need not take place at time t, nor need the switch be sudden. Many factors can account for gradual structural changes including adjustment costs. By the same token, it is more likely that changes in interrelationships among livestock prices occurs gradually rather than instantaneously at a discrete point in time. Furthermore, the Chow test requires a prior specification of the timing of the change, which is unknown a priori. For these reasons, we apply the Gradual Switching VAR model that identifies the timing of the change while allowing the speed of adjustment between alternative regimes to be gradual as well as instantaneous.

Structural changes in time series variables can be considered in several ways, such as the stationarity of individual variables, the cointegration of two or more variables, and the VAR or Vector Error Correction: VEC). In this study, the focus will be on the structural change in VAR.[3] The basic motivation of

[3] In order to apply VAR, all time series data should be stationary (Enders 1995). The Augmented Dickey Fuller (ADF) test results suggest that all eight time series used in this study have unit root implying non-stationarity. Thus, it should be recognized that the results of this study have limitations to a certain degree. The basic motivation of this study was to identify approximate timing of structural change, not the exact ones.

searching for structural change in this study is to estimate the approximate starting point and speed of adjustment for each pair of prices, which could give shed light on the nature of structural changes in livestock prices. The starting point and speed of adjustment estimated will provide a guideline to determine the regime.

A VAR system for time-ordered variables can be written as:

(1) $Y_t = \Psi Y_{t-s} + \varepsilon_t$

where t refers to time (t=1, ... ,T), Yt is an $mT \times 1$ vector of economic variables, Ψ is an $mT \times mT$ matrix of parameters, and ε_t is an $mT \times 1$ matrix of white noise innovations. An approach for allowing parameter drift in a VAR model can be found in the gradual switching method introduced by Bacon and Watts (1971) and has been applied by Tsurumi Wago, and Ilmakunna(1986), Moschini and Meilke (1989), and Goodwin (1992), Goodwin and Brewster (1995), and Goodwin, Harper and Schnepf (2000).

The gradual switching method allows structural change to occur gradually. A structural change can be interpreted as a shift in the parameter matrices from $\Psi(1)$ in the first regime to $\Psi(2)$ in the second regime. In this application, the change was allowed to start at an unknown joint μ^* and to occur at an unknown gradual rate of σ^*. A transition function λ_t, that is constrained by construction to lie in the open interval (0, 1), is used to represent shifting between regimes. Our specification of the mixing problem allows us to rewrite the VAR as:

(2) $Y_t = (1-\lambda_t)\Psi^{(1)} Y_{t-s} + \lambda_t \Psi^{(2)} Y_{t-s} + \varepsilon_t$

Following Tsurumi, Wago and Ilmakunnas (1986), the transition function can be any function that satisfies

(3) $\lim_{s_t \to \infty} \text{trn} \ (s_t/\sigma) = 1,$

$\text{trn}(0) = 0,$

$\lim_{\sigma \to \infty} \text{trn} \ (\mu/\sigma) = 1,$

and s_t is given by

$$s_t = 0 \text{ for } t \leq \mu_t^*$$

$$= t - \mu_t^* \text{ for } t > \mu_t^*$$

In this study, following Goldfeld and Quandts method (1972) that is applied by Goodwin et al. (2000), the transition function λ_t is given by:

(4) $\lambda_t = \Phi((t-\mu)/\sigma) \quad t = 1, \ldots, N;$

where Φ is the normal cumulative density function (Φ(c)=Prob($x \leq c$) with $x \sim$N(0,1)) and μ and σ are parameters to be estimated.[4] μ gives the mean point of the switch that is the observation lying one-half way between regime I and II (i.e., λ_t =0.50). The bandwidth parameter represents the speed of adjustment between the two regimes with the larger values of σ corresponding to more gradual adjustments between two regimes. Following Tsurumi (1986) and Goodwin et al. (1995), it is assumed that each equation in the VAR has the same value of the transition function.[5]

[4] In Bacon and Watts (1971), Tsurumi and Wago (1986), and Goodwin et al. (1992), the hyperbolic tangent function was used as the transition function. Results contained in Tsurumi et al. (1986), indicate that empirical results obtained from the application of transition functions are not, in general, sensitive to the choice of functional form for the transition function. In Ohtani et al. (1990), it is assumed that transition path is a polynomial of time, the degree of the polynomial being decided by a model selection criterion such as AIC.

[5] This does not imply that the extent of structural change must be the same for each equation. Rather, this only requires the timing of the change and speed of adjustment to the new equilibrium to be the same for each equation. It is possible that individual equations will not realize significant changes, in which case the transition parameters will be zero. Although individual joint points and speed of adjustment parameters for each equation are conceivable, the tractability of such an approach is limited by computational considerations (Goodwin and Brester 1995).

A test of the statistical significance of the transition function parameters (μ and σ) estimated is desirable. The restricted model is the VAR without the mixing term and the unrestricted model is the VAR with mixing term. Since this study uses iterative nonlinear SUR, the likelihood ratio statistics can be applied for testing significance of μ and σ (Greene 1997). The likelihood ratio test statistics is

(5) $\lambda = T(\ln|W_r| - \ln|W_u|)$,

where T is the number of observations, and Wr and Wu are the residual sum of squares and cross-product matrices from the constrained and unconstrained estimators respectively. The likelihood ratio statistic is asymptotically distributed as chi-squared with degrees of freedom equal to the number of restrictions (in our case, it is 1). One limitation of such a test approach is that as Davies (1977, 1987) examined, testing for structural breaks in a case where the break point is unknown a priori is complicated by the fact that parameters characterizing the break (μ and σ) are unidentified under the null hypothesis of no structural change. Under this situation, traditional large sample theory is not applicable and thus conventional test statistics have nonstandard distributions.[6] Since transition parameters are also obtained by maximizing the value of a log-likelihood for differences in parameters (log-likelihood test: LLT), the test statistic is a sup-LLT.[7]

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[6] Hansen (1996) has developed an approach to testing the statistical significance of parameter differences across regimes in threshold autoregression models. Under his approach, simulation methods are used to approximate the asymptotic null distribution of a test of parameter differences and to identify appropriate critical value. However, it is much too complex to follow his approach since it requires running a number of simulations whereby the dependent variables are replaced by standard normal random draws. For this reason, Dr. Goodwin personally recommended the author to use the likelihood ratio test technique.

[7] Davies (1987) defined sup test statistics as $M = \sup\{S(\theta) : L \leq \theta \leq U\}$, where [L, U] is the range of possible values of θ. The test would be to reject the hypothesis for large values of M.

IV. Estimation Results

Monthly data of beef, pork, and chicken prices from Jan. 1977 to Dec. 1999 are used. Four sets of prices are examined: (i) beef retail, wholesale, and farm prices, (ii) pork retail, wholesale, and retail prices, (iii) chicken retail, wholesale, and farm prices, and (iv) beef, pork, and chicken retail prices. Lag lengths of each VAR model are determined by the Akaike Information Criterions (AIC). Stating from the longest lag length deemed reasonable, the lag length with the smallest AIC is selected. The lag lengths chosen are 2, 5, 3 and 2 for beef, pork, chicken prices, and retail prices of three livestock products respectively. Restricted and unrestricted beef price model estimated is attached in Appendix.

The strategies of obtaining parameter μ (starting point) and ρ (speed of adjustment) are as follows: First, the restricted VAR model, (equation (1)) that does not include the transition parameters (λ) was estimated. Second, the unrestricted VAR model (equation (2)) that has transition parameters was estimated. Since the unrestricted VAR model with transition parameters is non-linear, it is better to give starting values for both regimes. Starting values were obtained by splitting the data into halves and running standard VAR models. Third, the optimal mean point and the speed of adjustment parameters were estimated from an iterative Grid search,[8] which yields maximum values of log-likelihood test statistic. A two-dimensional Grid search was used to specify the transition function parameters, μ and ρ. Finally, using the estimated mean point and speed of adjustment parameters, timing and speed of adjustment graphs for each pair of prices were drawn.

Since the ultimate goal of this study is to investigate the timing and the speed of adjustment of VAR models rather than

[8] The Grid search method is to search various combinations of parameter by which an optimal value of parameter that optimizes the objective function (here, log-likelihood test statistic) is selected.

to estimate the parameter itself, only parameter estimates of beef prices are presented in Table 1.[9] Although some of parameter estimates are not statistically significant, many of the parameter estimates are highly significant and the estimates appear to fit the data very well, as is evidenced by the R^2 measures. Unrestricted model that includes transition function parameters shows the highest the adjusted R raging from 0.97 to 0.99.

As presented in Table 2, the mean point parameter for beef prices has an estimated value of 125. A standard likelihood ratio test of the significance of the differences in the standard model and the regime switching model (estimate via grid search) had the value of 64.40, which strongly rejects the null hypothesis of parameter stability using conventional chi-square critical values. The mean point parameter 125 corresponds to a significant structural shift around in May 1987.[10]

The speed of adjustment parameter for beef prices has an estimated value of 1 implying a rather fast shift. As shown in Figure 1, it took about eight months for structural change in Korean beef prices to be completed. Unfortunately, no comparable research results on structural changes in beef prices and other livestock prices are available. This structural change in beef prices may be related to the beef price shock in the mid-1980s. A significant increase in the number of Korean native cattle along with expansion of imported live calves since the late 1970s meant the beef farm price per head plunged from 2,050,000 won in 1983 to 1,210,000 won in 1986 (MAF 2000). During that time, beef retail and wholesale prices also decreased. These changes in beef (relative) prices during the mid and late

[9] Other estimation results can be provided upon on request.

[10] It needs to be mentioned that although this study finds only one mean point with speed of adjustment, it does not necessarily mean that structural change occurs only once. The Grid search method finds out the most significant parameter pair among various pairs of parameters. Of course, it is possible to identify the second (or third) significant parameter pairs that also suggest the timing of structural change and speed of adjustment.

1980s might have contributed to structural changes in beef price relationships.

The mean point parameter for pork prices has an estimated value of 171 with corresponding log-likelihood ration test statistics of 115.251. This implies that a structural break in Korean pork prices occurred in early 1991 which was four years later than in the case of beef prices. The speed of adjustment parameter has an estimated value of 5 implying that structural change in Korean pork prices started in August 1989 and ended in October 1992. As shown in Figure 2, structural change in the pork price begins in mid-1989 and completes in the early 1991.

Kim S.H. (1998) pointed out three factors contributing to Korean pork price fluctuation after the early 1990s. According to him, short-run pork price fluctuation in Korea has been driven mostly by pork supply rather than by changes in demand. Instead, increased demand from a rise in national income could affect long-run movement of pork prices. Lastly, he argued that asymmetry in price transmission between pork retail and farm markets may be caused by retailers exercising market power could also deepen pork price fluctuations. Nonetheless, for most cases, it is not easy to identify the exact contributing factors to structural changes in parameters in the model since structural change may be caused by complicated factors, occurring over periods.

The mean point parameter for chicken prices has an estimated value of 95 with corresponding log-likelihood ratio test statistic of 65.762, which also strongly rejects the null hypothesis of parameter stability. This implies that chicken prices experienced a structural break centered on November 1984. The speed of adjustment parameter has an estimate of 1. This corresponds to a rather fast rate of adjustment (Figure 3); it suggests that it took about eight months for the structural shift to be completed. Vertical integration and contract among various stages of production and marketing beginning around the early 1980s and decrease in the number of broiler chickens raised may be the factors contributing to this structural change in chicken

prices.[11]

The mean point parameter and the speed of adjustment for Korean beef, pork and chicken retail prices are 80 and 3, respectively which implies that the relationship among the three retail prices shifted during the period from August 1983 to November 1984. During that time, both beef and pork retail prices showed a decreasing or stagnant trend while chicken retail price increased slightly. Compared with other mean points, the mean point of retail prices of the three products indicates a somewhat earlier structural change. It may be possible to hypothesize that in Korea along with the increase in national income, the demand structure for livestock products started to change in the early 1980s.

V. Summary and Concluding Remarks

Recognizing that if the parameters are not constant over sample observations, the estimates may lose their significance, a gradual switching VAR model was utilized for the VAR model of Korean livestock prices to investigate the structural break point and the speed of adjustment. Monthly data covering the period from Jan. 1977 to Dec. 1999 were analyzed in applying a nonlinear iterative SUR with Grid search.

Our results suggest that VAR for Korean beef prices experienced a rather rapid structural shift in the mid 1987 while pork prices showed gradual structural break in the early 1991. Chicken prices experienced a structural change centered on late 1984. The mean point for Korean beef, pork, and chicken prices suggests that these three price categories experienced a structural

[11] Oh and Son (1998) argued that from the late 1980 large-scale broiler production sites begun to be built and vertical integration in broiler production and marketing was launched. The number of boiler raised decreased from 16,736,000 (1983) to 14,156,000 (1984). Accordingly, broiler farm price per kg rose to 1,368 won (March 1984) from 812 won (March 1983). At the same period, broiler retail price per kg increased to 1,889 won from 1,275 won (NLCF 1985).

change in the early 1980s.

Estimation results by regime imply that ignoring the structural break among variables in the model could result in different estimates. When one uses data gathered over a relatively long period, the possibility of structural change needs to be suspected and examined. If a structural break is found, data should be divided and estimated by sub-period or regime.

TABLE 1. Unrestricted and restricted VAR model: Parameter Estimates

Parameter	Unrestricted Model (Full Sample)	Restricted Model	Regime I	Regime II
br11	0.016 (0.18)	-0.024 (-0.35)	0.018 (0.15)	0.045 (0.42)
br12	0.453 (6.33)**	0.733 (12.04)**	0.518 (5.61)**	1.136 (14.67)**
br13	0.431 (6.27)**	0.236 (4.05)**	0.371 (4.16)**	-0.168 (-2.28)**
br14	0.130 (2.75)**	0.078 (2.06)**	0.132 (2.02)*	0.087 (2.38)*
br15	0.024 (0.48)	-0.053 (-1.36)	0.008 (0.12)	-0.060 (-1.63)
br16	0.554 (6.00)**	0.435 (6.87)**	0.555 (4.37)**	0.213 (4.23)**
br17	-0.592 (-6.71)**	-0.413(-6.69)**	-0.582 (-4.81)**	-0.209 (-4.24)**
bbr11	0.022 (0.15)			
bbr12	1.222 (9.44)**			
bbr13	-0.250 (-2.02)*			
bbr14	0.097 (1.56)			
bbr15	-0.076 (-1.19)			
bbr16	0.205 (2.36)*			
bbr17	-0.199 (-2.34)*			
bw21	0.385 (2.30)*	0.291 (2.34)**	0.388 (2.12)*	0.948 (2.99)**
bw22	-0.211 (-1.59)	-0.278(-2.58)**	-0.216 (-1.58)	-0.657 (-2.88)
bw23	0.226 (1.77)	0.303 (2.90)**	0.230 (1.74)	0.607 (2.80)**
bw24	0.383 (4.36)**	0.589 (8.66)**	0.384 (3.99)**	0.907 (8.45)**
bw25	0.461 (5.00)**	0.341 (4.92)**	0.458 (4.57)**	0.020 (0.19)
bw26	0.716 (4.16)**	0.583 (5.13)**	0.721 (3.83)**	0.305 (2.05)*
bw27	-0.616 (-3.76)**	-0.571(-5.15)**	-0.618 (-3.45)**	-0.286 (-1.96)*
bbw21	0.573 (2.21)*			
bbw22	-0.539 (-2.24)*			
bbw23	0.516 (2.24)*			
bbw24	0.948 (8.15)**			
bbw25	-0.002 (-0.02)			
bbw26	0.317 (1.96)*			
bbw27	-0.303 (-1.91)			
bf31	0.011 (0.10)	-0.048(-0.62)	0.012 (0.12)	0.232 (1.04)
bf32	-0.248 (-2.95)**	-0.267(-3.99)**	-0.253 (-3.24)**	-0.557 (-3.48)**
bf33	0.258 (3.19)**	0.301 (4.65)**	0.262 (3.48)**	0.541 (3.56)**
bf34	0.172 (3.09)**	0.141 (3.34)**	0.172 (3.13)**	0.113 (1.50)
bf35	-0.039 (-0.68)	-0.072 (-1.69)	-0.040 (-0.70)	-0.031 (-0.41)
bf36	1.285 (11.81)**	1.368 (19.43)**	1.288 (11.99)**	1.383 (13.27)**
bbf37	-0.438 (-4.22)**	-0.471(-6.85)**	-0.439 (-4.29)**	-0.477 (-4.67)**
bbf31	0.040 (0.23)			
bbf32	-0.486 (-3.19)**			
bbf33	0.485 (3.33)**			
bbf34	0.137 (0.07)			
bbf35	-0.047 (-0.63)			
bbf36	1.387 (13.54)**			
bbf37	-0.483 (-4.81)**			
brp Adj R-Sq	0.9966	0.9962	0.9865	0.9916
bw Adj R-Sq	0.9749	0.9735	0.9625	0.8662
bfp Adj R-Sq	0.9938	0.9937	0.9896	0.9582

Note: 1) Subscripts correspond to (i=1) retail price equation, (i=2) wholesale price equation, (i=3) farm price equation, (j=1) constant, (j=2, 3) retail price lag 1 and 2, (j=4, 5) wholesale price lag 1 and 2, (j=6, 7) farm price lag 1 and 2. br, bw and bf represent the first regime parameters and bbr, bbw, and bbf denote the second regime parameters.

2) numbers in parenthesis are t-values.

3) *(**) denotes 0.05 (0.01) significance level.

TABLE 2. Mean Point and Speed of Adjustment Parameter Estimates

	Mean Point	Speed of Adjustment	Log likelihood ratio statistics [1]
BRP/BWP/BFP	125 (May 1987)	1	64.407
PRP/PWP/PFP	171 (March 1991)	5	115.251
CRP/CWP/CFP	95 (November1984)	1	65.762
BRP/PRP/CRP	80 (August 1983)	5	78.8466

Note: 1) The critical values are x^2 (1), 0.05 =3.8414, x^2(1), 0.01=6.63491.

FIGURE 1. Timing and Speed of Adjustment for Korean Beef Prices

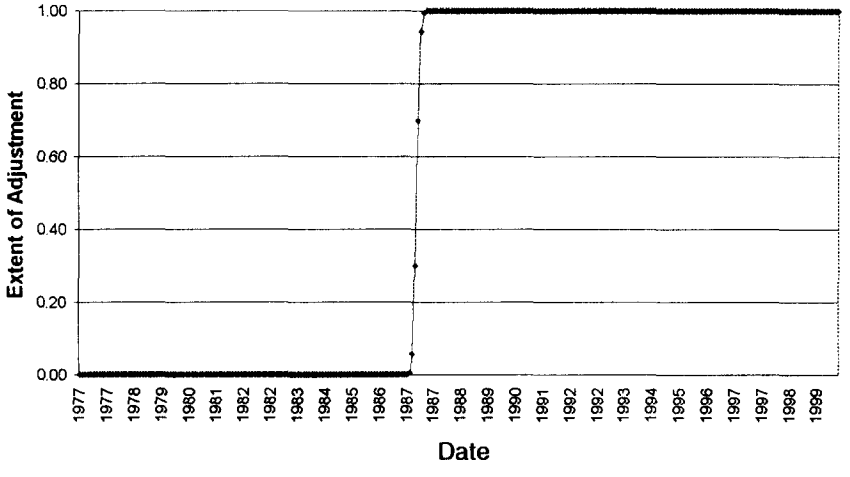


FIGURE 2. Timing and Speed of Adjustment for Korean Pork Prices

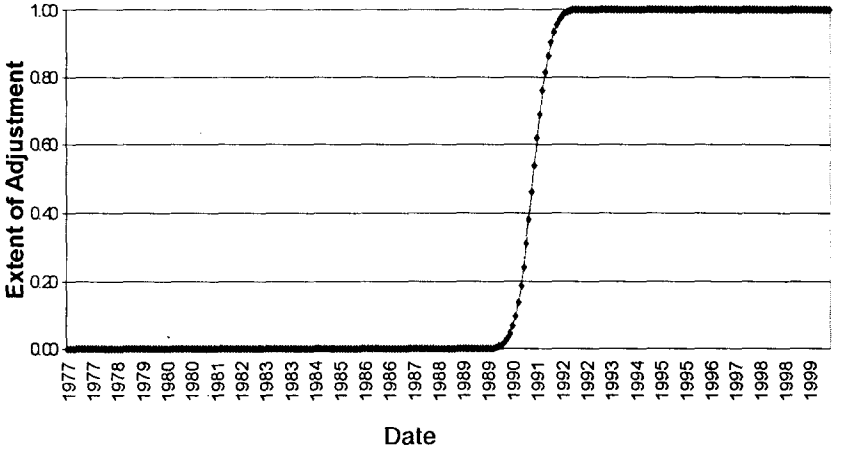


FIGURE 3. Timing and Speed of Adjustment for Korean Chicken Prices

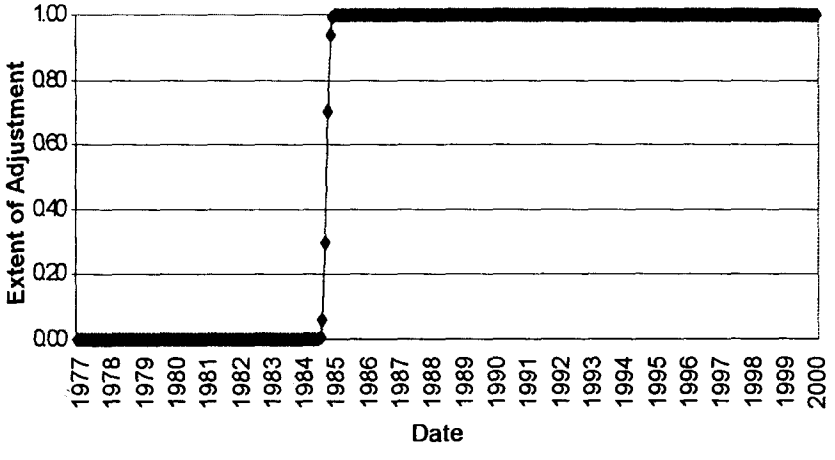
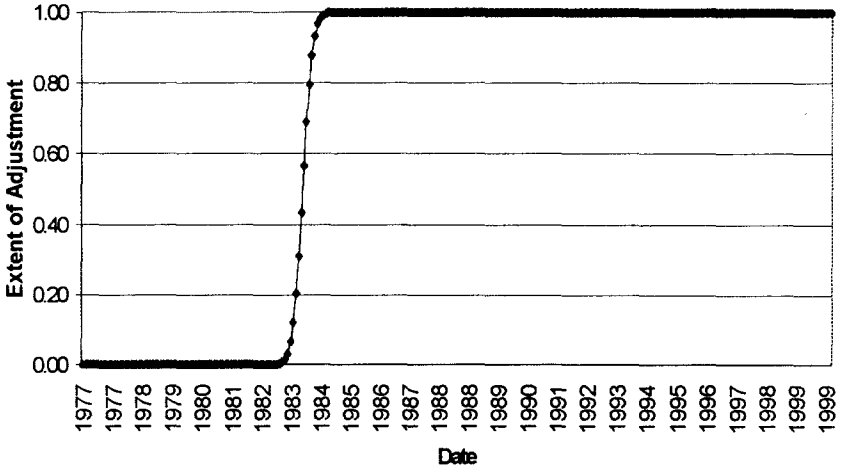


FIGURE 4. Timing and Speed of Adjustment for Korean Retail Prices of Beef, Pork and Chicken



APPENDIX

1. Unrestricted Model for Beef Prices

lbrp = (1-TRN)*(b11+b12*lag1(lbrp) + b13*lag2(lusbrp) + b14* lag1(lbwp) + b15*lag2(lbwp) + b16*lag1(lbfp) + b17*lag2(lbfp)) + TRN*(bb11 + bb12*lag1(lbrp) + bb13*lag2(lbrp) + bb14*lag1 (lbwp) + bb15*lag2(lbwp) + bb16*lag1(lbfp) + bb17*lag2(lbfp))

lbwp = (1-trn)*(b21 + b22*lag1(lbrp) + b23*lag2(lbrp) + b24* lag1(lbwp) + b25*lag2(lbwp) + b26*lag1(lbfp) + b27*lag2(lbfp)) + trn*(bb21 + bb22*lag1(lbrp) + bb23*lag2(lbrp) + bb24*lag1 (lbwp) + bb25*lag2(lbwp) + bb26*lag1(lbfp) + bb27*lag2(lbfp))
lbfp = (1-trn)*(b31 + b32*lag1(lbrp) + b33*lag2(lbrp) + b34*lag1 (lbwp) + b35*lag2(lbwp) + b36*lag1(lbfp) + b37*lag2(lbfp)) + trn*(bb31 + bb32*lag1(lbrp) + bb33*lag2(lbrp) + bb34*lag1(lbwp) +

bb35*lag2(lbwp) + bb36*lag1(lbfp) + bb37*lag2(lbfp))
Where, X = (T-TT)/NU, TRN = PROBNORM(X);
brp: beef retail price, bwp: beef wholesale price, and bfp: beef farm price

2. Restricted Model for Beef Prices

lbrp = b11 + b12*lag1(lbrp) + b13*lag2(lbrp) + b14*lag1(lbwp) + b15*lag2(lbwp) + b16*lag1(lbfp) + b17*lag2(lbfp) ;

lbwp = b21+b22*lag1(lbrp) + b23*lag2(lbrp) + b24*lag1(lbwp)+ b25*lag2(lbwp) + b26*lag1(lbfp) + b27*lag2(lbfp);

lbfp = b31 + b32*lag1(lbrp) + b33*lag2(lbrp) + b34*lag1(lbwp) + b35*lag2(lbwp) + b36*lag1(lbfp) + b37*lag2(lbfp) ;

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