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# DO CHANGES IN THE U.S. CHICKEN PRICES AFFECT KOREAN CHICKEN PRICES?

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**Key words:** price transmission analysis, VAR, forecasting error variance decomposition, structural changes in stationarity, chicken wholes prices


## ABSTRACT

Price transmission analysis relying on VAR and forecast error variance decomposition techniques was performed for the U.S. and Korean chicken wholesale prices, and for the period from 1990.1 to 2001.1. A special consideration was given to study whether the import liberalization of Korean chicken market in 1997 affected price relationships between two countries' chicken prices. It was not possible to find a statistically significant price relationship between chicken prices of the two countries, implying the import liberalization of Korean chicken market did not influence the price relationship between the two countries.


## I. Introduction

In general, we study demand-supply-price relationships in the context of price analysis (Tomek and Robinson 1990). As the sub-area of price analysis, the study of price transmission examines how price changes at one market level affect prices in

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other markets (Schwartz and Willett 1994).[1]

Schwartz and Willett (1994) also identified four main areas of study on price transmission: (a) causality - the direction of effect between changes in price from one market level to another-, (b) lags - the amount of time it takes for prices at other market levels to adjust to an initial change -, (c) asymmetry - price increase may affect other prices differently than price decreases -, and (d) market structure - the effect of market size, number of participants, product type on the price transmission process.

In addition, the degree of interrelationships among prices at different levels of markets and often, among prices of different commodities has also received much attention (Babula *et al.* 1990). The nature of price transmission has also been interpreted in the context of testing market performance or efficiency (Azzam 1999; Weaver *et al.* 1989). Understanding leads or lags is important in evaluating pricing efficiency among markets (Ward 1982).

Price transmission analysis is not only confined to domestic markets. Price transmission has featured in international trade literature as the bridge between world prices and a country's domestic price (Bolling 1988; Mohanty and Peterson 1995). The major theme of research on international price movement is to find out whether prices in different countries are closely aligned with one other, or whether they can vary more in relation to one another (Nambiar 1985).

In accordance with the Uruguay Agreement Commitment, Korea has already phased out import restrictions in the chicken market since July 1997. In this regard, changes in the U.S.

[1] The basic research questions regarding the price transmission include (a) What are the form and length of lags, if any, between changes in raw farm product prices and in retail prices? (b) Is the response to price increases and decreases symmetric? (in terms of the size and speed of price changes) and (c) Has the structure of price transmission changed with the passage of time and hence are markets more or less imperfect now than in the past? (Hall *et al.* 1981).

chicken prices may affect Korea's import demand for chicken, and in turn Korean chicken prices directly or indirectly. The United States has been the largest supplier of imported chicken to Korea, accounting for over 60% of its import.

The major purpose of this study is to determine whether there exists any relationship between the two countries' chicken prices. This will be an approach to consider cross-country effects on price transmission. If a such relationship exists, next question is how those prices are related. The answers to those questions will be pursued by analyzing impulse response function and forecast error decomposition. In doing this, special consideration will be given whether the Korean financial shock in 1997 influenced this price transmission process.

## II. Data Sources and Description

In general, the export price is based on wholesale price in the exporting country. In this context, Saghaian and Reed (2000) argued that domestic wholesale price could be understood as marginal cost of export price, while import price is also comparable with wholesale price of importing country. Accordingly, in this study, the chicken wholesale prices of the two countries will be compared. As data on transportation costs and quality differences are not readily available, following Yang et al. (2000), we will let constant term in the model capture parts of those differences.

United States' chicken wholesale price data were obtained from USDA-ERS, *Livestock, Dairy and Poultry Situation and Outlook* (http:// www.ers.usda.gov/Prodsrvs/rept-1dp.htm). Korean chicken wholesale prices were taken from the formerly National Livestock Cooperative Federation (currently part of the National Agricultural Cooperative Federation), *Monthly Survey Bulletin of the NLCF* and the NACFs statistical site (http://eis.nonghyup.com/cil-price/). The data covered the period from January 1990, to January 2001, totaling 132 observations for each price-time series. The U.S. chicken prices denoted by cents per pound were converted into Korean won per kilogram using appropriate exchange rates and tariffs of each month. Exchange rates were obtained from the Bank of Korea, http://www.bok.or.kr/bokis/m_daily. Tariff rates for chicken import were obtained from the Korea Customs Service (http://www.custom.go.kr/).

## III. Stationarity and Cointegration Tests

### 1. Stationarity Test

As the first step, a stationarity test needs to be performed since whether a time series is stationary or not determines model specification. The Augmented Dickey Fuller (ADF) test is applied for the U.S. and Korean chicken wholesale prices. The null hypothesis of stationary test is that each series has a unit root. Model incorporating constant and trend terms in the model, which is rather parsimonious to reject unit root hypothesis was tested in order to capture trend and other factors not included in the model than lagged values of own price.

Considering the potential effect of the Korean financial crisis during the late 1990s on Korean and the U.S. chicken prices denoted by Korean currency, the ADF was performed for three periods: full period (1990. 1～2001. 1), before the crisis (1990. 1～1996. 12), and after crisis (1997. 11～2001. 1). The lag lengths are determined by paring down the lags based on $t$ test statistic with the 0.01 significance level. In addition, the Ljung-Box $Q$ statistics are considered for assuring the residuals are white-noise process. The lags chosen are 10 months for the U.S. chicken wholesale price, 16 months for Korean chicken wholesale price (Table 1).

Regarding the U.S. chicken wholesale prices during the entire period (1990. 1～2001. 1), the absolute value of all estimated $t$ statistics is less than the critical value (-3.42) at the 0.05 percent implying that the U.S. chicken wholesale price series is nonstationary. Before the crisis, it is still nonstationary. Interestingly, after the crisis, it turns out stationary. More importantly, even though the U.S. chicken price series in level is nonstationaty, series in first differencing becomes stationary, implying that each price is integrated of order one I (1).

Korean chicken wholesale prices in level are also nonstationary for all three cases since the absolute values of estimated $t$ statistics are less than the critical values. Like the U.S. livestock prices, Korean chicken price in first differencing becomes stationary implying that it is also integrated of order one I (1).

There was a big jump in the U.S. chicken prices denoted by Korean currency around November 1997. This sudden jump in the U.S. price in terms of Korean currency was mainly caused by depreciation of Korean won. Due to the severe shortage of foreign currency, Korea experienced a financial crisis with its currency depreciating considerably against the U.S. dollar from 861won in January 1997 to 1,415 in December 1997. This led to approximate 60 percent increase in U.S. export price. We tested whether such shock resulted in structural change in stationarity or not.

Perron (1989) proposed a method for testing the null hypothesis of unit root allowing a one-time change in the structure occurring at a time $T_B$ $(1<T_B<T)$. He considered three different models: one that permits an exogenous change in the

**TABLE 1.** Results of Augmented Dickey-Fuller Test

| | ADF Test | | |
|---|---|---|---|
| | 90.1-2001.1 | 90.1-97.10 | 97.11-2001.1 |
| U.S. | | | |
| CWP | -2.44 | -1.93 | -4.71** |
| $\Delta$CWP | -7.68** | - | - |
| Korea | | | |
| CWP | -2.94 | -2.14 | -1.08 |
| $\Delta$CWP | -10.26** | - | - |

Note: ** denotes 0.05 (0.01) significance level.
$\Delta CWP = CWP_t - CWP_{t-1}$, CWP = Chicken Wholesale Price.

level of the series (a crash), one that permits an exogenous change in the rate of growth, and one that allows both change. Of the three models, we follow his third model in order to test one time change in the intercept and a permanent change in magnitude of the drift term. Thus the model chosen is

$$(1) \qquad Y_t = \alpha_0 + \mu_1 D_p + \mu_2 D_L + \alpha_2 t + \alpha_1 y_{t-1} + \sum_{t=1}^{k} \beta_i \Delta y_{t-1} + \varepsilon_t$$

Where $D_P$ = 1 for November1997
= 0 otherwise
$D_L$ = 1 for all $t$ beginning in November 1997
= 0 otherwise.

$D_P$ allows for a one-time change in intercept of the trend function while $D_L$ tests whether the drift parameter changes from $\alpha_0$ to $\alpha_0 + \mu_2$ at time of structural change. Regarding the dummy representing one time change in intercept, November 1997 was chosen as the break point in order to reflect not only the possible effect of the Korean financial crisis but also the effect of full-liberalization of Korean chicken import in July 31, 1997.

Under the alternative hypothesis of a permanent one-time break in trend stationary model, $\mu_1$=0, $\mu_2 \neq 0$ and $\alpha_1 < 0$, $\alpha_2 \neq 0$. Similarly, under the alternative hypothesis of a one-time change in intercept of a trend stationary process, $\mu_1 \neq 0$, $\mu_2$=0 and $\alpha_1 < 0$, $\alpha_2 \neq 0$. Perron provided the critical values for the null hypothesis $\alpha_1$=1, which depends on the proportion of observations occurring prior to the structural break $\lambda$ (= $\tau / T$ where, $\tau$ is the break point and $T$ is number of observations). For all price series, lag lengths (i.e. the values of $k$) were determined by using t-test on the coefficients $\beta_i$.

Table 2 reports Pierre Perrons test results for structural change in stationarity. As for the U.S. chicken wholesale price, the null hypotheses on one-time jump in constant term and permanent change in constant term cannot be rejected implying that there was a structural change in chicken wholesale prices denoted by Korean currency in November 1997. Since the estimated value of $\alpha_1$=0.746 is statistically different from unity, the U.S. chicken wholesale price denoted by Korean currency appears to be a trend stationary.

The null hypotheses on structural change in Korean chicken wholesale price cannot be supported. Both coefficients on $\mu_1$ and $\mu_2$ are not statistically different from zero. Moreover, since the estimated value of $\alpha_1$=0.357 is not statistically different from unity, Korean chicken price turns out to be non-stationary. In sum, the Korean financial crisis led to a change in the structure of U.S. chicken export price's stationarity denoted by Korean currency while it did not affect the stationarity of Korean chicken price at all.[2]

**TABLE 2.** Pierre Perrons Test for Structural Change

| | $T$ | $\lambda$ | $k$ | $\alpha_0$ | $\mu_1$ | $\mu_2$ | $\alpha_2$ | $\alpha_1$ |
|---|---|---|---|---|---|---|---|---|
| USCWP | 122 | 0.69 | 10 | 1.767** (4.28) | 0.117* (2.27) | 0.093** (3.57) | 0.0006* (2.17) | 0.746** (12.56) |
| KCWP | 116 | 0.72 | 16 | 4.724** (2.93) | -0.064 (-0.94) | 0.050 (0.77) | 0.003* (2.38) | 0.357 (1.63) |

Notes: 1. $T$ = number of observation
$\lambda$ = proportion of observation before the structure change
$k$ = lag length
2. The appropriate t-statistics are in parentheses. For $\alpha_0$, $\mu_1$, $\mu_2$, and $\alpha_2$, the null is that coefficient is equal to zero. For $\alpha_1$, the hypothesis is $\alpha_1$=1.
3. *(**) denotes the 0.05 (0.01) significance level.

[2] Of course, this finding assumes that exporting country (or firm) fully adjusts export price when exchange rate changes. If they do not fully adjust their export prices, export price could remain unchanged or change little. In this regard, Saghaian and Reed (2000) analyzed the relationship between exchange rate change and pricing-to-market of US meat exports. They found that export prices varied by country to some degree, and argued that exporters tended to delay adjusting export price in order to retain market share. Pick and Park (1991) also obtained similar evidence of market power in adjustment patterns of export prices for wheat and soybean products in response to exchange rate movement. Goldberg and Knetter (1997) argued that "…We are increasingly certain that deviation in the law of one price are not just a result of artifact of nonidentical goods, and incomplete pass-through… they both appear to be largely a result of third-degree price discrimination…."

Test results also suggest that when we consider shift in intercept, U.S. chicken wholesale prices turns out to be stationary while Korean chicken wholesale price remains nonstationary.

Although US chicken wholesale price with dummy was trend stationary, the ADF test suggests that both the U.S. and Korean chicken prices appeared integrated of order one implying that two prices should be estimated by the VEC. However in this study we will apply the VAR so as to distinguish statistically significant impulse responses from insignificant ones. While the standard error for impulse response function in the unrestricted VAR can be calculated using either asymptotic method or Monte Carlo repetition (the Eview reports these standard error), the standard error deviations of VEC are not available.

## 2. Cointegration Test

The components of the vector $x_t$ are said to be co-integrated of order *d, b,* denoted $x_t \sim CI(d, b)$, if (*i*) all components of $x_t$ are $I(d)$; and (*ii*) there exits a vector $\alpha (\neq 0)$ so that $Z_t = \alpha' x_t \sim I(d\text{-}b)$, b > 0. The vector $\alpha$ is called the co-integrating vector (Engle and Granger, 1987). The co-integrating vector $\alpha$ is not unique since if $\alpha' x_t$ is stationary, then so is $b\alpha' x_t$ for any nonzero scalar $b$ and if $\alpha$ is a cointegrating vector, so is $b\alpha$. Thus, it is typical that one of the variables is used to normalize the cointegrating vector by fixing its coefficient at unity. If $x_t$ has $N$ components, then there may be more than one co-integrating vector $\alpha$. It is clearly possible for several equilibrium relations to govern the joint behavior of the variables.

With $r$ linearly independent co-integrating vectors ($r \leq N\text{-}1$), the rank of $\alpha$ is $r$ which will be called the co-integrating rank of $x_t$. Cointegration means that although many developments can cause permanent changes in the individual elements of $x_t$, there are some long-run equilibrium relationships tying the individual components together. Under cointegration, the time path of the nonstationary time series must be linked. Two co-integrated series can be prices of the same commodity in different marketing channels or close substitutes in the same

**TABLE 3.** Results of Cointegration Test

| | trace tests 3) | | | max tests | | |
|---|---|---|---|---|---|---|
| | H0 | $H_1$ | $\lambda_{trace}$ | $H_0$ | $H_1$ | $\lambda_{max}$ |
| US CWP/ | r = 0 | r > 0 | 42.8276** | r = 0 | r = 1 | 41.6724** |
| ROK CWP | r ≤ 1 | r > 1 | 1.1552 | r = 1 | r = 2 | 1.1552 |

*(**) denotes rejection of the null hypothesis at the 5%(1%) significance level.

market. Many have attempted to investigate the long run relationship among economic variables using the idea of cointegration. (Enders 1998; In and Mount 1994; Chiu 1995; Clarida 1994). The Johansen cointegration test (JCT) is applied to investigate whether two price series have a long-run equilibrium relationship or not.

In order to perform the JCT, following Enders (1995), the lag lengths were determined by estimating the VAR using the undifferenced data. Of several criterions for selecting proper lag lengths, the Akaike Information Criterion (AIC) was referred to. As shown in Table 3, both trace and maximum value tests suggest that chicken wholesale prices of the two countries have common trend implying that there may exist an interrelationship between prices of two countries. The estimated long-run equilibrium relations are:

$$(2)\quad KLCWP_t - KLCWP_{t-1} = 4.283 + \underset{(-5.22)}{0.463} * (USLCWP_t - USLCWP_{t-1})$$

where, KLCWP denotes Korean chicken wholesale price and USLCWP is the U.S. chicken wholesale price, both in logarithm form. The coefficient on U.S. chicken wholesale price is statistically significant at the 0.01 level. Cointegration test results suggest that chicken wholesale prices of two countries have a statistically significant long-run equilibrium relation.[3]

[3] In this regard, the direction of causation may be mattered. The result of a block exogeneity test - a multi-variate generalization of Granger causality

# IV. Impulse Response Function and Forecast Error Variance Decomposition

## 1. Estimation Results of the Model

The VAR model has been widely used for three major purposes: forecasting, innovation accounting and error variance decomposition. Among them, the latter two are applied for searching interrelationships among variables. In order to perform innovation accounting and error variance decomposition, a VAR has to be estimated.

In a VAR system, all variables in the model are considered endogenous, that is, each variable influences itself and all other variables in the system with a certain degree of lags. A standard VAR system of order 1 (VAR (1)) can be written as follows:

(3) $$\begin{bmatrix} y_{1t} \\ y_{2t} \end{bmatrix} = \begin{bmatrix} \alpha_{10} \\ \alpha_{20} \end{bmatrix} + \begin{bmatrix} \alpha_{11} & \alpha_{12} \\ \alpha_{21} & \alpha_{22} \end{bmatrix} \begin{bmatrix} y_{1t-1} \\ y_{2t-1} \end{bmatrix} + \begin{bmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \end{bmatrix}, \text{ and}$$

(4) $$E\begin{bmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \end{bmatrix} = \begin{bmatrix} 0 \\ 0 \end{bmatrix} \quad \text{for all } t, \text{ and}$$

$$E\begin{bmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \end{bmatrix} [\varepsilon_{1t} \ \varepsilon_{2t}] = \begin{cases} \Omega & t = s \\ 0 & \text{otherwise} \end{cases}$$

where the variance and covariance matrix

(5) $$\Omega \begin{bmatrix} var(\varepsilon_{1t}) & cov(\varepsilon_{1t}, \varepsilon_{2t}) \\ cov(\varepsilon_{1t}, \varepsilon_{2t}) & var(\varepsilon_{2t}) \end{bmatrix}$$

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test - suggests that the U.S. chicken wholesale price Granger-causes Korean chicken wholesale price while Korean chicken wholesale price does not Granger-cause the U.S. chicken wholesale price. When we take into account that the United States is likely to be a large country in terms of trade theory compared with Korea, this result seems to agree with the general notion.

is assumed to be positive definite. Thus, the $\varepsilon_s$ are assumed to be serially uncorrelated but may be contemporaneously correlated. More compactly, standard VAR system of order $p$ can be defined as follows:

(6) $y_t = A + \sum_{s=1}^{p} B_s y_{t-p} + \varepsilon_t$ with $\varepsilon_t \sim i.i.d.N(0, \Omega)$

In this system of $N$ equations, $A$ is a $(n \times 1)$ vector and $y_{t-p}$ is an $N$-dimensional vector and the vector $B_s$ are $(n \times n)$ coefficient matrices. The $\varepsilon_t$ is a vector white noise process with the properties as before.

A VAR with lag length $p$ can be regarded as a reduced form of a simultaneous equation system in which each variable is regressed on a constant and $p$ of its own lags as well as on $p$ lags of each of the other variables in the VAR.[4] The right-hand side of the model contains only predetermined variables and error terms are assumed to be serially uncorrelated with constant variance. Hence, each equation on the system can be estimated using OLS.

The VAR for chicken wholesale prices of two countries are estimated.[5] In order to consider the possible effect of the Korean financial crisis and Korea's import liberalization of the chicken markets, estimation is performed for three periods: 1990. 1～2001. 1, 1990. 1～1996. 12, and 1997. 11～2001. 1. Lag length was determined based on the AIC and the Ljung-Box Q statistic criterions. Starting from the lag with the smallest AIC, lag length was pared down unless the Q statistic is greater than the critical value. Some of the parameter estimates are statistically significant,

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[4] Of course, in a near-VAR, lag lengths across equation can differ.

[5] Since both the U.S. and Korean chicken prices is shown to have unit root, it is absolutely right that two countries' chicken price series need to be estimated by other methods than VAR. If two price series are cointegrated, the Error Correction Model(ECM) may be applied. However, in this study VAR was used so as to distinguish statistically significant impulse responses from insignificant ones. Accordingly, a special consideration should be given for interpreting the results.

while the others are not.[6]

The estimation results of VAR for the U.S. and Korean chicken wholesale prices are represented in Table 4. The adjustment $R^2$s range between 0.54 and 0.96 for the entire period model, 0.37-0.79 for the first period model, and 0.26-0.77 for the second period model. All $F$-statistics are large enough to reject the null hypothesis of insignificance of parameter estimates. Based on the estimation results, impulse response function and forecast error variance decomposition are derived.

**TABLE 4.** Estimation Results of VAR for the U.S. and Korean Chicken Wholesale Prices[1)]

| | 1990.1-2001.1 | | 1990.1-1996.12 | | 1997.11-2001.1 | |
|---|---|---|---|---|---|---|
| | USLCWP | KLCWP | USLCWP | KLCWP | USLCWP | KLCWP |
| USLCWP(-1) | 1.11** (-12.99) | 0.25 (-1.09) | 0.88** (-8.31) | 0.16 (-0.42) | 1.16** (-7.34) | 0.29 (-1.10) |
| USLCWP(-2) | -0.15 (-1.72) | -0.01 (-0.05) | 0.03 (-0.32) | -0.08 (-0.20) | -0.39* (-2.53) | -0.06 (-0.21) |
| KLCWP(-1) | -0.03 (-1.05) | 0.72** (-8.24) | -0.07* (-2.28) | 0.75** (-6.83) | n0.08 (-0.76) | 0.53** (-2.89) |
| KLCWP(-2) | 0.07* (-2.31) | -0.25** (-2.93) | 0.09** (-2.92) | -0.26* (-2.44) | -0.01 (-0.12) | -0.14 (-0.76) |
| C | -0.03 (-0.18) | 2.33** (-4.59) | 0.47 (-1.08) | 3.29* (-2.06) | 1.20 (-1.42) | 2.93* (-2.06) |
| $R^2$ | 0.96 | 0.55 | 0.80 | 0.40 | 0.80 | 0.34 |
| Adjusted $R^2$ | 0.96 | 0.54 | 0.79 | 0.37 | 0.77 | 0.26 |
| $F$-statistics | 756.28 | 38.77 | 77.86 | 12.80 | 33.22 | 4.28 |

Note: *(**) denotes the 0.05(0.01) significance level.
Numbers in parenthesis are $t$ statistics.
1) From the fact that some of the values of parameters estimated are close1, one may reflect that each price series are nonstationary.

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[6] In this context, it is worthwhile to consider Doan(1992)'s argument that …"the goal of VAR analysis is to determine the interrelationships among variables, not the parameter estimate…".

## 2. Impulse Response Function Analysis

Innovation accounting refers to tracing out the systems reaction to a shock (innovation) in one of the variables. In order to perform innovation accounting and forecasting error variance decomposition, a form of the Vector Moving Average (VMA) that is expressed by lagged values of innovations is more convenient. From the VMA, impact multipliers are derived, which are used to generate the effects of shocks in innovations of each variable on the time path of the series in the system. The sets of coefficients in the VMA are called impulse response functions. The impulse response function simulates, over a period, the effects of a one-time shock in one of a VARs series on itself and on other series in the system.

A stationary VAR ($p$) process can be shown to have a moving average (MA) representation (Hamilton 1994)

$$(7) \qquad y_t = \mu + \varepsilon_t + \Psi_1 \varepsilon_{t-1} + \Psi_2 \varepsilon_{t-2 \ldots}$$
$$= \mu + \sum_{i=0}^{\infty} \Psi_i \varepsilon_{t-1}$$

where $\mu = E[y_t] = (I - B_1 - \ldots - B_p)^{-1} \varepsilon$, and $\Psi_0 = I$, $\Psi_1 = B_1$ ,

and $\Psi_i = \sum_{j=1}^{\min(p,i)} B_j \Psi_{i-j} = 1, 2, \ldots$

A plot of the row $i$, column $j$ element of $\Psi_i$,

$$(8) \qquad \frac{\partial y_i{}'t+s}{\partial \varepsilon_{jt}},$$

as a function of time $s$ is called impulse response function. It describes the response of $y_{i t+s}$ to a one-time impulse in $y_{jt}$ with all other variables dated $t$ or earlier held constant. However, this kind of analysis relying on non-orthogonal innovations whose variance-covariance matrix is not diagonal has some flaws in that it ignores the contemporaneous correlations among error terms, which may be unrealistic if the errors are not actually a white noise process (Judge *et al.* 1988; Johnston and Dinardo 1997; Hamilton 1994). For this reason, innovation accounting is often performed within a transformed VAR model in which the white-noise process has a diagonal covariance matrix so that there is no instantaneous correlation among the components.[7]

One of the methods to solve contemporaneous correlations among error terms is the Choleski decomposition.[8] As a practical matter for factorizing the covariance matrix, if we decide the order of the variables, then time series packages such as the RATS can compute the Choleski factorization. For instance, in the case of two variables, say $y_t$ and $z_t$, if we choose the order of $z_t$ is prior to $y_t$, then the Choleski decomposition constrains the system such that shock of $y_t$ has no instantaneous impact on $z_t$, but there is an indirect effect in that lagged values of $y_t$ affect the contemporaneous value of $z_t$.

In this study, a Choleski decomposition is also used. Taking into account a large country assumption on the United States, the ordering of the U.S. chicken wholesale price to Korean chicken wholesale price is chosen. The significance of each impulse response will be investigated based on *t* statistic.

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[7] Since the covariance matrix $\sum_{\varepsilon}$ of a VAR($p$) process is positive definite, there exists a nonsingular matrix $P$ such that

$$P \sum_{\varepsilon} P' = I$$

The new innovation $\nu_t = \varepsilon_t P^{-1}$ satifies $E(\nu_t \nu'_t) = I$. Such a matrix $P$ can be any solution of $P^{-1}(P^{-1}) = \sum_{\varepsilon}$. For any non-singular matrix $P$, $B_s$ in (6) can be replaced by $B_s P$ and $\varepsilon$. With this new matrix of covariance, the MA representation of $y_t$ in (7)

$$y_t = \mu + \sum_{s=0}^{\infty} B_s P P^{-1} \varepsilon_{t-s} = \mu + \sum_{i=0}^{\infty} \Gamma \nu_{t-i}$$

where $\Gamma i = B_i P$ and $v_t = P^{-1} \varepsilon_t$. The vector $v_t$ has the convenient property that its components are uncorrelated and all have unit variance.

[8] Besides the Choleski method for factorization, there are other methods such as the Eigen decomposition and structural decompositions suggested by Bernanke (1986) and Sims (1986).

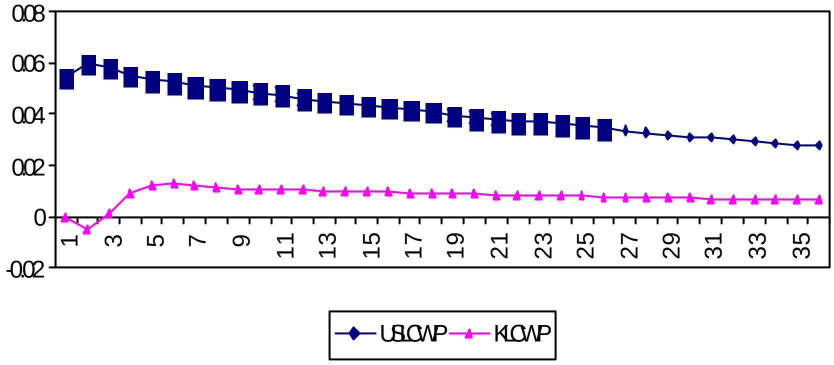
Figure 1 shows the impulse response of the U.S. chicken wholesale price. In order to distinguish significant responses from insignificant ones, responses exceeding two-standard error confidence bound are indicated by a "■" symbol. For the entire period, in response to own price shock, the U.S. chicken price takes a long response pattern of twenty-six months. Positive own price shock leads to an increase in the U.S. chicken price. In contrast, for the same period, the U.S. chicken price does not respond to the shock of the Korean chicken price. This response pattern of the U.S. chicken wholesale price seems acceptable.

In the first period (1990. 1～1996. 12), the U.S. chicken wholesale price rises for ten months in reaction to own price shock. The Korean chicken wholesale price shock results in decrease in the U.S. chicken price after two months of shock, which is counter intuition in that an increase in the Korean chicken price may also raise the U.S. chicken price if there exists a price relationship between two countries. In the second period (1997. 11～2001. 1), the U.S. chicken price reacts for four months in response to own price shock while the Korean chicken price shock does not affect the U.S. chicken price, which seems compatible with the general notion. Thus, it may be concluded that throughout the entire period, the main shock affecting the U.S. chicken price is own price rather than Korean price.

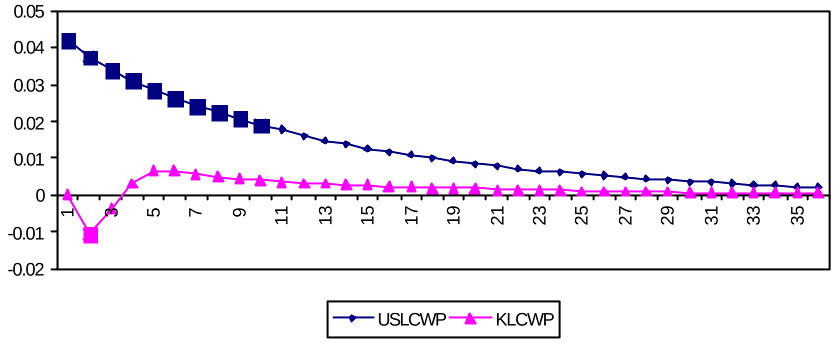
The impulse response function of the Korean chicken price is illustrated in Figure 2. For the entire period, in response to own price shock, the Korean chicken wholesale price increases sharply up to 12-16 percent for the first two months. At the third month, the Korean chicken price rises approximately 4 percent, thereafter its response dissipates. Interestingly, U.S. chicken price shock affects the Korean chicken price statistically significantly after four months of shock, and its effects on Korean chicken price last twenty-one months. For the period of 1990. 1～1996. 12, the Korean chicken price shows a similar response pattern to that in the entire period except it does not respond to changes in the U.S. chicken price. During the period between 1997. 11 and 2001. 1, in response to own price shock, Korean chicken price

**FIGURE 1.** Impulse Response Function for the U.S. Chicken Wholesale Price

[1990. 1~2001. 1]




[1990. 1~1996. 12]




[1997. 11~2001. 1]

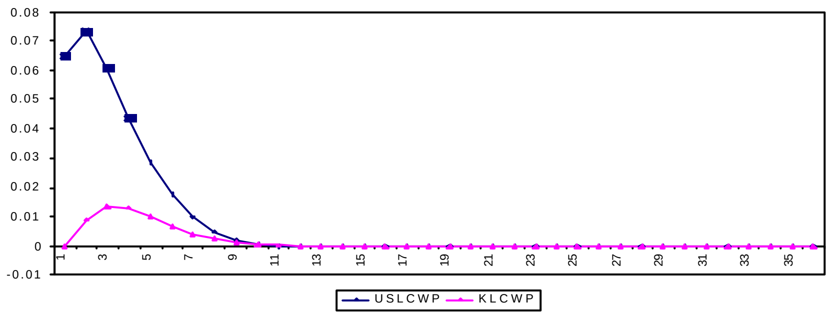
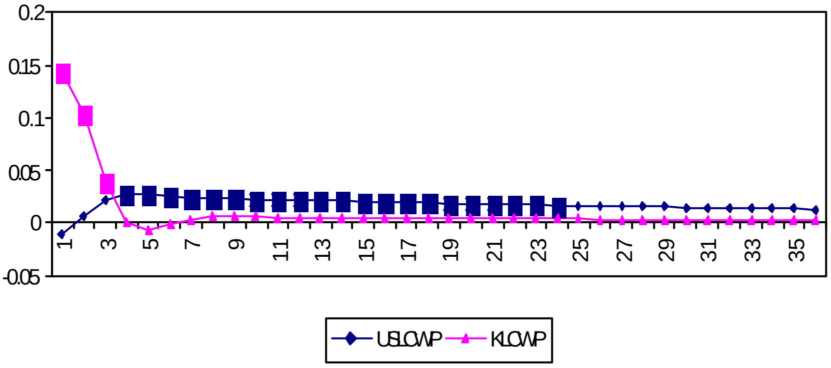
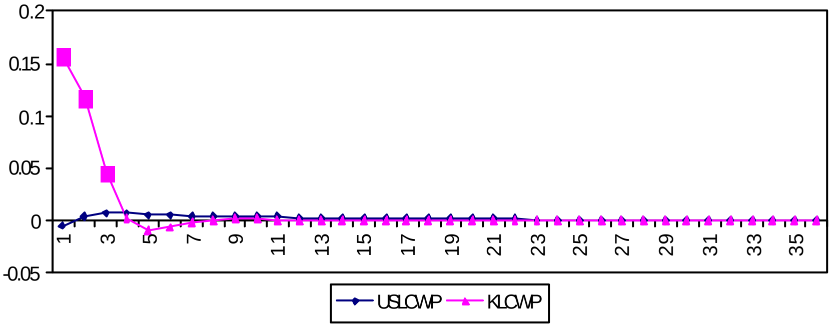


**FIGURE 2.** Impulse Response Function of Korean Chicken Wholesale Price

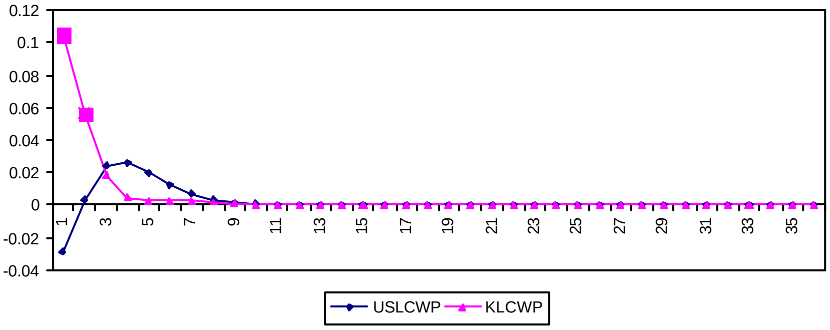
[1990. 1~2001. 1]




[1990. 1~1996. 12]




[1997. 11~2001. 1]



rises only for two months while the shock of the U.S. chicken wholesale price does not affect the Korean chicken wholesale price at all.

Considering the chicken wholesale price during the entire period, there seems to exist an interrelationship between the U.S and the Korean chicken prices, that is price movement from the United States to Korea. However, for each sub-period, the relationship between two countries' chicken prices is not found. Furthermore, there is no sufficient evidence that the Korean financial shock in 1997 affected the price relationship.

### 3. Forecast Error Variance Decomposition

The forecast error variance decomposition (FEV) is used for uncovering interrelationships among the variables in the system. Based on the vector moving average (VMA), the forecast errors and their variances can be obtained. It is also possible to decompose the $n$-step ahead forecast error variance due to each shock of variables in the model (Enders 1995). In the representation of VAR($p$) like (4), the error in forecasting a VAR $s$ periods into the future is:

(9) $$y_{t+s} - \hat{y}_{t+s|t} = \varepsilon_{t+s} + \Psi_1 \varepsilon_{t+s-1} + \Psi_2 \varepsilon_{t+s-2} + \ldots + \Psi_{s-1} \varepsilon_{t+1}$$

where $\Psi_s$ are coefficient matrices.

The mean squared error of this $s$-period-ahead forecast is

(10) $$MSE(\hat{y}_{t+s|t}) = E[(y_{t+s} - \hat{y}_{t+s|t})(y_{t+s} - \hat{y}_{t+s|t})'] = \Omega + \Psi_1 \Omega \Psi_1' + \ldots + \Psi_{s-1} \Omega \Psi_{s-1}'$$

where $\Omega = E(\varepsilon_t \varepsilon_t')$.

From $\nu_t = P^{-1} \varepsilon_t$,

(11) $$\varepsilon_t = p\,\nu_t = p_1 \nu_{1t} + p_2 \nu_{2t} + \ldots + p_n \nu_{nt}$$

where $p_j$ denotes the $j$-th column of the matrix $P$ given (2.18). Thus

(12) $\Omega = E(\varepsilon_t \varepsilon'_t) = p_1 p'_1 Var(\nu_{1t}) + p_2 p'_2 Var(\nu_{2t}) + \ldots + p_n p'_n Var(\nu_{nt})$

Substituting (12) into (10), the *MSE* of the $s$-period ahead forecast can be written as

(13) $MSE(\hat{y}_{t+s|t}) = \sum_{j=1}^{n} \{ Var(\nu_{jt}) \cdot [p_j p'_j + \Psi_1 p_j p'_j \Psi'_1 + \Psi_2 p_j p'_j \Psi'_2 + \ldots + \Psi_{s-1} p_j p'_j \Psi'_{s-1}] \}$

With this expression, it is possible to decompose the $s$-step-ahead forecast error variance due to each of the shocks. The contribution of the $j$-th orthogonalized innovation to the *MSE* of the s-period-ahead forecast is given by

(14) $Var(\nu_{jt}) \cdot [p_j p'_j + \Psi_1 p_j p'_j \Psi'_1 + \Psi_2 p_j p'_j \Psi'_2 + \ldots + \Psi_{s-1} p_j p'_j \Psi'_{s-1}]$

Dividing (2-24) by $MSE(\hat{y}_{t+s|t})$, we obtain the proportion of $MSE(\hat{y}_{t+s|t})$ due to shocks in $j$-th innovation.

If shocks to all other variables included in the system explain none of the forecast error variance of one series at all forecast horizons, we say that the series is exogenous. Likewise, a variable is highly endogenous to the system when small proportions of its FEV are attributed to own variance, and larger FEV proportions are attributed to the innovations of other variables (Bessler 1984). This FEV is used to measure the relative strength of relationships among variables.

Table 5 shows FEV of U.S. and Korean chicken wholesale price. This result could add more information about interrelation between two countries' prices analyzed previously. As expected, the U.S. chicken wholesale price is extremely exogenous for all three periods as more than 96 percent of its FEV is accounted for by own price innovations. And, there is no big difference of FEV between two periods. The Korean chicken wholesale price explains less than 4% of own price innovation for both regimes.

The Korean chicken wholesale price is also highly exogenous for all three periods with over 73 percent of its FEV being self-explained. During the period 1990. 1～1996. 12, the U.S. chicken price accounts for less than 0.7 percent of Korean chicken price implying that there is almost no relation between the two prices. However in the second period 1997. 11～2001. 1, in the short run of three months or less, the share of the U.S. chicken price in explaining FEV of Korean chicken price is 7-8 percent while at the six month or later, it increases to about 16 percent. This may imply that the interrelationship between U.S. and Korean chicken prices after the Korean financial crisis has become stronger than before.

## V. Summary and Conclusion

In order to determine whether the U.S. and Korean chicken wholesale prices are interrelated or not, the price linkage models linking the two countries' wholesale prices were analyzed. Data covers the period between 1990. 1 and 2001. 1.

The Augmented Dickey Fuller test (ADF) suggests that two countries' chicken wholesale prices have a unit root. Both prices in first difference turn out to be stationary. The Pierre Perron test for structural change indicates that there were structural changes in the U.S. chicken prices during the Korean economic crisis, implying that U.S. chicken wholesale prices are trend stationary. In contrast, Korean chicken wholesale prices have not been affected much by the Korean financial crisis.

**TABLE 5.** FEV of the U.S. and Korean Chicken Wholesale Prices

**Variance Decomposition of the U.S. Chicken Prices**

| Period | 1990.1-2001.1 | | | 1990.1-1996.12 | | | 1997.11-2001.1 | | |
|---|---|---|---|---|---|---|---|---|---|
| | S.E. | USLCWP | KLCWP | S.E. | USLCWP | KLCWP | S.E. | USLCWP | KLCWP |
| 1 | 0.05 | 100.00 | 0.00 | 0.04 | 100.00 | 0.00 | 0.06 | 100.00 | 0.00 |
| 6 | 0.14 | 97.70 | 2.30 | 0.08 | 96.75 | 3.25 | 0.13 | 96.56 | 3.44 |
| 12 | 0.18 | 96.62 | 3.38 | 0.10 | 96.40 | 3.60 | 0.13 | 96.45 | 3.55 |
| 18 | 0.21 | 96.30 | 3.70 | 0.10 | 96.35 | 3.65 | 0.13 | 96.45 | 3.55 |
| 24 | 0.23 | 96.15 | 3.85 | 0.10 | 96.33 | 3.67 | 0.13 | 96.45 | 3.55 |
| 30 | 0.25 | 96.06 | 3.94 | 0.11 | 96.32 | 3.68 | 0.13 | 96.45 | 3.55 |
| 36 | 0.26 | 96.00 | 4.00 | 0.11 | 96.32 | 3.68 | 0.13 | 96.45 | 3.55 |

**Variance Decomposition of Korean Chicken Prices**

| Period | 1990.1-2001.1 | | | 1990.1-1996.12 | | | 1997.11-2001.1 | | |
|---|---|---|---|---|---|---|---|---|---|
| | S.E. | USLCWP | KLCWP | S.E. | USLCWP | KLCWP | S.E. | USLCWP | KLCWP |
| 1 | 0.14 | 0.50 | 99.50 | 0.16 | 0.08 | 99.92 | 0.11 | 6.85 | 93.15 |
| 6 | 0.19 | 7.82 | 92.18 | 0.20 | 0.42 | 99.58 | 0.13 | 15.34 | 84.66 |
| 12 | 0.20 | 14.97 | 85.03 | 0.20 | 0.58 | 99.42 | 0.13 | 15.71 | 84.29 |
| 18 | 0.20 | 19.78 | 80.22 | 0.20 | 0.64 | 99.36 | 0.13 | 15.72 | 84.28 |
| 24 | 0.21 | 23.12 | 76.88 | 0.20 | 0.66 | 99.34 | 0.13 | 15.72 | 84.28 |
| 30 | 0.21 | 25.51 | 74.49 | 0.20 | 0.67 | 99.33 | 0.13 | 15.72 | 84.28 |
| 36 | 0.22 | 27.25 | 72.75 | 0.20 | 0.68 | 99.32 | 0.13 | 15.72 | 84.28 |

Ordering: USLCWP → KLCWP.

In order to analyze impulse response function and forecast error decomposition (FEV), three periods are investigated: entire period (1990. 1~2001. 1), regime I(1990. 1~1996. 12), regime Ⅱ (1997. 11~2001. 1). The U.S. chicken prices react to own prices as well as Korean chicken price shocks in regime I. Shocks to Korean chicken prices lead to a one percent fall in U.S. chicken prices, which seems not in agreement with general notions. For the entire analysis period, the Korean chicken wholesale price responds not only to own price shocks but also to U.S. chicken price shocks.

FEV suggests that both the U.S. and Korean chicken wholesale price is highly exogenous for all three periods, implying that each price has moved independently of each other. Interestingly, however, the Korean chicken wholesale price in regime II became less exogenous with 5-16 percent of its FEV being explained by U.S. chicken prices. This may imply that interrelationships between U.S. and Korean chicken wholesale price have become stronger after the Korean financial crisis.

The results of this analysis implies that as the livestock market is more interrelated with other countries, shocks originating from outside Korea would also be important factors influencing Korean livestock prices. Foreign factors could be buffer stabilizing domestic prices, while sometimes they may also make domestic prices more unstable.

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