Asset price movements and monetary policy in South Korea

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1. Introduction

In a market economy the prices of both financial assets and goods have the role of providing information. They give signals regarding the distribution of resources in the economy. This means that fluctuations in asset prices are useful as long as they reflect changes in expectations regarding the fundamental development of the economy. Optimistic expectations of future growth in the entire economy or in a particular industry can be expressed, for instance, in the form of rising share prices, as there are then expectations of a strong development in profits and a high real return on investment (ROI). However, in general, investors are often over-optimistic regarding future profits and asset prices rise more than is later proved to be warranted. Asset price bubbles can also be created by imperfections in the financial markets. Bonus systems designed to enable brokers to benefit when the stock market rises, while others have to bear the costs of a fall in share prices, are examples of this type of imperfection.

If households increase their consumption as a result of exaggerated increases in the value of assets, it could lead to the build-up of large real and financial imbalances. In addition, companies can increase their investments as a result of lower financing costs on a risk capital market with overvalued share prices, but also because a rise in the value of a company's assets increases its credit rating. Therefore, when the bubble bursts and asset prices fall, it could lead to heavily indebted households and companies increasing their savings considerably and to a fall in consumption and investment. Falling asset prices thus risk triggering processes that could eventually lead to a recession.

Following the widespread financial deregulation and increased globalisation of capital markets since the early 1980s, industrial economies have witnessed a clear upward trend in asset prices. However, in some cases, such as Japan and Scandinavia during the late 1980s and early 1990s, the asset price collapse after the boom turned out to have serious disruptive effects on the domestic financial system and contributed to prolonged recessions. Furthermore, as the harmful effect of large asset price movements appeared again in emerging markets in the Asian financial crisis, debate about the appropriate response of monetary policy to asset price movements has intensified recently. Nevertheless, to obtain an appropriate monetary policy response to asset price fluctuations, a better understanding of asset price movements and the linkages between asset prices and inflation is required.

This paper examines whether asset price bubbles have ever developed in South Korea and explores the relationship between changes in asset prices and inflation in order to present implications about monetary policy. Section 2 provides an interpretation of asset price changes to find whether asset price movements have been driven by fundamentals in South Korea. MRS, unit-root and cointegration tests are performed in this section. Section 3 investigates the relationship between asset price movements and inflation in South Korea to verify whether asset price movements serve as a leading indicator of upcoming inflation. A graphical analysis and an investigation of cross correlations are initially performed. Then, an empirical analysis using ordinary least squares (OLS) estimation is implemented to check the predictive power of asset price movements for future inflation. Based upon the results of the analyses given in Sections 2 and 3, Section 4 presents some implications for monetary policy with a focus on the response to asset price fluctuation. Section 5, which sets out concluding remarks, summarises the empirical results of this paper and presents some conclusions.

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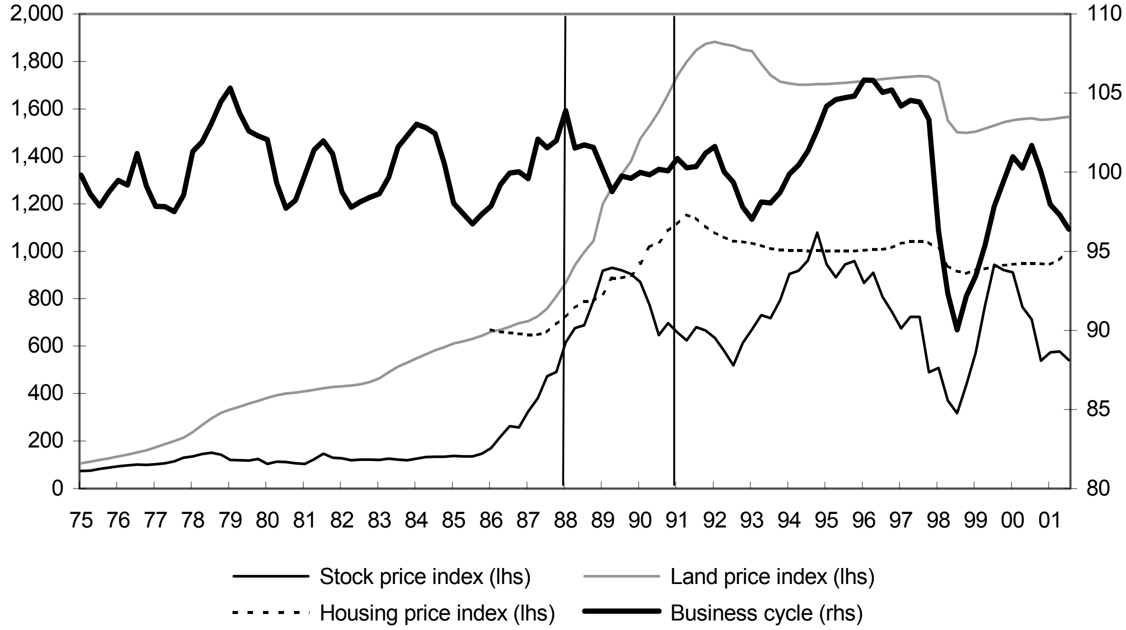
2. Interpreting asset price movements

For monetary policy to respond to asset price movements appropriately, it is crucial to distinguish asset price movements driven by economic fundamentals from asset price bubbles. However, the interpretation of asset price changes is not a straightforward task due to the various ways of deriving the asset price determined by fundamentals.

Graph 1 shows the stock, land and housing price indices and the business cycle in South Korea. The trends of these indices indicate that whereas the business cycle was moving into a trough from the late 1980s to the early 1990s, all the asset price indices were increasing. Therefore, if there have been asset price bubbles in South Korea in the recent past, this graphical information shows that they may have appeared in this period, from the late 1980s to the early 1990s.

Graph 1

Asset price indices and the business cycle



The bulk of the bubble-testing literature may be divided according to the use of one of the three types of tests. The first type examines the relationship between the observed price and the present-value price or the fundamentals used to forecast it. For example, tests of the bubble hypothesis in exchange rates - Meese (1986), Chinn and Meese (1995) and Taylor (1995) - examine the existence of a long-run equilibrium (cointegrating) relationship among the exchange rate, money supply and prices. Tests of the bubble hypothesis in stock markets - Campbell and Shiller (1987) and Campbell et al (1997) - examine the existence of equilibrium (cointegrating) relationships between prices and dividends. A second type of bubble test compares the volatilities of observed prices and the present-value prices, for example Shiller (1981), LeRoy and Porter (1981), Mankiw et al (1985) and West (1987). The third type of test is more elaborate and indirect: it estimates a reduced-form price equation by two alternative methods and verifies whether the parameter values are the same.

For the analyses in this section, the first- and second-type bubble tests[2] are chosen due to their simple structure.

[2] I would like to thank Dr Seong-Hun Yun (Research Department, The Bank of Korea) for allowing me to quote his MRS test results in “Impact of Rapid Asset Price Movements on Consumption” (2002).

2.1 Volatility tests

The bubble test initially suggested by Shiller (1981) uses volatilities of asset prices as a measure of determining whether or not the asset prices form a bubble. Shiller's test may be summarised as follows.

Consider the standard present value relation:

$$P_t = \sum_{k=0}^{\infty} \gamma^{k+1} E_t D_{t+k} \tag{1}$$

where P_t is the price of the stock (or other asset) at time t, D_{t+k} is the dividend paid at time $t+k$, E_t is the expectation conditional on information available at time t, and γ is the discount factor, or $1/(1+r)$, where r is the required rate of return.

Define P_t^* as the "perfect foresight", or "ex post rational" stock price. That is:

$$P_t^* = \sum_{k=0}^{\infty} \gamma^{k+1} D_{t+k} \tag{2}$$

P_t^* is the present value of actual, rather than expected, dividends. Since $P_t = E_t(P_t^*)$, $P_t^* = P_t + \nu_t$ where ν_t is the error in forecasting P_t^*. As a rational forecast error, ν_t is uncorrelated with information available at time t, for example P_t. Thus:

$$V(P_t^*) = V(P_t) + V(\nu_t) \tag{3}$$

where $V(x)$ is the variance of x. Therefore:

$$V(P_t^*) \geq V(P_t) \tag{4}$$

The variance of P_t^* thus presents an upper bound to the variance of the observed stock price in an efficient market. However, as there is a possibility of a non-normal distribution problem in Shiller's test with a small sample size, Mankiw et al (1985) suggest a modified volatility test (MRS test), which is immune to the problem.

Let P_t^0 be some "naive forecast" stock price:

$$P_t = \sum_{k=0}^{\infty} \gamma^{k+1} F_t D_{t+k} \tag{5}$$

where $F_t D_{t+k}$ denotes a naive forecast of D_{t+k} made at time t. This naive forecast need not be a rational one. It is important, however, that the rational agents at time t have access to this naive forecast. From this identity:

$$P_t^* - P_t^0 = (P_t^* - P_t) + (P_t - P_t^0) \tag{6}$$

We can derive an equation as follows:

$$E_t(P_t^* - P_t^0)(P_t - P_t^0) = 0 \tag{7}$$

since P_t and P_t^0 are known at time t. Squaring both sides of equation (6) and taking the expectation this produces:

$$E_t(P_t^* - P_t^0)^2 = E_t(P_t^* - P_t)^2 + E_t(P_t - P_t^0)^2 \tag{8}$$

This equality implies:

$$E_t(P_t^* - P_t^0)^2 \geq E_t(P_t^* - P_t)^2 \tag{9}$$

and

$$E_t(P_t^* - P_t^0)^2 \geq E_t(P_t - P_t^0)^2 \tag{10}$$

As equations (9) and (10) are derived under the assumption that the market is efficient, the violations of these inequalities indicate that there are bubbles in the asset prices.

2.1.1 Testing for stock prices

To verify whether there have been bubbles present in stock prices, an MRS test is performed. For the actual price, the average price of stocks traded in the market is used. Earnings per share (EPS), instead of the dividend,[3] are discounted using the discount rate to derive the ex post rational stock price. The discount rate is determined by the real rate of return, which is assumed to be 10-15%. The naive value at $t+k$ is computed under the assumption that the EPS at $t+k$ continue in the future. As can be seen from Table 1, equations (9) and (10) are violated at the discount rates of both 10% and 15%. Therefore, the existence of asset price bubbles is demonstrated by an MRS test. Despite uncertainty as to the precise period, Graph 1 strongly implies that it lasted from the late 1980s to the early 1990s.

Table 1

MRS test for stock prices[1]

Mankiw-Romer-Shapiro test[2]

$E_t(P_t^* - P_t^0)^2 \geq E_t(P_t^* - P_t)^2$ and $E_t(P_t^* - P_t^0)^2 \geq E_t(P_t - P_t^0)^2$

Discount rate	$E_t(P_t^* - P_t^0)^2$	$E_t(P_t^* - P_t)^2$	$E_t(P_t - P_t^0)^2$
10%	0.137	0.477	0.761
15%	0.092	0.232	0.352

[1] Time period: 1981-2000. [2] The actual stock price is used as a weight for the derivation to prevent the impact of the stock price level on the mean square error.

2.1.2 Testing for housing prices

Due to problems of data availability for land prices, the MRS test is applied only for housing prices. Since only the index is available for housing prices, the actual, ex post rational and naive housing prices are derived[4] using the monthly housing and house-leasing price index announced by Housing and Commercial Bank of Korea. As shown in Table 2, equation (10) is violated. Therefore, the existence of asset price bubbles is proved by the MRS test. Also, Graph 1 strongly implies that the period when the bubble existed was from the late 1980s to the early 1990s.

[3] Earnings per share are more closely related to actual stock prices than are dividends in South Korea.

[4] I assume that the ratio of the house-leasing price to the housing price is 60% and that the housing price is 1,000 in November 2001. Then, the monthly housing prices and house-leasing prices are derived using the housing and house-leasing price index. To obtain the ex post rational housing price which is the sum of the discounted monthly rent, the monthly rent is calculated from the multiplication of the house-leasing price and market interest rate. For the discount rate,

Table 2

MRS test for housing prices[1]

Mankiw-Romer-Shapiro test [2]

$$E_t(P_t^* - P_t^0)^2 \geq E_t(P_t^* - P_t)^2 \text{ and } E_t(P_t^* - P_t^0)^2 \geq E_t(P_t - P_t^0)^2$$

Discount rate	$E_t(P_t^* - P_t^0)^2$	$E_t(P_t^* - P_t)^2$	$E_t(P_t - P_t^0)^2$
1%	0.102	0.021	0.117
1%[3]	0.096	0.022	0.115

[1] Time period: January 1986-March 2001. [2] The actual housing price is used as a weight for the derivation to prevent the impact of housing price levels on the mean square error. [3] The MRS test statistics are adjusted to minimise the impact on monthly rent of the unusually high market interest rate right after the financial crisis in 1997.

2.2 Unit-root and cointegration tests

An asset price bubble may be thought of as an explosive component of the asset price which is not present in the underlying fundamentals such as the dividend and which, therefore, induces an explosive wedge between the stock price series and the underlying fundamentals. If stock prices and dividends are realisations of I(1) processes, in the absence of bubbles the standard present-value model of stock prices implies cointegration between the stock price and dividend series, implying that the difference between the stock prices and a multiple of the dividend should define a stationary process - Campbell and Shiller (1987), Diba and Grossman (1988) and Campbell et al (1997).

Consider the ex post stock return r_{t+1}, defined as:

$$r_{t+1} \equiv \log(P_{t+1} + D_{t+1}) - \log(P_t) \tag{11}$$

where P is the stock price and D is the dividend. Taking a Taylor series approximation of equation (11), Campbell et al (1997) derive the relationship:

$$r_{t+1} \approx k + \rho_{t+1} + (1-\rho)d_{t+1} - p_t \tag{12}$$

where $\rho \equiv \dfrac{1}{[1+\exp(d-p)]}$, $k \equiv -\log(\rho) - (1-\rho)\log(\dfrac{1}{\rho} - 1)$ and d–p is the average log dividend-price ratio, where lower-case letters denote the logarithms of the variables. Solving equation (12) forwards, imposing the transversality condition that:

$$\lim_{j \to \infty} \rho^j p_{t+1} = 0 \tag{13}$$

and taking expectations conditional on information at time t, we obtain:

$$p_t = \frac{k}{1-\rho} + E_t\left[\sum_{j=0}^{\infty} \rho^j [(1-\rho)d_{t+1+j} - r_{t+1+j}]\right] \tag{14}$$

Rearranging this equation, we can derive an equation for the log dividend-price ratio:

$$d_t - p_t = \frac{k}{1-\rho} + E_t\left[\sum_{j=0}^{\infty} \rho^j [-\Delta d_{t+1+j} - r_{t+1+j}]\right] \tag{15}$$

an average long-term deposit rate (1986-2001) is used. The naive housing price at $t+k$ is derived under the expectation that the future monthly rent is the same as the one at $t+k$.

Equation (5) indicates that the log dividend-price ratio will be a stationary I(0) process if and only if the stock price return series r_t is generated by a stationary process under the assumption that d_t and p_t are each generated by I(1) processes. However, testing for stationarity of the log dividend-price ratio is not eligible in the non-stationary r_t model. Therefore, equation (15) is rearranged as:

$$d_t - p_t - \frac{1}{1-\rho} r_t = \frac{k}{1-\rho} + E_t\left[\sum_{j=0}^{\infty} \rho^j [-\Delta d_{t+1+j} + \frac{1}{1-\rho} \Delta r_{t+1+j}]\right] \tag{16}$$

which only requires a test for cointegration between the log dividend-price ratio and the stock return to verify the presence of a bubble in stock prices. In the presence of bubbles in stock prices, cointegration between the log of prices and the log of dividends, or between the log dividend-price ratio and the real rate of return cannot be established.

The procedure for testing for bubbles is as follows. First, unit-root tests for stationarity of the log dividend-price ratio and the ex post rate of return are performed. Then, if r_t is tested to be non-stationary, cointegration between the log dividend-price ratio and the rate of return is examined through cointegration tests. If the log dividend-price ratio series and the ex post return series were both stationary, or if the log dividend-price ratio series and ex post returns cointegrated to a stationary series, this would suggest a rejection of the hypothesis of stock price bubbles.

2.2.1 Testing for stock prices

Results of unit-root tests for stock prices in South Korea are provided in Table 3. Whereas the MRS tests apply annual data, monthly data are used for unit-root tests. Since the I(1) null hypothesis can be rejected at the 5% level for r_t and $(d - p)_t$, without the assumption of the non-stationarity of d_t and p_t, it is concluded that there have been no bubbles in the stock prices. However, as the non-stationarity of d_t is rejected in the Phillips-Perron test and that of p_t is rejected in both the ADF and Phillips-Perron tests, the claim that there have been no bubbles is not persuasive. Furthermore, due to the stationarity of r_t, no additional cointegration tests between r_t and $(d - p)_t$ can be applied in this model. Therefore, the existence of bubbles in stock prices cannot be decided by unit-root and cointegration tests.

Table 3

Unit-root tests[1] on d_t, p_t, $(d - p)_t$ and r_t[2]

	ADF test		Phillips-Perron test	
	I[3]	II[4]	I[3]	II[4]
d_t	–1.8453	–2.1269	–6.2979*	–8.1541*
p_t	–3.0612*	–3.6714*	–6.1498*	–6.2401*
$(d - p)_t$	–4.1412*	–4.4562*	–3.2835*	–3.0090*
r_t	–7.7829*	–7.8299*	–23.4129*	–23.5522*

* Significant at the 5% level.

[1] The 5% critical values for I and II (ADF test) are –2.8777 and –3.4359 for d_t, p_t and $(d - p)_t$, and –2.8771 and –3.4350 for r_t. For the Phillips-Perron test, the 5% critical values for I and II are –2.8767 and –3.4344 for d_t, p_t and $(d - p)_t$, and –2.8768 and –3.4345 for r_t. [2] Time period: January 1986-November 2001. [3] Regression with an intercept. [4] Regression with an intercept and a linear trend.

2.2.2 Testing for housing prices

While the unit-root and cointegration tests for stock prices have been performed frequently, tests for housing prices have not been tried. Regarding housing prices and house-leasing prices, P_t and D_t, respectively, Table 4 shows the results of unit-root tests for d_t, p_t, r_t and $(d - p)_t$. Although the null hypothesis of non-stationarity for $(d - p)_t$ cannot be rejected in the ADF test and Phillips-Perron test in

Table 4, we cannot determine that there have been bubbles in housing prices due to the non-stationarity of r_t indicated by the results of the ADF test.

Table 4

Unit-root tests[1] on d_t, p_t, $(d - p)_t$ and r_t[2]

	ADF test		Phillips-Perron test	
	I[3]	II[4]	I[3]	II[4]
d_t	–2.5198	–2.0239	–2.2716	–1.8205
p_t	–3.2805*	–3.1149	–2.0121	–1.5136
$(d - p)_t$	–2.2955	–2.1419	–2.5146	–2.3278
r_t	–2.2585	–2.7076	–7.2651*	–7.2615*

* Significant at the 5% level.

[1] The 5% critical values for I and II (ADF test) are –2.8777 and –3.4359 for d_t, p_t and $(d - p)_t$, and –2.8771 and –3.4350 for r_t. For the Phillips-Perron test, the 5% critical values for I and II are –2.8767 and –3.4344 for d_t, p_t and $(d - p)_t$, and –2.8768 and –3.4345 for r_t. [2] Time period: January 1986-November 2001. [3] Regression with an intercept. [4] Regression with an intercept and a linear trend.

Therefore, to find evidence for the presence of bubbles, residual-based tests for cointegration of r_t and $(d - p)_t$ are employed. Besides the ADF test, the Phillips-Ouliaris-Hansen (1990) test is applied to consider the spurious regression problem generated by using a non-stationary process, r_t and $(d - p)_t$. Since r_t and $(d - p)_t$ are proved not to have a deterministic trend, test statistics for the model without a deterministic time trend are reported in Table 5. All the test statistics indicate that the null hypothesis of non-cointegration can be rejected at the 5% significance level. Therefore, the results of the tests suggest that there have been no bubbles in housing prices.

Table 5

Residual-based tests[1] for cointegration of r_t and $(d - p)_t$

ADF test[2]	Phillips-Ouliaris-Hansen test[2]	
	$Z\rho$[3]	Z_t[3]
–7.8261*	–33.4714*	–4.2949*

* Significant at the 5% level.

[1] The 5% critical value for the ADF test is –3.37. For the Phillips-Ouliaris-Hansen test, the 5% critical values for Z_ρ and Z_t are –20.5 and –3.37. [2] Regression with intercept. [3] See appendix for the derivation of Z_ρ and Z_t.

2.3 Implications from the results of the tests

While the results of the MRS tests performed in Section 2.1 suggest that there have been bubbles in asset prices in South Korea, the unit-root and cointegration tests performed in Section 2.2 reject the presence of bubbles in some asset prices, for example housing prices. Therefore, the results of tests depend heavily on the frequency of data and the type of test.

Even though the graphical information in Graph 1 suggests that the bubbles could exist from the late 1980s to the early 1990s, empirical tests do not provide us with any decisive information about the presence of bubbles. As mentioned before, this implies that distinguishing bubbles from the fundamentals is not a straightforward task.

3. Change in asset prices and inflation

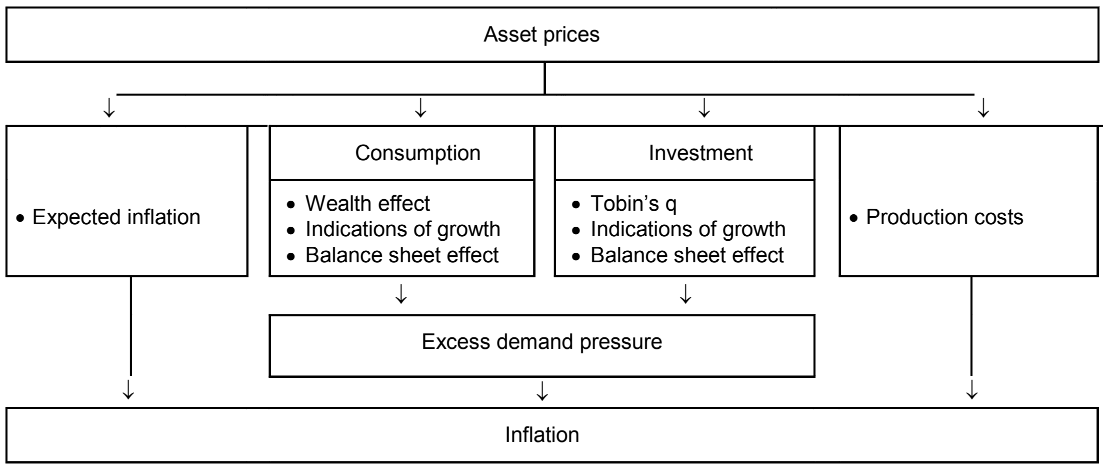
A private agent in an economy has various assets such as bonds, equities and foreign exchange categorised as financial assets and a house and land grouped as real estate. Recent research - Borio et al (1994), Goodhart and Hofmann (2000) and Ray and Chatterjee (2001) - suggests that asset price fluctuations have predictive content for future inflation, such as the expected inflation rate reflected in asset price movements, and induce future inflation via transmission channels running from asset prices to inflation.

First, the predictive power of asset price movements for future inflation can be understood by the Fisher equation. The Fisher equation, $i = r + \pi^e$, where i, r and π^e denote the nominal interest rate, the real interest rate and the expected inflation rate, respectively, shows that the information content of financial asset prices as embodied in nominal interest rates can be assessed in the absence of risk premia and money illusion and with r as a constant. Assuming a relationship can be drawn between the expected inflation rate and the actual future inflation rate, the one-to-one relationship between the nominal interest rate and expected inflation rate in the Fisher equation implies information about the future inflation rate. Some empirical studies - Fama (1977) and Mishkin (1990a) - have generally confirmed that current asset prices, for example nominal interest rates, provide reliable forecasts of future inflation up to a certain period.

Second, it has been argued that the impact of asset price movements is transmitted to inflation via various channels (see Figure 1). The two main channels are private consumption and investment. Rising asset prices, such as stock or property prices, affect private consumption by raising lifetime wealth, signalling higher expected wage incomes and increasing the value of collateral, which influences the borrowing capacity of private agents. An increase in asset prices affects investment by lowering the cost of new capital relative to existing capital (Tobin's q), providing an impetus to current investment based on expected future growth of output (the "flexible accelerator" model) and improving banks' balance sheets and thus inducing banks to lower interest charges on loans. Empirical evidence - Kent and Lowe (1997) and Browne et al (1998) - confirms the impact of stock and property prices on private consumption and investment in industrialised countries, although the magnitude of the effect varies, depending on the share of the assets in national wealth and the nature of corporate and banking laws.

Figure 1

A transmission mechanism of asset prices to inflation



Based on the implications suggested in the above discussion, the relationship between asset price movements and the inflation rate in South Korea is investigated in the following subsections. For this purpose, recent trends in the nominal interest rate, interest rate spreads, stock prices and property prices are first reviewed. For stock prices and property prices, their troughs and peaks are compared with those of the inflation rate. In addition, cross correlations between asset prices and the inflation rate are checked. The predictive power of asset price movements for the inflation rate is examined by the estimation of the inflation forecasting models in the final subsection.

3.1 Relationship between asset prices and the inflation rate[5]

3.1.1 Nominal interest rate and interest rate spreads

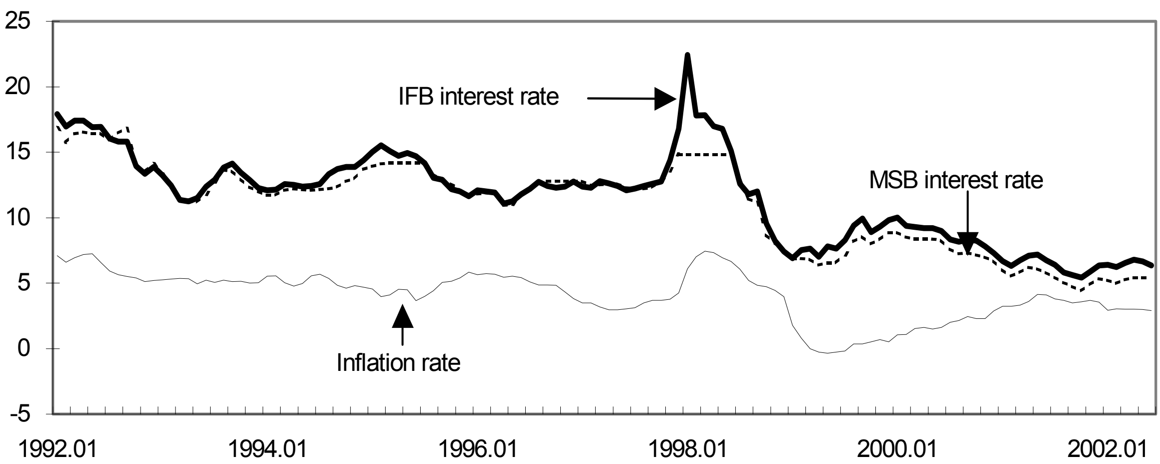
The period selected for investigation in this analysis is 1992-2001 in view of the fact that the first stage of interest rate deregulation was carried out in late 1991. For long-term interest rates, yields on monetary stabilisation bonds (one year) and industrial finance bonds (three years) are employed. For short-term interest rates, yields on certificates of deposit (CDs) (91 days) are used. A modified consumer price index (CPI) excluding farm products[6] and petroleum prices is employed in this subsection to minimise the effect of exogenous components, such as weather and foreign product prices, on the inflation rate index.

The trends of the nominal interest rate and the inflation rate in Graph 2 do not reveal any graphically detectable correlation. In particular, the nominal interest rate does not lead the inflation rate during the period 1992-2001. However, the trend in the interest rate spread in Graph 3 indicates that it has led the change of the inflation rate[7] by approximately four quarters since the mid-1990s.

Graph 2

Trends in the nominal interest rate and the inflation rate

In percentages



Note: IFB = industrial finance bonds; MSB = monetary stabilisation bonds.

The graphical analysis performed in the above discussion shows that the interest rate spread has more information about the future inflation rate than the nominal interest rate.

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5 I would like to thank Mr Jong Wook Kim (Research Department, The Bank of Korea) for allowing me to cite his paper "Analysis of change in asset prices as predictors of inflation" (2002).

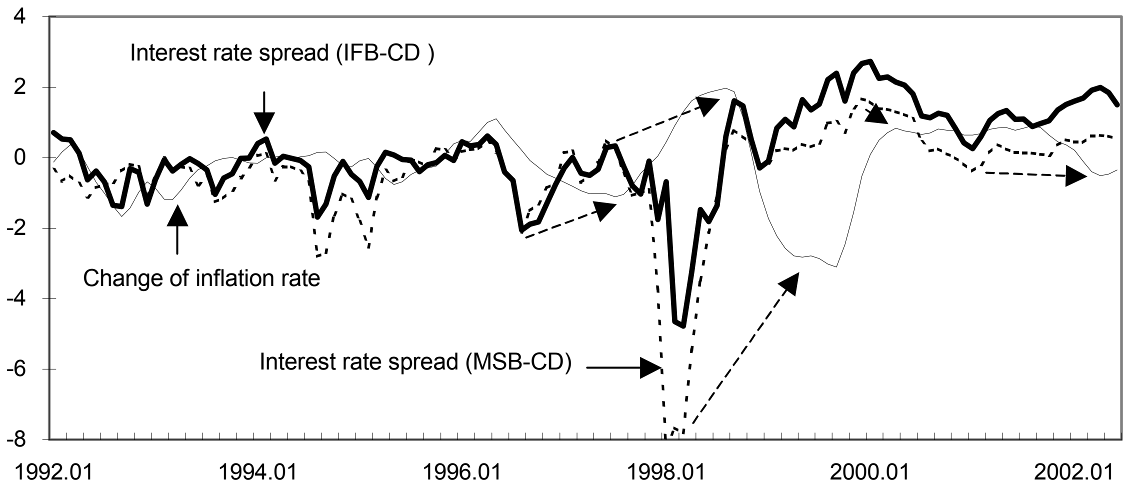
6 Only grain prices are included in this modified CPI (core inflation).

7 Change of the inflation rate is defined as the gap between the average inflation rate in the previous four quarters and the inflation rate four quarters before, based on the inflation forecasting equation - Mishkin (1990a, 1990b).

Graph 3

Trends in interest rate spread and change of inflation rate

In percentages



To confirm this result, cross correlation coefficients between the nominal interest rate and the inflation rate and between the interest rate spread and the change of the inflation rate are derived. Table 6 indicates that while the cross correlation coefficient between the nominal interest rate and the inflation rate is large in the same lag but becomes smaller as the time lag increases, the other cross correlation coefficient between the interest rate spread and the change of the inflation rate turns to a positive value from a negative one as the time lag increases by more than four quarters. Recent trends in the interest rate spread and the change of the inflation rate show an apparent positive cross correlation in five-quarter time lags. These results also imply that the interest rate spread is suitable as a leading indicator of future inflation.

Table 6

Cross correlation coefficients[1]

	0	**1**	**2**	**3**	**4**	**5**	**6**	**7**	**8**	**9**	**10**	**11**	**12**
INF, MSB	0.71	0.67	0.56	0.41	0.29	0.17	0.10	0.01	–0.04	–0.08	–0.08	–0.05	–0.01
INF, IFB	0.71	0.68	0.54	0.38	0.19	0.06	0.00	–0.06	–0.09	–0.11	–0.10	–0.04	0.02
INF, SP1[2]	–0.54	–0.51	–0.34	–0.13	0.18	0.28	0.25	0.24	0.25	0.26	0.26	0.19	0.03
1992-1997	0.07	0.15	0.21	0.04	–0.19	–0.39	–0.39	–0.20	0.05	0.21	0.32	0.30	0.13
1998-2001	–0.71	–0.65	–0.43	–0.14	0.31	0.44	0.40	0.34	0.26	0.19	0.11	0.04	–0.07
INF, SP2[3]	–0.71	–0.62	–0.47	–0.26	–0.04	0.04	0.05	0.11	0.18	0.25	0.30	0.26	0.14
1992-1997	0.23	0.36	0.37	0.22	–0.01	–0.23	–0.30	–0.16	0.01	0.09	0.17	0.10	–0.09
1998-2001	–0.82	–0.65	–0.37	–0.02	0.39	0.50	0.45	0.37	0.27	0.17	0.08	0.00	–0.10

[1] Cross correlation coefficients between the inflation rate (*INF*) in period *t* and the yield on MSBs (*MSB*), the yield on IFBs (*IFB*) in period *t – i* and the interest rate spread (*SP1*, *SP2*) for 1992:1-2001:4. [2] *SP1*: yield on MSBs (one year) - yield on CDs (91 days). [3] *SP2*: yields on IFBs (three years) - yields on CDs (91 days).

3.1.2 *Stock prices*

The relationship between stock prices and the inflation rate is investigated using a systematic cyclic analysis[8] for time series considering the close correlation between stock prices or the inflation rate and the business cycle. The troughs and peaks of the rate of change in the stock price index (*KSP*) and the inflation rate in Table 7 show that *KSP* and the inflation rate have had three and five cycles, respectively, since the mid-1980s (see also Graph 4). The cycles in Table 7 reveal that troughs and peaks of *KSP* lead those of the inflation rate by four to eight quarters.

Table 7

Cycles of the inflation rate and the rate of change in the stock price index (*KSP*)

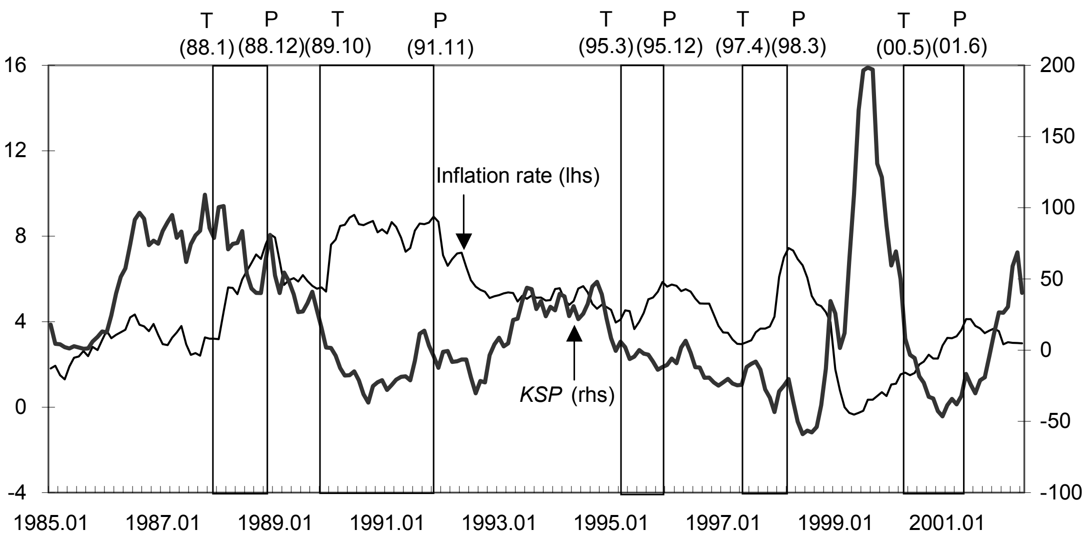
Inflation rate			*KSP*			
Cycle	**Trough**	**Peak**	**Trough**	**Lead period[1]**	**Peak**	**Lead period[1]**
1st	Jan 1988	Dec 1988	Sep 1985	28	–	–
2nd	Oct 1989	Nov 1991	–	–	Mar 1989	32
3rd	Mar 1995	Dec 1995	Aug 1992	31	Nov 1994	13
4th	Apr 1997	Mar 1998	–	–	–	–
5th	May 2000	Jun 2001	Jun 1998	23	Dec 1999	18

[1] In months.

Graph 4

Trends in the inflation rate and the rate of change in the stock price index (*KSP*)

In percentages



[8] The Bry-Boschan analysis method is used for detecting cycles applying enough time series data for the analysis. The inflation rate is deseasonalised by X-12 ARIMA.

The cross correlations between *KSP* and the inflation rate and between *KSP* and excess real demand pressure (*GAP*) are derived in Table 8. The cross correlation between *KSP* and excess real demand pressure indicates an indirect impact of *KSP* on the inflation rate via excess real demand caused by an increasing wealth effect and a forthcoming growth effect.

KSP in Table 8 shows a positive correlation with the GDP gap rate[9] and with the inflation rate after two and seven quarters, respectively. This implies that the fluctuation of *KSP* affects the inflation rate through changes of real demand.

Table 8

Cross correlation coefficients between *KSP* and other macroeconomic variables[1]

	0	1	2	3	4	5	6	7	8	9	10	11	12
GAP, KSP	0.08	0.21	0.28	0.30	0.29	0.22	0.18	0.13	0.09	0.10	0.11	0.12	0.13
INF, KSP	–0.39	–0.33	–0.23	–0.11	–0.02	0.09	0.19	0.29	0.36	0.41	0.44	0.43	0.42

[1] Cross correlation between the inflation rate (*INF*) or GDP gap rate (*GAP*) and the rate of increase of the stock price index (*KSP*) for 1985:1-2002:4.

3.1.3 Real estate prices

The troughs and peaks of the rate of change in housing prices (*KHP*) and the inflation rate in Table 9 obtained by systematic cyclic analysis show that *KHP* and the inflation rate have had four and five cycles, respectively, since the mid-1980s (see also Graph 5). In addition, the cycles in Table 9 indicate that the troughs and peaks of *KHP* lead those of the inflation rate by two to five quarters.

Table 9

Cycles of the inflation rate and the rate of change in housing prices (*KHP*)

Inflation rate			*KHP*			
Cycle	Trough	Peak	Trough	Lead period[1]	Peak	Lead period[1]
1st	Jan 1988	Dec 1988	Aug 1987	5	–	–
2nd	Oct 1989	Nov 1991	–	–	Apr 1991	7
3rd	Mar 1995	Dec 1995	Oct 1993	17	Dec 1994	12
4th	Apr 1997	Mar 1998	Jan 1996	15	Oct 1997	5
5th	May 2000	Jun 2001	Nov 1998	18	Nov 1999	19

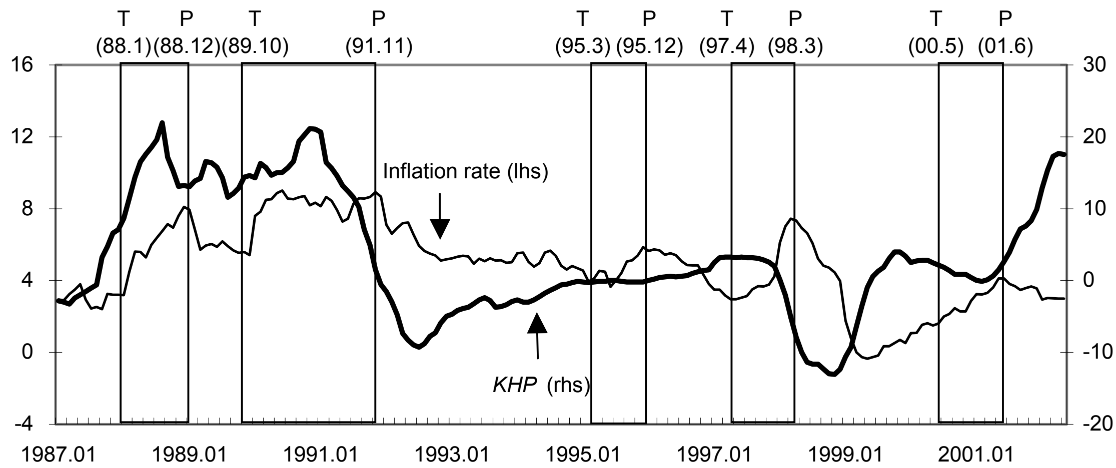
[1] In months.

[9] GDP gap rate = (actual GDP/potential GDP − 1) × 100.

Graph 5

Trends in the inflation rate and the rate of change in housing prices (*KHP*)

In percentages

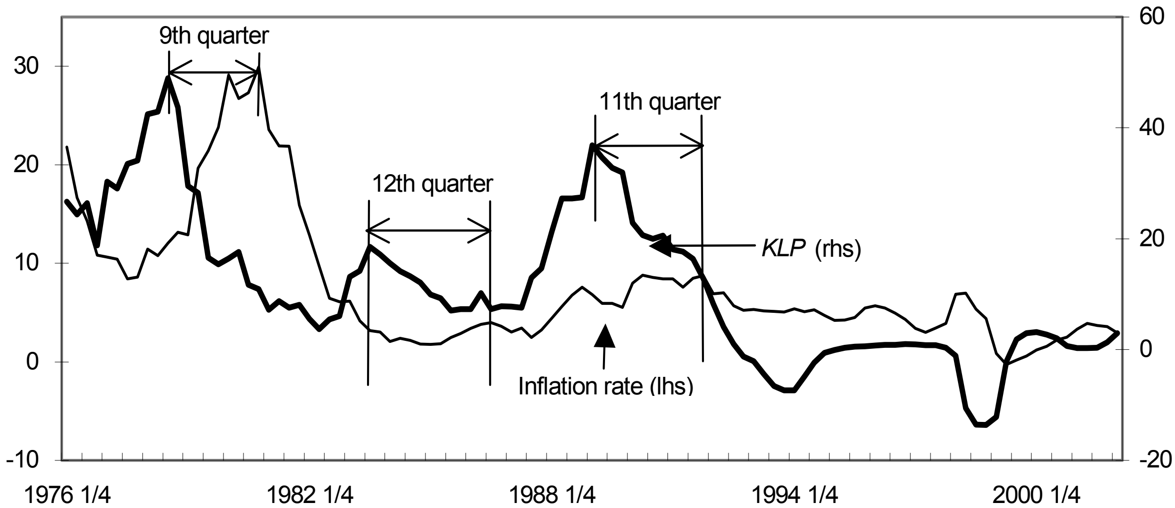


For land prices, the analysis applied for *KSP* and *KHP* is not employed since they do not show any noticeable fluctuation in the short run and have no high-frequency data (monthly). Therefore, the trends in the rate of increase of land prices (*KLP*) and the inflation rate are compared in Graph 6. *KLP* shows steep increases in 1978, 1983 and 1989. The inflation rate then shows an upward trend mirroring each increase of *KLP*, two to three years later.

Graph 6

Trends in the inflation rate and the rate of change in land prices (*KLP*)

In percentages



The cross correlations between *KHP* or *KLP* and *GAP* or the inflation rate are derived in Table 10. The positive cross correlations between *KHP* or *KLP* and *GAP* are largest in the same quarter. However, for the inflation rate, it becomes largest in the fourth or fifth quarter. These results imply that the impact of real property price movements on the inflation rate may be effected via the cost transmission channel as well as the real demand transmission channel.

Table 10

Cross correlation coefficients between *KHP* or *KLP* and other macroeconomic variables[1]

	0	1	2	3	4	5	6	7	8	9	10	11	12
GAP, KHP	0.38	0.34	0.25	0.13	0.02	–0.04	–0.11	–0.14	–0.14	–0.14	–0.13	–0.13	–0.14
INF, KHP	0.39	0.56	0.67	0.72	0.71	0.65	0.60	0.56	0.52	0.48	0.43	0.37	0.32
GAP, KLP	0.30	0.24	0.17	0.08	0.00	–0.05	–0.09	–0.10	–0.10	–0.09	–0.08	–0.09	–0.11
INF, KLP	0.45	0.56	0.65	0.72	0.74	0.74	0.72	0.70	0.68	0.63	0.60	0.54	0.48

[1] Cross correlation between the inflation rate (*INF*) or GDP gap rate (*GAP*) and the rate of increase of housing prices (*KHP*) or the increasing rate of land prices (*KLP*) for 1987:1-2001:4.

3.2 Asset price movements as a leading indicator of inflation

As the interest rate spread, *KSP*, *KHP* and *KLP* were shown to have cross correlations with the inflation rate in the above subsections, an empirical analysis is performed to check whether or not asset price movements have predictive power for future inflation.

3.2.1 The model and estimation

A multivariate inflation forecasting model without asset prices is set as a basic model as in Borio et al (1994) or Goodhart and Hofmann (2000). The basic model, which is a reduced-form equation, is as follows:

$$\pi_{t+i} = \alpha + \beta_0 \pi_{t-j} + \beta_1 \Delta WG_{t-j} + \beta_2 \Delta M_{t-j} + \beta_3 GAP_{t-j} + \beta_4 \Delta PIM_{t-j} + \varepsilon_{t+j} \quad (17)$$

with $i = 4, 6, 8$ $j = 0, 1, 2, \ldots$

where π is inflation, ΔWG is the rate of change in unit labour costs, ΔM is the rate of change in total liquidity, *GAP* is the rate of the GDP gap, and ΔPIM is the rate of change in import goods prices. The forecasting horizons are assigned as four, six and eight quarters based on the cross correlations between asset prices and the inflation rate.

With so many regressors, the question of how to determine lag lengths becomes crucial. As a first measure to save degrees of freedom, only information of the current and the previous year is taken into account when assessing future inflation. However, this measure alone would still have left 41 coefficients to be estimated in the basic model. Therefore, to economise on the number of freely estimated parameters, the average of lags 0-3 and 4-7 is included for each variable in the regression. This also avoids arbitrary restriction on the lag lengths.[10] In addition, significance was assessed on the basis of Newey-West autocorrelation and heteroskedasticity-consistent standard errors.

The estimated coefficients of the basic model in Table 11 indicate that the coefficient of the GDP gap rate (*GAP*) is significantly different from zero at the 5% level at four-quarter horizons but not at six- and eight-quarter horizons. The coefficient of the rate of change in import goods prices (ΔPIM) also implies that the impact of the fluctuation of import goods prices on the inflation rate is not strong enough to be significant. However, the coefficients of the rates of change in unit labour costs and in total liquidity are significantly different from zero at the 5% level.

[10] See Goodhart and Hofmann (2000).

Table 11

Estimated coefficients in the basic model[1]

Forecasting horizon (Dependent variable)[2]	Independent variable at time *t*					$\bar{R}^2$
	ΔWG	ΔM	GAP	ΔPIM	α	
4 (π_{t+4})	0.23* (5.75)	0.12* (4.11)	0.18* (2.53)	0.03* (1.71)	0.50* (1.98)	0.88
6 (π_{t+6})	0.16* (4.32)	0.17* (4.56)	0.05 (1.18)	–0.05* (3.75)	0.19 (1.31)	0.86
8 (π_{t+6})	0.12* (1.87)	0.14* (2.61)	–	–0.07* (3.28)	1.23 (1.66)	0.77

* = significance at the 5% level; *t*-statistics in brackets.

[1] Time period: 1988:1-2001:4. [2] In quarters.

Based on the estimation of the basic model, an extended model, which is the basic model with each asset price variable, is estimated to check the predictive power of each asset price for the inflation rate. Table 12 presents the estimated coefficient for each variable. The third column shows that the interest rate spread has predictive power for future inflation at four-quarter horizons. The fifth column reveals that housing prices provide effectual information for future inflation at four- and six-quarter horizons. In addition, as shown in the eighth and 10th columns, land and stock prices are suitable as leading indicators of future inflation at six- and eight-quarter horizons. These results coincide with the implications presented by the analyses performed in the previous subsections.

Table 12

Estimated coefficients in the extended model[1]

Forecasting horizon (Dependent variable)[2]	Basic model	Added asset price variable at time *t*							
		Interest rate spread[3]		*KHP*		*KLP*		*KSP*	
	$\bar{R}^2$	Co-efficient	$\bar{R}^2$	Co-efficient	$\bar{R}^2$	Co-efficient	$\bar{R}^2$	Co-efficient	$\bar{R}^2$
4 (π_{t+4})	0.88	0.80* (2.62)	0.90	0.06* (2.61)	0.90	0.03 (1.49)	0.88	0.00 (0.15)	0.88
6 (π_{t+6})	0.86	0.63 (1.55)	0.87	0.04* (1.80)	0.87	0.04* (2.00)	0.87	0.01* (2.44)	0.88
8 (π_{t+6})	0.77	–0.56 (0.87)	0.77	0.00 (0.06)	0.77	0.05* (1.81)	0.79	0.02* (2.62)	0.82

* = significance at the 5% level; *t*-statistics in brackets.

[1] Time period: 1988:1-2001:4. [2] In quarters. [3] Time period: 1992:1-2001:4.

3.2.2 *Simulation and ex ante forecasting*

To ascertain how greatly asset prices contributed to the forecast of the inflation rate in 1990-2001, an ex post simulation is performed with significant asset price variables at four- and eight-quarter horizons.

In general, as can be seen from Table 13, the forecasting error significantly decreases in the forecasts with asset prices considering the mean absolute error (*MAE*) and the root mean square error (*RMSE*). While the forecasting error in the basic model is 0.56% (*MAE*), it decreases by 0.07 percentage points per quarter with *IRS* and by 0.05 percentage points per quarter with *KHP* at four-quarter horizons. At eight-quarter horizons, also, the forecasting error decreases by 0.10 percentage points with *KSP* and by 0.05 percentage points with *KLP*.

However, the predictive power of the model depends on the time period of the data. In particular, whereas the improvement in forecasting ability is considerable in 1990-93 owing to the addition of asset prices, it is not in 1994-97. In 1998-2001, only the interest rate spread and *KHP* contribute to the forecast of the inflation rate. When the excess real demand induces ongoing inflation, asset price movements serve as a leading indicator. However, in the case of inflation induced by high costs, the information from asset price movements for future inflation is not useful.

Table 13

Results of simulation[1]

Forecasting horizon	Added variable	1990-93		1994-97		1998-2001		1990-2001	
		RMSE	***MAE***	***RMSE***	***MAE***	***RMSE***	***MAE***	***RMSE***	***MAE***
Fourth quarter	(Basic model)	0.78	0.70	0.65	0.45	0.71	0.55	0.71	0.56
	IRS[2]	–	–	0.63	0.48	0.58	0.50	0.61	0.49
	KHP	0.53	0.49	0.63	0.49	0.66	0.54	0.61	0.51
Eighth quarter	(Basic model)	1.21	1.03	0.74	0.67	1.28	0.94	1.10	0.88
	KSP	0.70	0.61	0.73	0.65	1.25	1.07	0.93	0.78
	KLP	0.99	0.81	0.73	0.64	1.46	1.03	1.10	0.83

[1] $RMSE = \sqrt{\frac{1}{T}\sum_{t=1}^{T}(Y_t^s - Y_t^a)^2}$, $MAE = \frac{1}{T}\sum_{t=1}^{T}\left|Y_t^s - Y_t^a\right|$, where Y_t^s is a forecasted value, and Y_t^a is an actual value for time *t*.

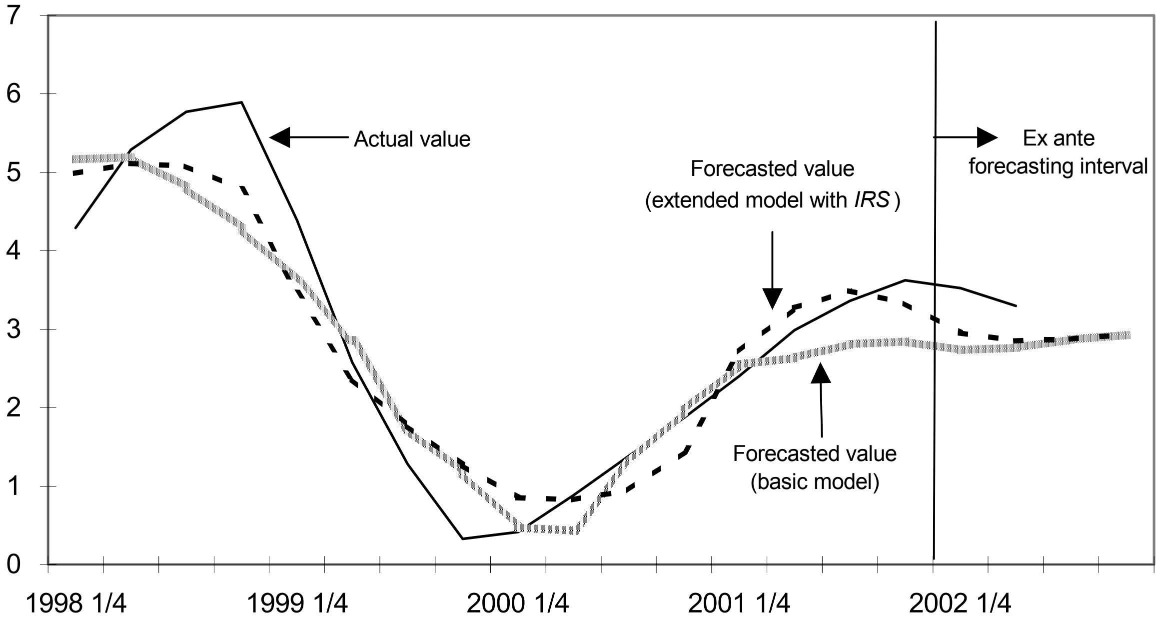
[2] Interest rate spread.

As shown in Graphs 7 and 8, ex ante forecasting for the four-quarter horizon indicates that the values forecasted by the extended models approach the actual values more closely than the values forecasted by the basic models. This implies that while the basic model does not adequately capture the impact of asset price movements, for example rising housing prices, on the inflation rate, the extended model reflects the impact of the fluctuation of asset prices on the inflation rate owing to the addition of housing prices, *KHP*. However, ex ante forecasting for the eight-quarter horizon did not show any crucial improvement in the predictive power through the addition of asset prices (Graphs 9 and 10).

Graph 7

Actual and forecasted values of the interest rate spread[1]

In percentages

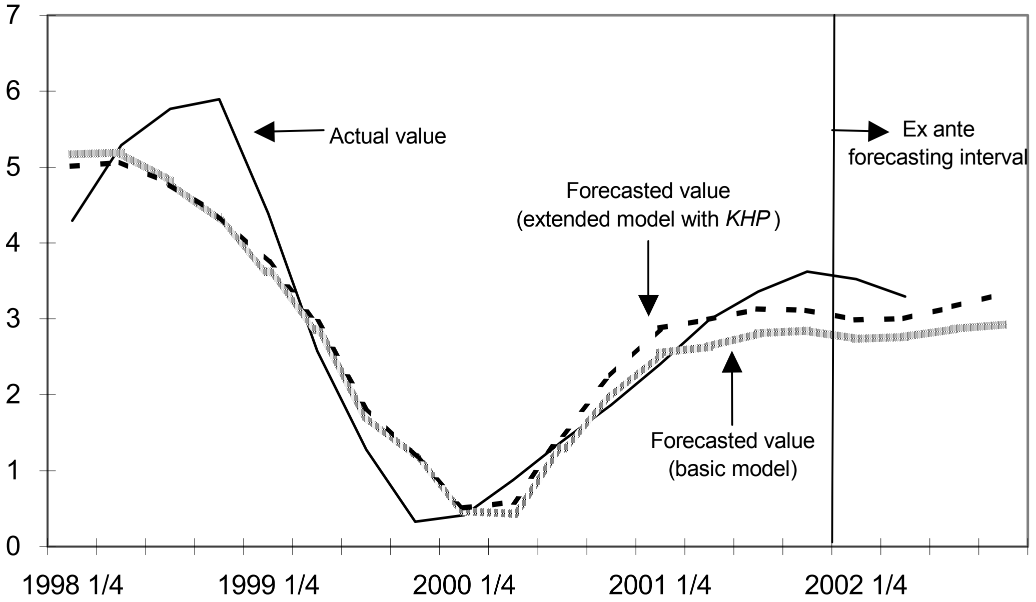


[1] Forecasting horizon is four quarters.

Graph 8

Actual and forecasted values of housing prices (*KHP*)[1]

In percentages

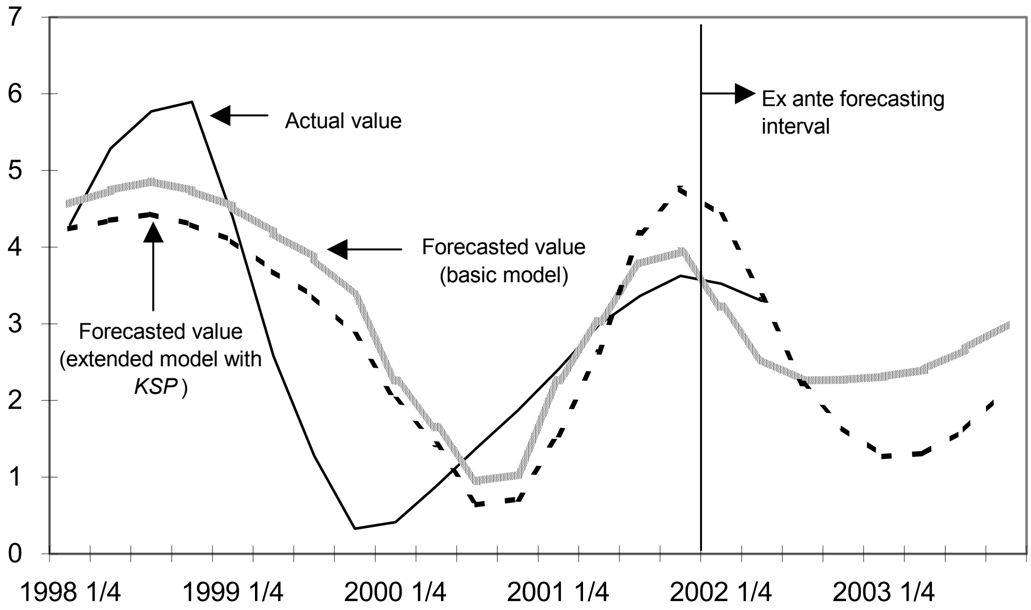


[1] Forecasting horizon is four quarters.

Graph 9

Actual and forecasted values of the stock price index (*KSP*)[1]

In percentages

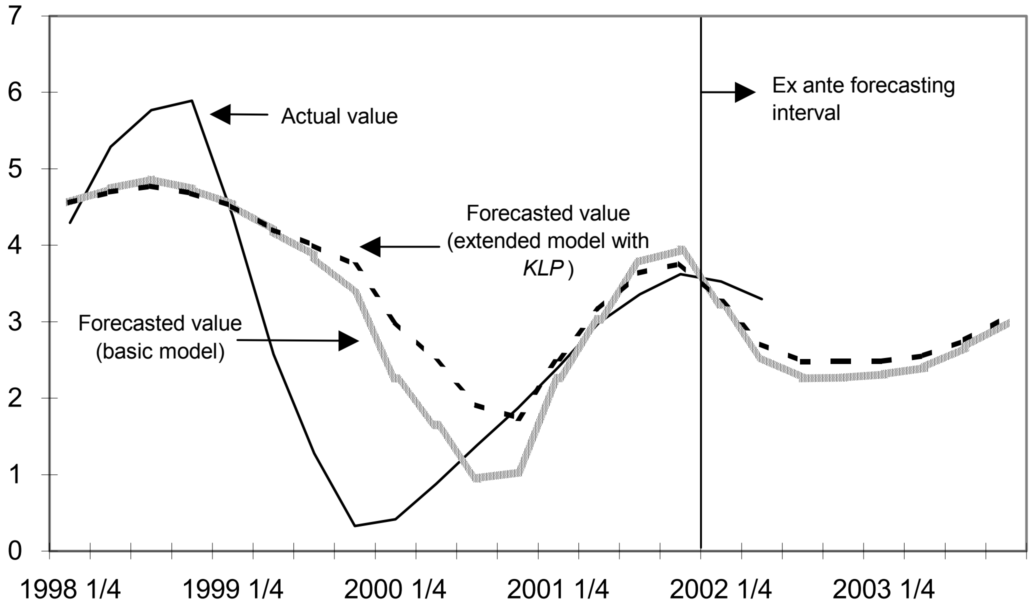


[1] Forecasting horizon is eight quarters.

Graph 10

Actual and forecasted values of land prices (*KLP*)[1]

In percentages



[1] Forecasting horizon is eight quarters.

4. Implications for monetary policy

4.1 Main views in the recent literature

There have been two main views in the literature about how the central bank should respond to asset price movements to achieve sustainable price stability. One of them is that the central bank should respond to asset price fluctuations directly - Kent and Lowe (1997) and Cecchetti et al (2000, 2002). The possible instability of the economy following the financial disturbance caused by asset price fluctuations is one of the reasons for their assertion. In general, the bursting of a bubble after a boom has been sustained for a long period causes a serious imbalance in the balance sheets of financial institutions through insolvencies of lendings backed by collateral. This aggravates the weakness of financial institutions and induces them to reduce or to be more cautious about corporate lending. Therefore, the possible credit crunch could contract investment and, finally, trigger recession. In addition, they argue that the crucial role of the impact of asset price movements on consumption and investment via the monetary transmission mechanism is another reason for monetary policy to respond to asset price movements. Rises and falls in asset prices affect real economic activity mainly through channels such as consumption via the wealth effect and investment through capital gains due to changes in collateral and net asset prices.

However, the prevailing consensus among economists and central bankers is that monetary policy should not directly target asset prices, but should respond to the effects of asset price fluctuations insofar as they signal changes in expected inflation - Bernanke and Gertler (2000) and Batini and Nelson (2000). First, the difficulty of distinguishing asset price movements driven by excess optimism from those led by fundamentals forces the central bank to hesitate about targeting asset prices for monetary policy. For example, to attain the information on a discounted future dividend stream from fluctuation in the stock price index is not straightforward due to the irrational exuberance reflected in stock price index movements. Second, even if the central bank were able to distinguish asset price movements driven by irrational expectations, it does not have adequate policy tools to excise the bubble alone. The obscurity of the relationship between interest rates and asset prices prevents it from responding to asset price fluctuations efficiently. However, the cited authors argue that by focusing on reducing the inflationary or deflationary pressures generated by excess real demand, a central bank can respond effectively to the harmful side effects of asset booms and busts without getting into the business of deciding what is or is not a fundamental.

4.2 Asset price movements and inflation targeting in South Korea

Financial markets in South Korea experienced rapid structural changes in the 1990s owing to a four-stage interest rate deregulation plan, the opening of financial markets, and the liberalisation of the foreign exchange and capital markets. Aligned to the changing financial environment, the provisions of the revised Bank of Korea Act, which came into effect on 1 April 1998, required The Bank of Korea to assume the responsibility for setting an annual target inflation rate, and the conduct of monetary policy in order to attain it (Table 14). Since an inflation targeting regime is a forward-looking pre-emptive framework for monetary policy, which is generally based on a medium-term inflation target, a strong capacity for inflation forecasting ability is required for the success of inflation targeting. From this viewpoint, the results of the analyses in Section 3, indicating that asset price movements are reliable leading indicators for future inflation, imply that the application of information signalled by asset price movements would enhance the inflation forecasting capacity.

For monetary policy to respond to asset price fluctuations effectively, it is essential to find out whether the asset price movements are driven by fundamentals. Regarding this question, the results of research in The Bank of Korea have shown that if the central bank is able to detect bubbles in asset prices, a monetary policy targeting inflation and asset prices is more effective in controlling the business cycle than one with only inflation targeting. However, they suggest that the conduct of a monetary policy that targets inflation and asset prices without detecting the presence of bubbles in asset prices may induce an aggravation of the recession after the bubbles burst.

As shown in Section 2, the disagreement between the results of MRS tests and those of unit-root and cointegration tests about the presence of asset price bubbles shows the difficulty of distinguishing bubbles from fundamentals. In addition, the difficulty of calculating precisely an ex post real asset value for the derivation of the wedge between actual and ex post real asset prices adds another reason for the monetary policy authority to hesitate about responding to asset price movement directly. This implies that further research is necessary for the central bank to investigate the presence of bubbles in asset prices in order to implement an inflation and asset price targeting regime. However, although the central bank cannot perfectly distinguish asset price bubbles from fundamentals, the current inflation targeting regime in South Korea, which responds to the excess real demand pressure on inflation caused by asset price fluctuation, has been to some degree effective in stabilising asset price movements. Furthermore, as shown in Section 3, since it is more evident when asset price movements put pressure on real demand due to the predictive power of asset prices for future inflation, the inflation targeting regime in South Korea is expected to become more effective than in the past.

Table 14

Key points of the sixth revision of the Bank of Korea Act

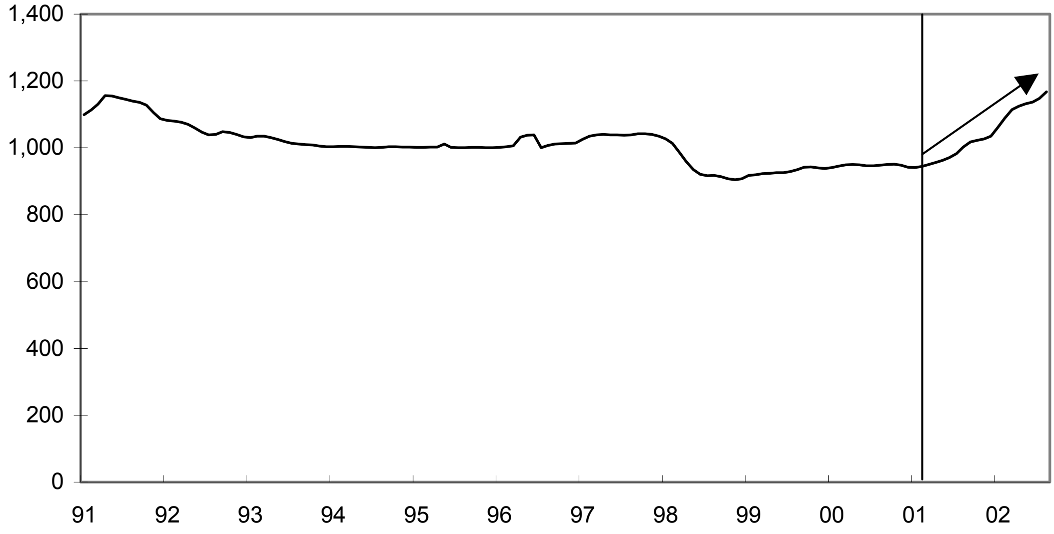
Previous	Revised
Objective - Stabilisation of the value of money and maintenance of the soundness of the banking and credit system	- Price stability
Policymaking body - Minister of Finance and Economy also served as Chairman of the Monetary Board	- Governor of The Bank of Korea (BOK) also serves as the Chairman of the Monetary Policy Committee
Monetary policies - (Added)	- The BOK determines an annual price stability target in consultation with the government and formulates and publishes an operational plan for monetary and credit policies that includes the said target
Relationship with the National Assembly - (Added)	- At least once a year, the BOK shall prepare a report on the implementation of its monetary and credit policies and submit it to the National Assembly
- (Added)	- At the request of the National Assembly or its committees, the Bank Governor should attend and answer the questions of the National Assembly or its committees concerning the report

One of the recent notable features of asset price movements in South Korea is that the housing price index has shown a steep upward trend since early 2001, as shown in Graph 11. The ratio of housing prices to national disposable income per capita (NDI) in South Korea, as may be seen from Graph 12, also indicates that housing prices in South Korea have been increasing steeply, at rates comparable to those of the United States and Spain. Furthermore, the increasing gap between the rate of return on housing prices and yields on corporate bonds in Graph 13 suggests that the upward tendency of housing prices will continue for the time being.

As the excess real demand pressure on inflation caused by increasing housing prices is expected to come to the surface after four or six quarters according to the analysis in Section 3, a response by the monetary policy authority could be considered. However, since an increase in the interest rate causes a rise in the financial costs of consumption and investment and destabilises the financial market, the central bank should be cautious about conducting monetary policy in reaction to asset price movements so as not to unsettle the economy.

Graph 11

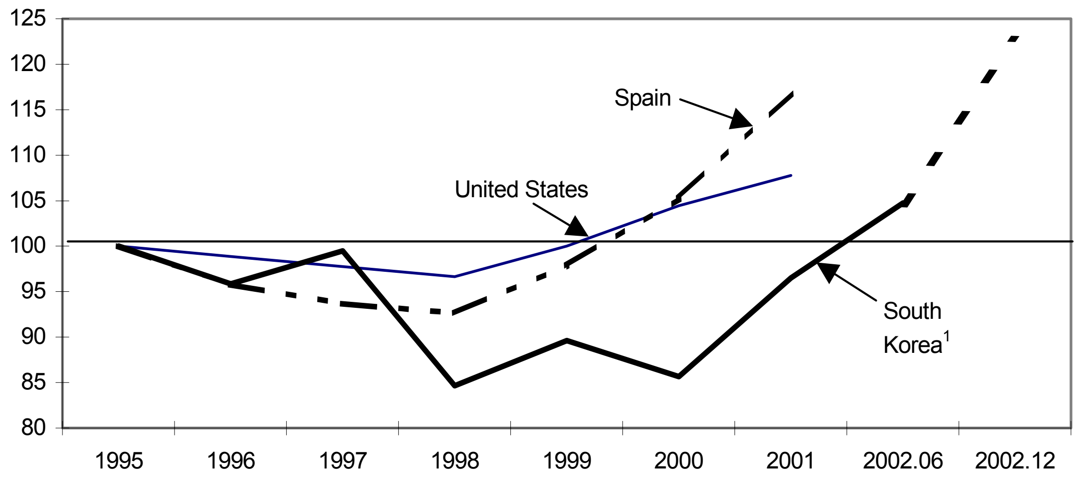
The housing price index in South Korea



Graph 12

Ratio of housing prices to NDI in South Korea and other countries

1995 = 100



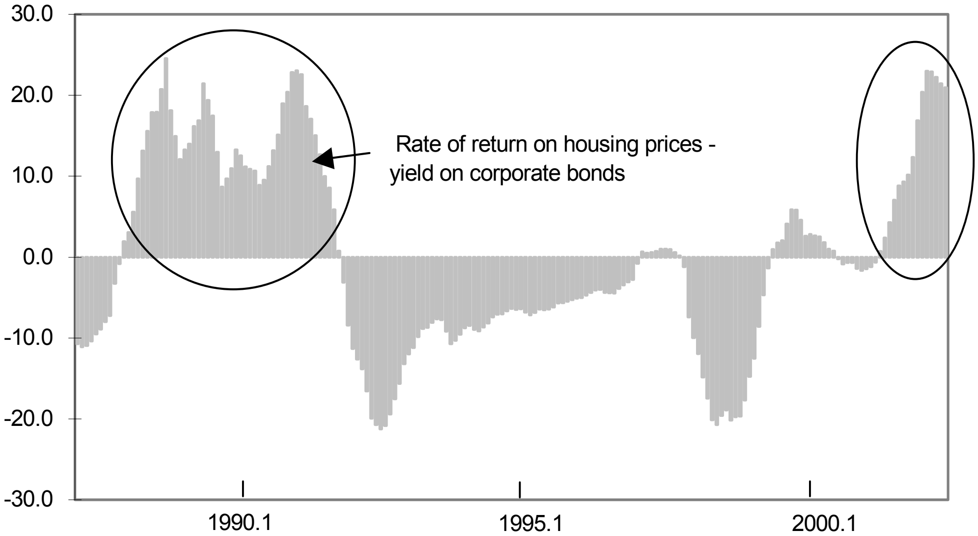
NDI = national disposable income per capita.

[1] Average apartment price in Seoul.

Graph 13

The gap between the rate of return on housing prices and the yield on three-year corporate bonds

In percentage points



5. Concluding remarks

The Asian currency crisis in 1997-98 reminded central banks that the impact of asset price movements on the economy should not be ignored in the conduct of monetary policy. However, without distinguishing bubbles from fundamentals, a direct response to asset price movements from monetary policy could well aggravate the recession after the bubbles burst.

To verify whether there have been bubbles in asset prices in South Korea, trends of asset prices were visually inspected in Section 2. While the graphical information suggests that bubbles could exist from the late 1980s to the early 1990s, empirical tests do not provide us any decisive information about the presence of bubbles. This suggests that distinguishing bubbles from fundamentals is not a straightforward task.

Section 3 checked whether asset price movements have predictive power for future inflation. First, the trends of asset prices, the interest rate spread, and stock and real estate prices show that asset prices lead inflation by four to eight quarters. Second, to confirm this result by empirical analysis, OLS estimations of inflation forecasting models were provided. These estimations show that the interest rate spread has predictive power for future inflation at four-quarter horizons; housing prices provide effective information for future inflation at four- and six-quarter horizons; and land and stock prices could serve as leading indicators at six- and eight-quarter horizons.

Based on the results obtained in the above analyses, the implications of asset price volatility for the conduct of monetary policy were stated in Section 4. While there is little agreement over how monetary policy should react to asset price movements within the literature, it is the mainstream position that central banks should refer to asset price movements only as an information variable for expected inflation. The difficulties of calculating precise ex post real asset values and specifying the period when bubbles exist, as described in Section 2, strongly suggest that a direct response from the central bank to asset price movements is inappropriate.

The inflation targeting regime in South Korea, which responds to excess real demand pressure on inflation by adjusting the interest rate, has been successful in controlling the variability of both the inflation rate and output since it was introduced in 1998. Therefore, even though the central bank may not be able to pinpoint the causes of asset price movements, the conduct of monetary policy that seeks to reduce the excess real demand pressure on inflation caused by asset price movements appears promising. However, research to find the causes of asset price movements and specify the periods when bubbles exist should be continued in order to raise the effectiveness of monetary policy in view of the crucial role of asset price movements in the business cycle.

Appendix

Let y_t be an $(n \times 1)$ vector partitioned as

$$\underset{(n\times 1)}{y_t} = \begin{bmatrix} \underset{(1\times 1)}{y_{1t}} \\ \underset{(g\times 1)}{y_{2t}} \end{bmatrix} \tag{1}$$

for $g \equiv (n-1)$. Consider the regression

$$y_{1t} = \alpha + \gamma' y_{2t} + u_t \tag{2}$$

Let $\hat{u}_t$ be the sample residual associated with OLS estimation of (2) in a sample of size T:

$$\hat{u}_t = y_{1t} - \hat{\alpha}_T - \hat{\gamma}'_T y_{2t} \text{ for } t = 1, 2, ...T \tag{3}$$

The residual $\hat{u}_t$ can then be regressed on its own lagged value $\hat{u}_{t-1}$ without a constant term:

$$\hat{u}_t = \rho \hat{u}_{t-1} + e_t \text{ for } t = 2, 3, ..., T \tag{4}$$

yielding the estimate

$$\hat{\rho}_T = \frac{\sum_{t=2}^{T} \hat{u}_{t-1}\hat{u}_t}{\sum_{t=2}^{T} \hat{u}_{t-1}^2} \tag{5}$$

Let s_T^2 be the OLS estimate of the variance of e_t for the regression of equation (4):

$$s_T^2 = (T-2)^{-1} \sum_{t=2}^{T} (\hat{u}_t - \hat{\rho}_t \hat{u}_{t-1})^2 \tag{6}$$

and let

$$\hat{\sigma}_{\hat{\rho}_T} = s_T^2 \Big/ \left\{ \sum_{t=2}^{T} \hat{u}_{t-1}^2 \right\} \tag{7}$$

Finally, let $\hat{a}_{j,T}$ be the jth sample autocovariance of the estimated residuals associated with equation (4):

$$\hat{a}_{j,T} = (T-1)^{-1} \sum_{t=j+2}^{T} \hat{e}_t \hat{e}_{t-j} \text{ for } j = 0, 1, 2, ..., T-2 \tag{8}$$

for $\hat{e}_t \equiv \hat{u}_t - \hat{\rho}_T \hat{u}_{t-1}$; and let the square of $\hat{\lambda}_T$ be given by

$$\hat{\lambda}_T^2 = \hat{a}_{0,T} + 2 \sum_{j=1}^{q} [1 - j/(q+1)] \hat{a}_{j,T} \tag{9}$$

where q is the number of autocovariance to be used. Phillips's Z_ρ statistic (1987) can be derived as follows:

$$Z_{\rho,T} = (T-1)(\hat{\rho}_T - 1) - (1/2)\{(T-1)^2 \hat{\sigma}_{\hat{\rho}_T}^2 \div s_T^2\}\{\hat{\lambda}_T^2 - \hat{a}_{0,T}\} \tag{10}$$

Similarly, Phillips's Z_t statistic associated with the residual autoregression equation (4) would be

$$Z_{t,T} = (\hat{a}_{0,T} / \hat{\lambda}_T^2)^{1/2} t_T - (1/2)\{(T-1)\hat{\sigma}_{\hat{\rho}_T} \div s_T\}\{\hat{\lambda}_T^2 - \hat{a}_{0,T}\} / \hat{\lambda}_T \tag{11}$$

and for t_T, the usual OLS t-statistic for testing the hypothesis $\rho = 1$:

$$t_T = (\hat{\rho}_T - 1) / \hat{\sigma}_{\hat{\rho}_T} \tag{12}$$

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