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The Impact of Monetary Tightening on the Won/Dollar Exchange Rate*

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This study examines the impact of the widening interest rate differential between Korea and the U.S. on the won/dollar exchange rate. Reflecting characteristics of the Korean economy such as its small open economy, currency and financial crises, and structural economic changes, I estimate VAR models and ECMs, treating the dollar index and oil prices as exogenous variables alongside dummy variables. The impulse response analysis reveals that during the pre-currency crisis period (April 1990–October 1997), an upward shock to (call rate - federal funds rate) caused industrial production and consumer prices to decline while increasing the won/dollar exchange rate. From the post-currency crisis period to the pre-global financial crisis period (January 1999–August 2008), a price puzzle phenomenon emerged during monetary tightening shocks, and the won/dollar exchange rate exhibited an inverse U-shaped response. However, from the period when Korea became a net external financial asset holder until recently (January 2014–May 2025), monetary tightening shocks have caused both prices and the won/dollar exchange rate to decline. Because the effects of increased liquidity demand and deflation outweigh the effect of reduced industrial production, a so-called "advanced-economy-type forward premium puzzle" emerges—in which a monetary tightening shock causes the domestic currency to appreciate—contrary to the claims of existing international studies.

Keywords: Interest Rate Difference, Forward Premium Puzzle, Price Puzzle, Block Exogeneity

JEL Classification: C13, E44, F31

I. Introduction

As a small open economy, Korea's domestic economic conditions are significantly influenced above all by the global economic situation. Changes in the global economic environment directly impact the domestic economy through the current account balance and the won/dollar exchange rate. Among these, the current account balance has

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consistently posted surpluses since the 1997 currency crisis until recently, making a significant contribution to domestic economic stability. However, despite the won/dollar exchange rate's volatility gradually decreasing due to the current account surplus sustained since the currency crisis, it has repeatedly surged dramatically during periods like the global financial crisis and the COVID-19 pandemic. Therefore, research examining the causal relationship between domestic and international macroeconomic variables and the won/dollar exchange rate is a challenging task.

According to traditional exchange rate determination theory, exchange rates are determined not only by external factors but also by the income, prices, interest rates, expected inflation, and current account balances of the two countries, as exchange rates represent the ratio of the two countries' currencies. In particular, the representative macroeconomic theories—the Chicago School's flexible price exchange rate determination theory and the Keynesian sticky price exchange rate determination theory—focus on the relationship between the interest rate differential between the two countries and the exchange rate. The former argues that an increase in the domestic interest rate relative to the foreign interest rate raises expected inflation, thereby decreasing the value of the domestic currency, while the latter presents the opposing view that it increases demand for the domestic currency, thereby raising its value.

When covered interest rate parity holds, the interest rate differential between two countries equals the forward premium. Under the assumptions of risk neutrality and rational expectations, if the foreign exchange market is efficient, the expected future spot rate equals the forward rate. If they differ, this indicates either an inefficient foreign exchange market or the existence of a risk premium. According to Fama (1984), when regressing the expected future exchange rate change rate against the forward premium, the regression coefficient should equal 1 if the foreign exchange market is efficient. However, empirical analysis results show a negative value. Existing studies refer to this phenomenon as the "forward premium puzzle or bias."

Since Eichenbaum and Evans (1995) analyzed the impact of U.S. monetary tightening shocks on the U.S. nominal and real exchange rates using a VAR model, numerous studies have examined the effect of interest rate differentials between the two countries on exchange rates using various VAR models. In particular, Hnatkovska et al. (2016) and Kalemli-Özean and Varela (2025) demonstrated, using panel VAR models and local projections, respectively, that positive interest rate differential shocks appreciated exchange rates in advanced countries while depreciating those in developing countries. While these studies included Korea among developing countries to reach such conclusions, their empirical findings on the won/dollar exchange rate warranted reconsideration on several points. First, despite including the periods of the currency crisis and the global financial crisis in their analysis, they did not adequately account for these events in their estimations. In addition, despite the won/dollar exchange rate being significantly influenced by major global variables due to Korea's status as a small open economy, these variables were not adequately incorporated into the analytical model. Moreover, given the structural changes that occurred in the Korean economy during their analysis period or thereafter, their conclusions could not be considered valid anymore.

This study aims to examine whether an "advanced-economy-type forward premium" exists by analyzing the impact of tight monetary policy, measured by the interest rate differential between Korea and the United States, on the won/dollar exchange rate using an analytical model that addresses the limitations of previous studies. The analysis period spans from March 1990, when the market-average exchange rate system was introduced, to May 2025. Given the occurrence of multiple crises and changes in economic structure during this time, the analysis covers not only the entire period (April 1990–May 2025) but also subdivides it into the pre-currency crisis period (April 1990–October 1997), the post-currency crisis to pre-global financial crisis period (January 1999–August 2008), and the period from when Korea became a net external financial asset holder to the present (January 2014–May 2025). After selecting the appropriate level or difference VAR model for each period based on AICc criteria and cointegration tests, I conduct impulse response analysis. To reflect the influence of global variables, I include the error terms of the dollar index and international oil prices, derived from an ARMA(1,1) model, as exogenous variables in the VAR model or error correction model (ECM). Additionally, I estimate a VAR model based on Lastrapes (2005, 2006), which incorporates global variables as endogenous variables but assumes domestic shocks do not affect global variables. This analysis focuses on how the so-called "advanced-economy forward premium puzzle"—in which tight monetary policy causes the domestic currency to appreciate—manifests itself over different time periods.

Impulse response analysis results indicate that during the period before the currency crisis (April 1990–October 1997), an upward shock in (call rate – federal funds rate) did not produce a price puzzle phenomenon, and the won/dollar exchange rate increased. From the post-currency crisis period to the pre-global financial crisis period

(January 1999–August 2008), unlike the pre-currency crisis period, a price puzzle phenomenon emerged during interest rate hikes, and the won/dollar exchange rate exhibited an inverse U-shaped response, consistent with the overall pattern. From the period when Korea became a net external financial asset country until recently (January 2014–May 2025), not only does the price puzzle phenomenon not occur, but the won/dollar exchange rate declines. Contrary to claims in existing foreign studies, a "forward premium puzzle" phenomenon similar to that seen in advanced economies has recently emerged.

The contents of this study are organized as follows. Section II reviews existing studies that theoretically and empirically analyze the effects of a widening interest rate differential on exchange rates in relation to the "forward premium puzzle", and discusses the differences between this study and those studies. Section III conducts unit root tests and cointegration tests and examines basic statistics. Section IV estimates a traditional OLS model concerning the interest rate differential. Sections V and VI estimate various VAR models and ECMs including global variables as exogenous shock variables and analyze the impulse response functions. Section VII discusses the implications of this study regarding the causal relationship between interest rate differentials and the won/dollar exchange rate. Finally, Section VIII summarizes the overall research findings and concludes.

II. Literature Review

The monetary approach to exchange rate determination is broadly divided into two opposing schools of thought: the Chicago School theory and the Keynesian theory. The Chicago School theory argues that, under the assumption of perfectly flexible prices, changes in the nominal interest rate differential reflect changes in expected inflation, thus establishing a positive relationship between the nominal interest rate differential and the exchange rate (e.g., Bilson, 1978). Conversely, the Keynesian theory posits that, assuming prices are sticky in the short run, changes in the nominal interest rate differential reflect shifts in contractionary monetary policy, thus establishing a negative relationship between the nominal interest rate differential and the exchange rate (e.g., Dornbusch, 1976). Meanwhile, the real interest rate differential theory of Frankel (1979) combines the Keynesian assumption of sticky prices with the Chicago School's assumption of persistent inflation. It states that the exchange rate has a negative relationship with the nominal interest rate differential but a positive relationship with the long-term expected inflation differential.

Eichenbaum and Evans (1995) demonstrated through VAR models that, contrary to previous studies, U.S. monetary tightening shocks not only caused a significant and persistent appreciation of the U.S. nominal and real exchange rates but also led to a persistent and substantial deviation from uncovered interest parity. This empirical result supported a monetary business cycle model that allowed for liquidity effects rather than a money-neutral real business cycle model and aligned with the Dornbusch (1976) and Frankel (1979) models, which emphasized nominal rigidity. However, the consistent appreciation of the dollar was inconsistent with the overshooting model of Dornbusch (1976) or the uncovered interest rate parity assumption.

The empirical results of Eichenbaum and Evans (1995) are closely related to previous studies on the "forward premium puzzle". When the covered interest rate parity holds, the interest rate differential between two countries is identical to the forward premium. If the uncovered interest rate parity holds, then when regressing the expected exchange rate change rate against the forward premium, the constant term should be zero and the regression coefficient should be one. However, according to Fama (1984), the regression coefficient takes on a negative value. Thus, the "forward premium puzzle" implies that the expected exchange rate change rate and the forward premium have a negative relationship. The prevailing view, including Lustig et al. (2011) and Hassan and Mano (2019), attributes this negative relationship between expected exchange rate changes and the forward premium to the existence of a risk premium. Conversely, some view this relationship as stemming from market participants' bias in expectations or information constraints (e.g., Froot and Frankel, 1989; Candian and de Leo, 2025).

While most studies on the "forward premium puzzle" have focused on advanced economies, Bansal and Dahlquist (2000) argued, through an empirical analysis using data from 28 countries including both advanced and emerging market economies, that the inclusion of emerging market economies played a decisive role in establishing a positive relationship between interest rate differentials and exchange rates. The "forward premium puzzle" is seen as a phenomenon confined to advanced economies. Frankel and Poonawala (2010) also demonstrated, using expanded data from both advanced and emerging market economies, that the Fama coefficient was small but positive.

Hnatkovska et al. (2016) demonstrated through a panel VAR model using data from 25 advanced economies and 47 emerging market economies that tight monetary policy appreciated the exchange rates of advanced economies while depreciating those of emerging market economies. They explained this differential response using a calibrated small open economy model with three monetary policy channels: the liquidity demand channel, the fiscal channel, and the output channel. The liquidity demand channel increases demand for domestic currency-denominated liquid assets during monetary tightening, thereby appreciating the domestic currency. Conversely, when interest rates rise, the output channel and fiscal channel each depress domestic economic activity or increase the fiscal burden, thereby depreciating the domestic currency. The strength of the liquidity effect is an increasing function of the money/GDP ratio and the deposit/cash ratio, which are systematically higher in advanced economies. Therefore, in advanced economies, the liquidity demand channel is more important than the indirect output and fiscal channels, so that tight monetary policy causes the exchange rate to fall. In emerging market economies with low money/GDP and deposit/cash ratios, the indirect output and fiscal effects outweigh the liquidity effect, so that the exchange rate falls during monetary tightening.

While previous studies regressed actual future exchange rate change rates on the interest rate differential or forward premium between two countries, Kalemli-Özean and Varela (2025) used expected exchange rate change rates derived from survey data. Using panel data from 22 emerging market economies and 12 advanced economies, their OLS and local projections estimates showed that emerging market economy exchange rates rose during monetary tightening, even when using survey exchange rate data, just as they did when using actual exchange rate data.

This study differs from existing foreign research by analyzing how the so-called "forward premium puzzle" phenomenon has changed over specific sub-periods during the past 35 years since the market-average exchange rate system was implemented in March 1990. Representative recent studies using the won/dollar exchange rate include Hnatkovska et al. (2016) and Kalemli-Özean and Varela (2025). The former used actual won/dollar exchange rate data from December 1997 to December 2007, while the latter utilized survey data on expected won/dollar exchange rates from November 1996 to December 2018. Both papers classified Korea as an emerging market economy in their analysis. Hnatkovska et al. (2016) demonstrated through impulse response analysis of bivariate and multivariate VAR models that Korea's monetary tightening policies, while not statistically significant, did cause the won/dollar exchange rate to rise. Kalemli-Özean and Varela (2025) included Korea in the developing country panel, showing that developing country exchange rates rose when their interest rates increased.

However, not only is the analysis period too short, but events such as the currency crisis, the global financial crisis, and structural changes in the Korean economy occurred during this time; therefore, it is difficult to place full confidence in the findings of these studies regarding the won-dollar exchange rate. The extreme values of the Korea-U.S. interest rate differential and the won/dollar exchange rate, as well as regime switching, are highly likely to distort the estimation results of simple linear analysis. Furthermore, as Korea is a small open economy, its won/dollar exchange rate is highly susceptible to significant influence from uncontrollable global factors. Therefore, this study aims to examine how the "forward premium puzzle" phenomenon changes by incorporating these factors into the estimation model.

III. Data Characteristics

This study utilizes monthly data from March 1990, when the market-average exchange rate system was introduced, through May 2025. However, since both level variables and difference variables are used, the actual analysis period is April 1990 to May 2025, with a sample size of 422. Given that the analysis period encompasses the currency crisis, the global financial crisis, and changes in economic structure, the data is segmented not only into the entire period (April 1990–May 2025) but also into the pre-currency crisis period (April 1990–October 1997), the post-currency crisis period until the pre-global financial crisis (January 1999–August 2008), and the period from when Korea became a net external financial asset holder until the present (January 2014–May 2025).[1]

The interest rate differential between Korea and the U.S. (call rate minus federal funds rate), the won/dollar exchange rate (15:30 closing rate and daily average rate), and domestic and international control variables such as industrial production (*IP*), consumer price index (*CPI*), US dollar index (*$index*), and international oil price (*WTI*) are used. The won/dollar exchange rate is sourced from the Bank of Korea, while the call rate, industrial production, and consumer prices are sourced from the IMF. The dollar index and WTI oil prices are sourced from the Federal Reserve Board. The broad

[1] In Appendix A, I examine the Qu and Perron (2007) test to verify at which point in time structural breaks occurred during the analysis period. For the results of the empirical analysis covering the period immediately following the global financial crisis, see Appendix D.

dollar index is used as the dollar index.[2] Level variables, excluding interest rates, are used by taking the logarithm. Difference variables are expressed as a percentage by taking the logarithm, differencing, and multiplying by 100 ($\Delta x = 100 \times \Delta \log x$). The difference between the two interest rates is expressed in percentage points (%p). Before proceeding with the main analysis, unit root tests, cointegration tests, and basic statistics are examined.

1. Unit Root and Cointegration Tests

The VAR model is employed as the primary analytical method in this study. Criteria for selecting the lag number in VAR models include AIC, AICc, and SIC. This study adopts the lag number determined by AICc criterion (Jorda, 2005), which lies between AIC and SIC. According to the AICc criterion, the number of lags is determined as follows: 4 for the period '90.04–'25.05, 1 for '90.04–'97.10, 2 for '99.01–'08.08 and '14.01–'25.05, resulting in 4, 1, 2, and 2, respectively. To conserve space, only the test results corresponding to these lag lengths are described.

Table 1 shows the results of the ADF unit root test. For each level variable, the ADF unit root test results indicate that during the period '90.04–'25.05 with a lag of 4, *CPI* and (*call-ffr*), and during the period '14.01–'25.05 with a lag of 2, *$index*, among others, do not have a unit root at the 10% significance level. For the difference variables, none of them are found to have a unit root regardless of the period or variable.

Table 2 shows the cointegration test results for 4-variable and 6-variable models with the same lag number as Table 1. In Table 2, H_0: r=0 or r=1 (3, 5) indicates the null hypothesis that no cointegration vector exists or that one (3, 5) exists, respectively. First, examining the case of the 4-variable model composed solely of domestic variables [*IP*, *CPI*, (*call-ffr*), *won/$*], I observe that the cointegration relationship varies across periods. For the period '90.04–'25.05, the Trace statistic for the null hypothesis H_0: r=3 is 4.770, which is greater than the 5% critical value of 3.842, leading to the rejection of this null hypothesis. Conversely, for the periods '90.04–'97.10 and '99.01–'08.08, the null hypothesis (H_0: r=0) that no cointegration relationship

[2] Effective February 4, 2019, the dollar index was restructured into the broad dollar index and two supplementary dollar indices—the AFE (advanced foreign economies) dollar index and the EME (emerging market economies) dollar index.

exists is accepted. However, for the period ’14.01–’25.05, the null hypothesis that no cointegration relationship exists is rejected. For the 6-variable model including the global variable *$index* and *WTI*, the test results are identical to those for the 4-variable model, except that the null hypothesis of one cointegration relationship existing during

Table 1. ADF Unit Root Test

Period	Lag		*IP*	*CPI*	*call-ffr*	*won/$*	*$index*	*WTI*
’90.04–’25.05	4	Level	−2.161	−4.853**	−2.649+	−2.357	−2.418	−1.601
		Difference	−9.562**	−8.266**	−8.401**	−8.565**	−8.361**	−9.926**
’90.04–’97.10	1	Level	−0.742	−1.592	−2.517	0.641	−0.260	−3.915**
		Difference	−9.860**	−6.772**	−7.935**	−4.064**	−5.528**	−6.303**
’99.01–’08.08	2	Level	−1.561	0.581	−1.197	−1.290	0.426	−0.689
		Difference	−6.212**	−8.165**	−4.398**	−5.024**	−5.693**	−6.272**
’14.01–’25.05	2	Level	−1.021	1.909	−1.422	−1.115	−2.730*	−2.569
		Difference	−7.136**	−7.917**	−4.735**	−7.727**	−7.043**	−7.560**

Notes: 1) *IP*, *CPI*, and (*call-ffr*) denote industrial production, consumer price index, and call rate–federal funds rate, respectively; all variables except (*call-ffr*) are logarithmically transformed.

2) (*call-ffr*) is expressed as the difference (%p) between the two levels.

3) + and ** indicate significance at the 10% and 1% levels, respectively.

Table 2. Cointegration Tests

Variable	Period	H_0	Lag	Trace Statistics	Critical Value (5%)
IP, *CPI*, (*call-ffr*), *won/$*	’90.04–’25.05	r=3	4	4.770	3.842
	’90.04–’97.10	r=0	1	25.119	47.856
	’99.01–’08.08	r=0	2	45.675	47.856
	’14.01–’25.05	r=0	2	53.122	47.856
IP, *CPI*, (*call-ffr*), *won/$*, *$index*, *WTI*	’90.04–’25.05	r=5	4	4.416	3.842
	’90.04–’97.10	r=0	1	90.915	95.754
	’99.01–’08.08	r=0	2	95.264	95.754
	’14.01–’25.05	r=1	2	77.583	69.819

Notes: 1) *IP*, *CPI*, and (*call-ffr*) denote industrial production, consumer price index, and call rate-federal funds rate, respectively; all variables except (*call-ffr*) are logarithmically transformed.

2) (*call-ffr*) is expressed as the difference (%p) between the two levels.

3) H_0: r=0 and r=1 (3, 5) represents the null hypothesis that there is no cointegrating vector and that there is 1 (3, 5) cointegrating vector, respectively.

the period ’14.01–’25.05 is rejected. Therefore, the level VAR model is used for estimation and impulse response analysis for the period ’90.04–’25.05, the difference VAR model for the periods ’90.04–’97.10 and ’99.01–’08.08, and either the difference VAR model or the ECM is used for the period ’14.01–’25.05.

2. Sample Statistics

Table 3 shows the sample statistics for the rate of change (%). For the *IP* rate of change, it increased by an average of 0.412% over the entire analysis period (April 1990–May 2025). However, examining by period reveals that the growth rate decreased to 0.108% between January 2014 and May 2025 from 0.629% between April 1990 and October 1997, a reduction to one-sixth. The average growth rates for *CPI* and (*call-ffr*) (%p) also show a continuous decline over time. The average change rate of *won/$* was negative during the period ’99.01–’08.08, but it is not statistically significantly different from zero for the entire period or for each sub-period. The average change rate of *$index*, like *won/$*, was negative during the period ’99.01–’08.08, but it has statistically significant positive values in other periods. The average change rate of *WTI* was positive throughout the entire period but turned negative during the period from January 2014 to May 2025. For the entire period, the absolute values of the maximum and minimum of all variables were very large. However, for the sub-periods excluding the currency crisis period and the global financial crisis period, the absolute values decreased as extreme values were excluded. Q(10) represents the Ljung-Box test statistic for the 10th-order serial correlation. For the entire period, the influence of extreme values indicates a significantly higher serial correlation compared to the sub-periods.

Figure 1 shows the trends and rates of change in key variables. Since these variables changed significantly during the currency crisis period and the global financial crisis period, it demonstrates that simply estimating a linear model that includes these periods can lead to distortion of the estimation results due to the influence of extreme values. Therefore, this study examines periods excluding these periods in detail, in addition to the entire period. The won/$ exchange rate shows movements similar to *$index* when excluding the currency crisis period. *WTI* rose sharply starting in 2020 due to the COVID-19 pandemic and the Russia-Ukraine war, but has been declining again since 2022.

Table 3. Sample Statistics for Change Rate (%)

Period		*IP*	*CPI*	*call-ffr*	*won/$*	*$index*	*WTI*
’90.04–’25.05	Mean	0.412**	0.268**	2.812**	0.163	0.146**	0.264
	STD	2.263	0.428	3.956	2.903	1.267	9.561
	Max	9.265	2.498	20.070	37.236	7.161	52.799
	Min	–11.546	–0.747	–1.866	–8.857	–3.338	–54.654
	Q(10)	29.241 [0.001]	105.772 [0.000]	3035.93 [0.000]	102.857 [0.000]	75.869 [0.000]	47.102 [0.000]
’90.04–’97.10	Mean	0.629**	0.467**	8.453**	0.310	0.413**	0.043
	STD	1.783	0.432	2.311	0.712	1.144	7.873
	Max	4.047	2.121	14.890	2.304	3.723	37.617
	Min	–5.453	–0.512	4.280	–1.709	–2.778	–18.892
	Q(10)	13.491 [0.198]	21.912 [0.016]	261.660 [0.000]	58.488 [0.000]	36.661 [0.000]	41.075 [0.000]
’99.01–’08.08	Mean	0.669**	0.256**	0.854**	–0.126**	–0.133	2.015**
	STD	2.116	0.415	1.521	1.858	1.039	7.744
	Max	7.058	1.289	3.210	5.089	2.841	19.994
	Min	–4.950	–0.604	–1.390	–4.976	–3.338	–19.658
	Q(10)	15.388 [0.119]	60.614 [0.000]	751.025 [0.000]	7.884 [0.640]	12.319 [0.264]	6.973 [0.728]
’14.01–’25.05	Mean	0.108	0.161**	0.129	0.201	0.198+	–0.331
	STD	2.193	0.332	1.188	1.846	1.235	11.705
	Max	6.318	1.057	2.430	5.607	3.634	52.799
	Min	–7.861	–0.747	–1.866	–4.884	–2.426	–54.654
	Q(10)	46.385 [0.000]	54.213 [0.000]	959.117 [0.000]	35.815 [0.000]	38.957 [0.000]	23.188 [0.010]

Notes: 1) *IP*, *CPI*, and *(call-ffr)* denote industrial production, consumer price index, and call rate–federal funds rate, respectively; all variables except (*call-ffr*) are logarithmically transformed.

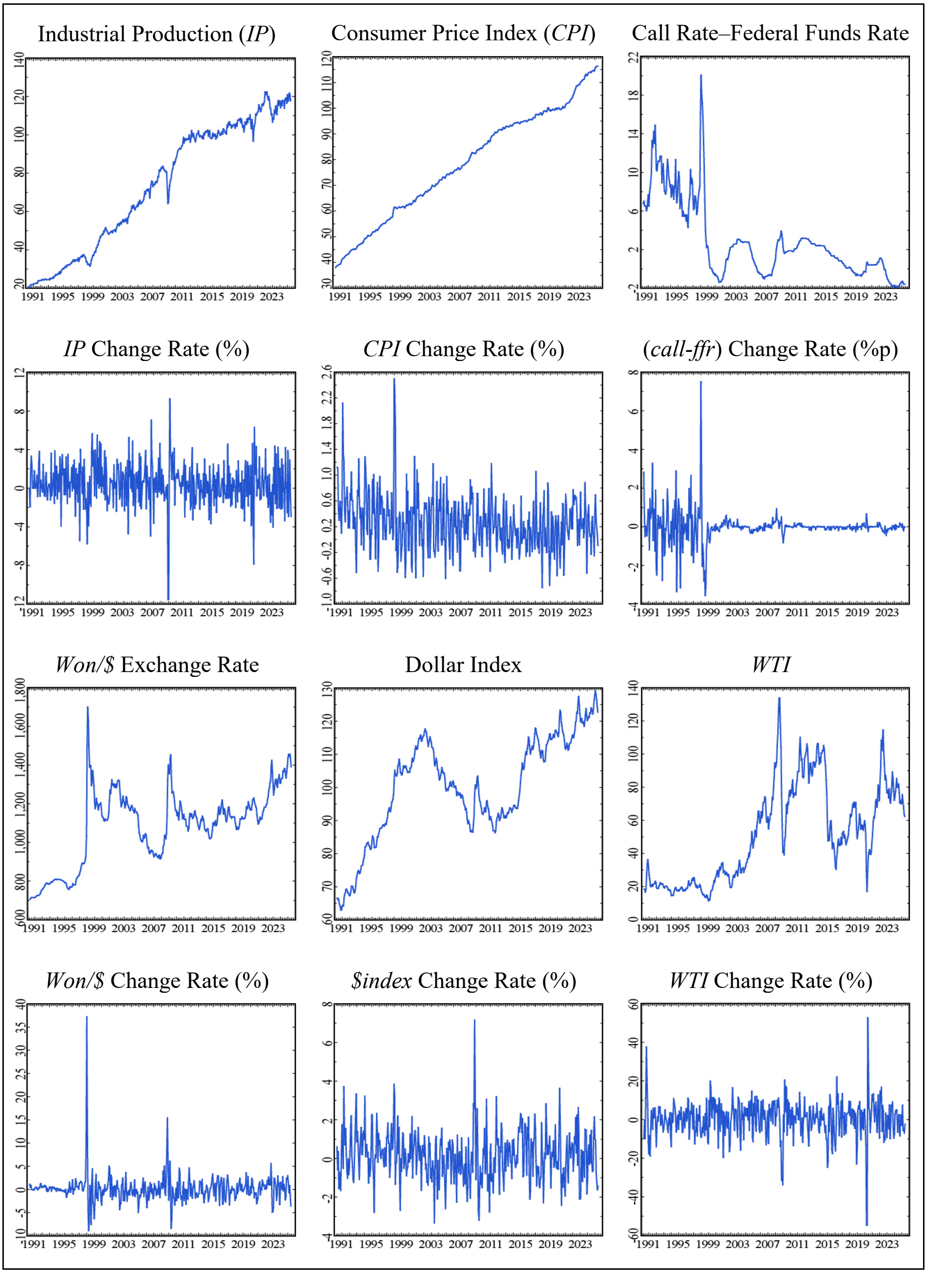
2) (*call-ffr*) is expressed as the difference (%p) between the two levels.

3) $^{+}$, * and ** indicate significance at the 10%, 5%, and 1% levels, respectively.

4) Q(10) represents the Ljung-Box test statistic for the 10th-order serial correlation.

5) Values in [ ] indicate p-values.

Figure 1. Trends and Rates of Change in Key Variables



IV. Estimated Results

As mentioned earlier, Fama (1984) estimated the following equation to examine whether the forward premium could explain the rate of change in future spot exchange rates.

$$s_{t+1} - s_t = \beta_0 + \beta_1 (f_t - s_t) + \varepsilon_{t+1} \tag{1}$$

where s_{t+1} and f_t denote the won/dollar spot exchange rate and the won/dollar forward exchange rate, respectively, both taken as logarithms. Under the assumptions of risk neutrality and rational expectations, β_1 should equal 1. However, Fama (1984) argued that empirical analysis results showed a negative value for β_1, which was due to the negative correlation between $E_t(s_{t+1} - s_t)$ and the premium, the two components of $(f_t - s_t)$.

Under the covered interest parity assumption, Equation (1) is equivalent to the following equation.

$$s_{t+1} - s_t = \beta_0 + \beta_1 (i_t - i_t^*) + \varepsilon_{t+1} \tag{2}$$

where i_t and i_t^* represent the interest rates of Korea and the United States, respectively, using the call rate and the federal funds rate. As observed in previous studies, β_1 takes a negative value in advanced economies, whereas it takes a positive value in emerging market economies. Notably, Hnatkovska et al. (2016) and Kalemli-Özean and Varela (2025) included Korea in the emerging market economy panel, demonstrating that emerging market economy exchange rates depreciated when their interest rate differentials rose. If β_1 is positive, the more the call rate exceeds the federal funds rate, the more the won depreciates over the holding period. When Korea implements tight monetary policy, the won/dollar exchange rate tends to rise rather than fall.

In this study, I aim to estimate the following equation by extending Equation (2).

$$\begin{aligned}\Delta s_t = \beta_0 + \beta_1(i_{t-1} - i_{t-1}^*) + \beta_2 \Delta dindex_t + \beta_3 \Delta s_{t-1} + \beta_{D1} D_{1,t} \\ + \beta_{D2} D_{2,t} + \epsilon_t\end{aligned} \tag{3}$$

where Δs_t and $\Delta dindex_t$ denote the percentage change in the won/dollar exchange rate and the percentage change in the dollar index at time *t*, respectively. D_1 is a currency crisis dummy variable equal to 1 for the period from November 1997 to December 1998 and 0 for all other periods. Similarly, D_2 is a dummy variable for the global financial crisis period, equal to 1 for the period from September 2008 to June 2009 and 0 for all other periods.

The estimation results for Equation (3) are shown in Table 4. Examining the entire period (April 1990–May 2025), when considering additional dummy variables for the

Table 4. OLS Estimates

Period	β_0	β_1	β_2	β_3	β_{D1}	β_{D2}	R^2
'90.04–'25.05	0.039 (0.175)	0.007 (0.038)			1.809 (0.840)*	1.812 (0.924)*	0.022
	0.017 (0.146)	–0.042 (0.032)	1.258 (0.094)**		1.685 (0.703)*	0.996 (0.775)	0.317
	0.059 (0.136)	–0.056 (0.030)**	1.091 (0.090)**	0.314 (0.040)**	1.091 (0.661)$^{+}$	0.462 (0.727)	0.408
'90.04–'97.10	0.197 (0.286)	0.012 (0.033)					0.002
	0.092 (0.282)	0.017 (0.032)	0.152 (0.064)*				0.062
	0.056 (0.253)	0.006 (0.029)	0.152 (0.058)**	0.449 (0.095)**			0.255
'99.01–'08.08	–0.067 (0.198)	–0.070 (0.115)					0.003
	–0.015 (0.181)	–0.011 (0.105)	0.763 (0.153)**				0.183
	0.018 (0.174)	0.015 (0.101)	0.784 (0.147)**	0.254 (0.095)**			0.252
'14.01–'25.05	0.197 (0.160)	0.022 (0.133)					0.000
	0.011 (0.115)	–0.130 (0.096)	1.062 (0.093)**				0.496
	0.001 (0.115)	–0.122 (0.096)	1.017 (0.100)**	0.080 (0.067)			0.501

Notes: 1) D_1 is a dummy variable for the currency crisis period, which equals 1 from November 1997 to December 1998 and 0 for other periods. Similarly, D_2 is a dummy variable for the global financial crisis period, which equals 1 from September 2008 to June 2009 and 0 for other periods.

2) $^{+}$, * and ** indicate significance at the 10%, 5%, and 1% levels, respectively.

currency crisis and the global financial crisis in Equation (2), the estimated value of β_1 is 0.007, nearly zero, while the estimates of the both dummy variables are statistically significant. When considering *$index* additionally as a proxy variable for risk premium, the estimate of β_1 changes to a negative value. Even when the change rate of *won/$* with a one-period lag is added as a partial adjustment term, the estimate for β_1 also takes a negative value, and the parameter estimates for *$index* or the partial adjustment term also exhibit high statistical significance. The estimate of β_1 is positive during the pre-currency crisis period (April 1990–October 1997), but negative during the post-currency crisis to pre-global financial crisis period (January 1999–August 2008), though not statistically significant. Meanwhile, the estimates of β_2 and β_3 have high statistical significance in both periods, as shown by the t-value and R^2. In the period after Korea became a net external financial asset holder and net external creditor nation (January 2014–May 2025), when *$index* is considered an explanatory variable, the estimate of β_1 has an absolutely larger negative value compared to previous periods. Unlike the period before the currency crisis, during the period after the global financial crisis, when the interest rate differential between Korea and the US widens due to Korea's tight monetary policy, the value of the won does not fall but rises. Rather than offsetting profits linked to investment in Korea, expectations of a stronger won exchange rate further increase these profits.

V. Estimation Results of the VAR Model

1. Estimation Model

This study considers a four-variable structural VAR model, similar to recent studies.

$$\Gamma_0 Z_t = a_0 + \sum_{i=1}^{I} a_{Di}\, D_{i,t} + \sum_{j=1}^{J} a_{Sj}\, \eta_{j,t} + \sum_{p=1}^{P} A_p Z_{t-p} + u_t \tag{4}$$

In Equation (4), $Z_t = [IP_t, CPI_t, (\mathit{call\text{-}ffr})_t, won/\$_t]'$ is used for VARs with level variables, $Z_t = [\Delta IP_t, \Delta CPI_t, (\mathit{call\text{-}ffr})_t, \Delta won/\$_t]'$ for VARs with difference variables, and Γ_0 indicates the contemporaneous causality. Parameters Γ_0 and A_p are 4×4 matrices. IP_t, CPI_t, and $won/\$_t$ denote the logarithms of industrial production, consumer prices, and the won/dollar exchange rate, respectively, and Δ indicates the rate of change (%). $(\mathit{call\text{-}ffr})_t$ represents the difference between the call rate and the federal funds rate. For the difference VAR, the results are similar regardless of whether $(call\text{-}ffr)_t$ or $\Delta(call\text{-}ffr)_t$ is used, but the former is more explicit and thus described in the main text.

$D_{i,t}$ denotes a dummy variable. Since extreme values occurring during crisis periods are highly likely to distort the estimation results of the linear model, this study uses two dummy variables (D_1 and D_2): one for the currency crisis period and another for the global financial crisis period (I=2).[3]

Because Korea is a small open economy, its domestic macroeconomic variables are significantly affected by global shocks. In particular, estimating a reduced-form model instead of a Keynesian structural model is highly likely to overlook not only reverse causality but also third factors. Therefore, this study considers the dollar index and WTI oil price as exogenous factors representing global influences. To avoid estimation complexity, the two error terms (J=2), $\eta_{dindex,t}$ and $\eta_{WTI,t}$ obtained via the ARMA (autoregressive moving average) or MA (moving average) model are treated as exogenous shock variables in Equation (4).

Assume that $u_t = (u_{IP,t}, u_{CPI,t}, u_{call-ffr,t}, u_{won/\$,t})'$ and $E(u_t u_t') = \Omega$. Equation (4) can be expressed in the following reduced form.

$$\begin{aligned} Z_t = {} & \Gamma_0^{-1} a_0 + \Gamma_0^{-1} \textstyle\sum_{i=1}^{I} a_{Di} D_{i,t} + \Gamma_0^{-1} \sum_{j=1}^{J} a_{Sj} \eta_{j,t} \\ & + \Gamma_0^{-1} \textstyle\sum_{p=1}^{P} A_p Z_{t-p} + \Gamma_0^{-1} u_t \end{aligned} \tag{5}$$

$$Z_t = b_0 + \textstyle\sum_{i=1}^{I} b_{Di} D_{i,t} + \sum_{j=1}^{J} b_{Sj} \eta_{j,t} + \sum_{p=1}^{P} B_p Z_{t-p} + \varepsilon_t \tag{6}$$

At time *t*, the response function of Z_{t+k} to a structural prediction error (u_t) shock can be easily estimated by converting Equation (6) into MA form and assuming a Cholesky decomposition.

2. Estimation Results

First, Table 5 shows the results of estimating the dollar index and WTI oil prices using the ARMA or MA model. For the entire period, level variables are used in the

[3] Since treating the COVID-19 period as a dummy variable yields no significant difference in the impulse response analysis results, the main text describes the case where this dummy variable is not used.

VAR model estimation. Therefore, the error terms for *$index* and *WTI* are derived from the level ARMA(1,1) estimation results, and these error terms are used as exogenous global shock variables in the VAR model. Both the AR and MA parameter estimates are statistically significant at the 1% level. Even when using error terms obtained from the difference variable ARMA(1,1) estimation results instead of the level variable, the VAR model's impulse response analysis results are similar. For the other subdivided periods, the impulse response analysis results for the difference variable MA(1) and the difference variable ARMA(1,1) are similar. However, since the LR test results indicate that the null hypothesis that the AR parameter estimate is zero is accepted, the difference variable MA(1) is used.

Table 5. ARMA Estimates

Variable	Period	Estimation Model	AR	MA	Constant
$index	'90.04–'25.05	Level ARMA(1,1)	0.998 (0.003)**	–0.416 (0.044)**	4.532 (0.251)**
	'90.04–'97.10	Difference MA(1)		–0.377 (0.080)**	0.197 (0.133)
	'99.01–'08.08	Difference MA(1)		–0.298 (0.094)**	–0.124 (0.121)
	'14.01–'25.05	Difference MA(1)		–0.377 (0.080)**	0.197 (0.133)
WTI	'90.04–'25.05	Level ARMA(1,1)	0.982 (0.009)**	–0.312 (0.047)**	3.710 (0.290)**
	'90.04–'97.10	Difference MA(1)		–0.375 (0.080)**	–0.344 (1.303)
	'99.01–'08.08	Difference MA(1)		–0.102 (0.095)	2.009 (0.792)**
	'14.01–'25.05	Difference MA(1)		–0.375 (0.080)**	–0.344 (1.303)

Note: ** indicates significance at the 1% level.

Table 6 shows only the estimation results for the global shock variable and crisis dummy variable for the last variable, *won/$* in the four-variable VAR model of Equation (6). To save space, the estimation results for the constant term and lagged variables are omitted. The estimation results for the other three variables are also omitted for the same reason. For the entire period (April 1990 to May 2025), upward shocks to both *$index* and *WTI* cause *won/$* to rise. Notably, a 1%p upward shock to *$index* increases *won/$* at nearly a 1:1 ratio and is statistically significant at the 1%

level. The parameter estimates for the currency crisis and global financial crisis dummy variables are both positive, indicating that the won/$ exchange rate rises more during crisis periods than during normal times. Even excluding the crisis periods, the parameter estimates for both the *$index* and *WTI* shocks are positive. The parameter estimates of the *$index* shock have been increasing more recently.

Table 6. Estimates for Global Shock and Crisis Dummy Variables in the Won/$ Equation

Period	Estimation Model	$\eta_{dindex,t}$	$\eta_{WTI,t}$	$D_{currency,t}$	$D_{global,t}$	R^2
’90.04–’25.05	Level VAR(4)	1.083 (0.095)**	0.017 (0.012)	0.045 (0.008) **	0.022 (0.007) **	0.988
’90.04–’97.10	Difference VAR(1)	0.194 (0.067) **	0.007 (0.010)	–	–	0.272
’99.01–’08.08	Difference VAR(2)	0.870 (0.162) **	0.024 (0.020)	–	–	0.316
’14.01–’25.05	Difference VAR(2)	1.111 (0.102) **	0.209 (0.011) **	–	–	0.589

Notes: 1) Only the estimation results for the global shock variable and crisis dummy variable for the last variable, won/$ in the 4-variable VAR model of equation (6), are shown.

2) $\eta_{dindex,t}$ ($\eta_{WTI,t}$) denotes the dollar index (WTI) shock variable estimated from the ARMA or MA model, and $D_{currency,t}$ ($D_{global,t}$) denotes the currency crisis (global financial crisis) dummy variable, respectively.

3) ** indicates significance at the 1% level.

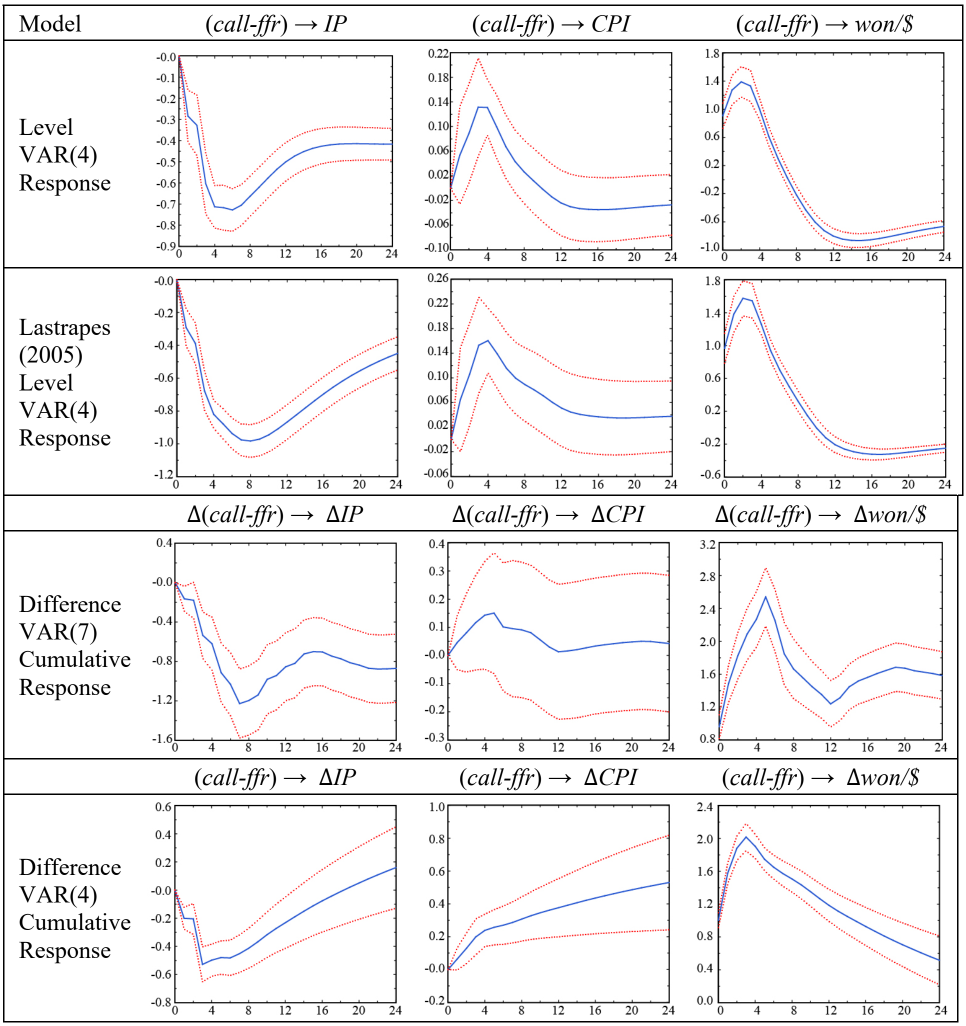
VI. Impulse Response Analysis

1. Entire Period (April 1990–May 2025)

Figure 2 examines the impulse response curves over the entire period (April 1990–May 2025). The first row of figures shows the responses of *IP*, *CPI*, and *won/$* to a 1%p increase in (*call-ffr*) in a level VAR model with four variables and lags. In Figure 2, the solid lines show the estimates of the impulse response function, and the dotted lines above and below the solid lines indicate the 95% confidence bands (response estimate ± 1.96 × one standard deviation) obtained through 1,000 Monte Carlo simulations. As seen in the cointegration test, the null hypothesis H_0: *r*=3 is rejected, so level variables are used. An upward shock to (*call-ffr*) causes *IP* to decline in a U-shaped pattern. Conversely, the *CPI* exhibits an inverse U-shaped response: it rises until the fourth quarter, then declines, showing negative values after the eighth quarter.

An initial price puzzle phenomenon emerges. The *won/\$* also displays an inverse U-shaped response similar to the *CPI*. It rises until the third quarter, then the rate of increase slows, showing a negative response after the eighth quarter.

Figure 2. Impulse Response Results I (Entire Period: April 1990–May 2025)



Notes: 1) For difference variables, the cumulative impulse response curve is displayed.

2) The dotted lines above and below the solid lines indicate the 95% confidence bands (response estimate ± 1.96×one standard deviation) obtained through 1,000 Monte Carlo simulations.

The figures in the second row show the impulse response curves using Lastrapes' (2005, 2006) level VAR model with 6 variables and 4 lags. The 4-variable VAR model in the first row uses the error terms of *$index* and *WTI*, obtained from the ARMA(1,1) model, as exogenous variables. In contrast, the 6-variable VAR model in the second row assumes that domestic shocks do not affect the global variables *$index* and *WTI*.[4] The responses of *IP*, *CPI*, and *won/$* to a 1%p shock (*call-ffr*) are not significantly different from those in the first row. The lag number of 4 are selected based on the AICc criterion, which represents an intermediate lag order among AIC, AICc, and SIC.

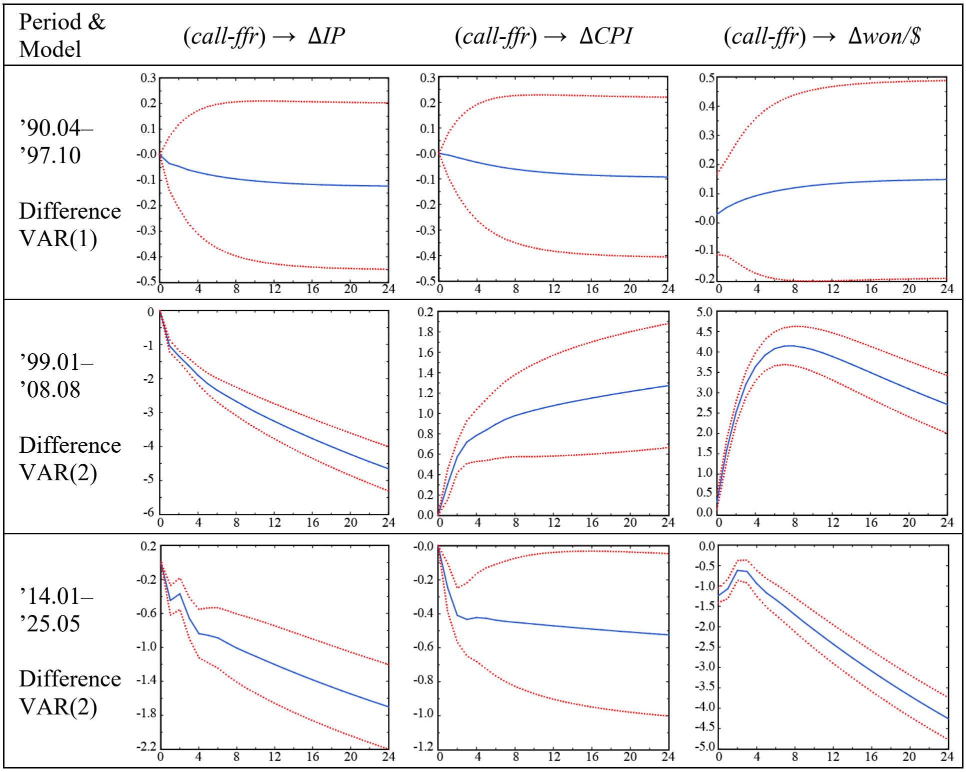
The third row of figures shows the cumulative impulse response curves when using the difference variable VAR(7) model instead of the level variable VAR(4) model in the first row. *IP*, *CPI*, and *won/$* all exhibit responses similar in shape to those of the level variables, but unlike the level variables, *CPI* and *won/$* do not show negative responses. The fourth row of figures displays the cumulative response of the difference variable VAR(4), but uses (*call-ffr*) instead of Δ(*call-ffr*) as the interest rate differential between the two countries. The *CPI* response shows a continuous increase, while the *won/$* response shows a continuous decrease.

2. Period Classification

Since the impulse response analysis results in Figure 2 are unclear, Figure 3 shows the impulse response curves for each sub-period, excluding the currency crisis and global financial crisis periods, which contain extreme values. For the period before the currency crisis (April 1990–October 1997), the lag number is determined as 1 based on the AICc criterion, and as seen in the cointegration test, the null hypothesis H_o: r=0 is accepted. Therefore, the figures in the first row show the cumulative responses of *IP*, *CPI*, and *won/$* to a 1%p upward shock to the (*call-ffr*) rate, estimated using a difference variable VAR(1) model. An increase in (*call-ffr*) causes *IP* and *CPI* to decline. The initial price puzzle phenomenon observed in the impulse response curve for the entire period is not visible. Meanwhile, the won/$ exchange rate rises in response to the relative increase in domestic interest rates. This result is largely consistent with existing research findings that relative monetary tightening causes developing country exchange rates, including the *won/$*, to rise.

[4] For the VAR model of Lastrapes (2005, 2006), see Appendix B.

Figure 3. Cumulative Impulse Response Results II



Notes: 1) For difference variables, the cumulative impulse response curve is displayed.

2) The dotted lines above and below the solid lines indicate the 95% confidence bands (response estimate ± 1.96×one standard deviation) obtained through 1,000 Monte Carlo simulations.

For the period from the post-currency crisis to the pre-global financial crisis (January 1999–August 2008), the lag number is determined to be 2 based on the AICc criterion, and the null hypothesis H_0: r=0 is not rejected in the cointegration test. Therefore, the figures in the second row show the cumulative responses of the *IP*, *CPI*, and *won/$* to a 1%p shock to the (*call-ffr*), estimated using a difference variable VAR(2) model. An upward shock to (*call-ffr*) reduces *IP* while increasing *CPI*. Unlike the pre-currency crisis period, a price puzzle phenomenon emerges. The *won/$* exhibits an inverse U-shaped response, consistent with the overall period. Overshooting gradually appears over eight quarters before declining, consistent with existing research findings.

Finally, since Korea has become not only a net foreign creditor nation but also a net foreign financial asset holder following the global financial crisis, I examine the period from this point until recently (January 2014 to May 2025). During this period, for the four variables, the lag number is determined to be 2 based on the AICc criterion, and the cointegration test rejects the null hypothesis: $H_0: r=0$. First, the figures in the third row show the cumulative responses of *IP*, *CPI*, and *won/$* to a 1%p shock to the (*call-ffr*) rate, estimated using a VAR(2) model with difference variables.[5] The upward shock to the (*call-ffr*) reduces not only *IP* but also *CPI*, indicating that the price puzzle phenomenon does not occur. Furthermore, unlike in previous periods, a rise in domestic interest rates relative to U.S. interest rates causes the won/$ exchange rate to decline.[6]

The first row of figures in Figure 4 shows the cumulative responses of the difference VAR(2) model of Lastrapes (2005, 2006) over the same period. As in the previous case, *IP*, *CPI*, and *won/$* decline. However, as the second row of figures indicates, when overseas shock variables such as *$index* and *WTI* are not considered, *won/$* increases. For small open economies, it is evident that major foreign variables significantly influence the causal relationships between variables.[7] The third row of figures shows the response curves obtained from the ECM(2) model, reflecting the cointegration relationship present during this period. The ECM(2) can be represented as a level

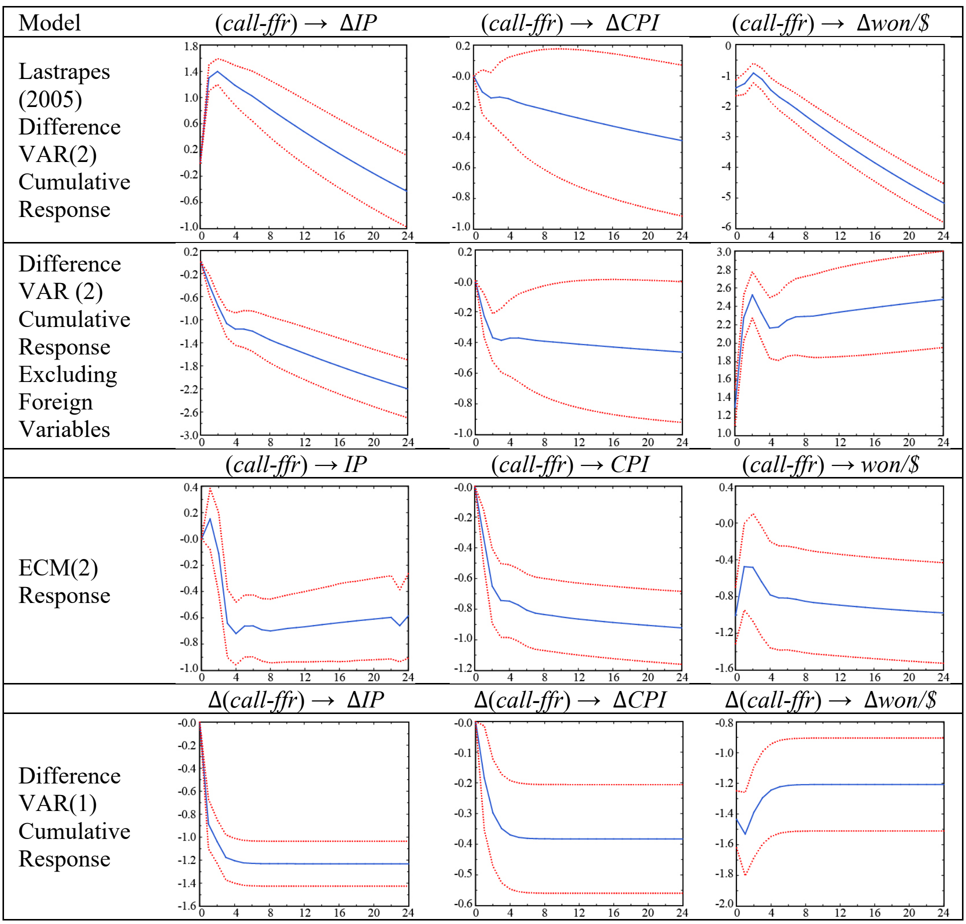
[5] For information on changing the order of variables or using *call* instead of (*call-ffr*), see Appendix C.

[6] During President Lee Myung-bak's term in office (February 2008–February 2013), the average inflation rate reached 3.6%, despite the revision of the Consumer Price Index in 2011, due to the aftermath of the global financial crisis and a sustained policy of maintaining a high exchange rate (a weak won). Such artificial government intervention in exchange rates and prices can make it difficult to accurately assess the effectiveness of monetary policy. Reflecting the low growth and high inflation of this period, Figure D1 in the Appendix D shows that during the periods '10.01–'25.05 or '11.01–'25.05, the estimated responses of industrial production and consumer prices derived from Cholesky decomposition deviate from or hover near the 97.5th percentile obtained using sign restrictions based on economic theory, suggesting that a relevant sign reversal may have occurred at an earlier point in time. In contrast, during the period from '14.01–'25.05, when Korea became both a net foreign creditor and a net foreign asset holder, these estimates do not differ significantly from the median; therefore, the results for this period are described in the main text. Even during the period from the post-global financial crisis to the present (July 2009 to May 2025), monetary tightening shocks cause the won/dollar exchange rate to decline. For an analysis of impulse responses using both contemporary sign restrictions and Cholesky decomposition, see Appendix D.

[7] For cases where additional global variables related to volatility or term spreads are used, refer to Appendix E.

variable VAR(3) with parameter constraints (see Hamilton, 1994). The impulse response curve shows that, similar to the cumulative response curve of the difference VAR(2), a 1%p increase in (*call-ffr*) causes declines in *IP*, *CPI*, and *won/\$*. The fourth row of figures shows the cumulative response from the difference VAR(1) model to a 1%p increase in Δ(*call-ffr*) instead of (*call-ffr*). As with the other models, *IP*, *CPI*, and *won/\$* all decline. The lag number is 1 for all criteria: AIC, AICc, and SIC.

Figure 4. Impulse Response Results III (Period: January 2014–May 2025)



Notes: 1) For difference variables, the cumulative impulse response curve is displayed.

2) ECM(2) is converted and displayed as a level variable VAR(3).

3) The dotted lines above and below the solid lines indicate the 95% confidence bands (response estimate ± 1.96×one standard deviation) obtained through 1,000 Monte Carlo simulations.

Previous studies using panel data have included Korea among emerging market economies, arguing that emerging market economies exhibit the price puzzle phenomenon and experience rising exchange rates. However, examining the recent period (January 2014–May 2025) using a model incorporating overseas variables reveals that, like advanced economies, Korea does not exhibit the price puzzle phenomenon. A rise in the interest rate differential between Korea and the U.S. causes the won/dollar exchange rate to fall, not rise.

3. Reasons for Differences in Reactions in the VAR Model

In Equations (5) and (6), u_t and ε_t denote the error terms ($\varepsilon_t = \Gamma_0^{-1} u_t$) of the structural and reduced-form VAR models, respectively. In Table 7, Γ_0 indicates the direct causal relationship during the same period, while Γ_0^{-1} represents the overall causal relationship during the same period, encompassing both direct and indirect effects. As shown in Table 7, for the period January 2014 to May 2025 using the difference VAR(2) model, a 1%p increase in (*call-ffr*) causes a 1.240%p decrease in *won/$* during the same period. Assuming Cholesky decomposition and given the variable order *IP*, *CPI*, (*call-ffr*), *won/$* (from left to right), the direct effect and the overall effect are identical during the same period.[8] However, unlike the period after the global financial crisis, in the period from April 1990 to October 1997 and the period from January 1999 to August 2008, when the difference variable VAR(1) and VAR(2) models are used, a 1%p increase in (*call-ffr*) raises the won/$ exchange rate by 0.030%p and 0.323%p, respectively, during the same period.

Table 8 presents the estimated values of Γ_0 and Γ_0^{-1} for the period January 2014 to May 2025. Unlike Table 7, it uses a differenced 4-variable VAR(2) model excluding the global exogenous shock variable composed of *$index* and *WTI*, and a differenced

[8] For the period January 2014 to May 2025, a 1%p increase in ΔIP leads to a 0.01718%p rise in ΔCPI during the same period. Additionally, it directly increases (*call-ffr*) by 0.00848%p during the same period, but indirectly reduces it by 0.00058%p (-0.01718×0.03393) through ΔCPI, resulting in an overall increase in (*call-ffr*) of 0.00790%p. Meanwhile, a 1%p increase in ΔIP directly raises *won/$* by 0.00215%p during the same period. Indirectly, through ΔCPI and (*call-ffr*), it increases *won/$* by 0.00429%p ($0.01718\times0.81980+0.01718\times(-0.03393)\times(-1.23984)+0.00848\times(-1.23984)=0.00429$), resulting in an overall increase of 0.00644%p in the *won/$* during the same period. A 1%p increase in ΔCPI directly raises the *won/$* by 0.820%p during the same period and increase it by 0.862%p overall. For the difference between direct effects and overall effects, see Appendix F.

6-variable VAR(2) model incorporating these variables, as proposed by Lastrapes (2005). In the former model without foreign variables, a 1%p increase in (*call-ffr*) raises the won/\$ exchange rate by 1.306%p, similar to the pre-global financial crisis period. In contrast, the latter model, assuming block exogeneity, lowers it by 1.405%p, similar to Table 7. These responses for the same period are marked at period 0 in each impulse response curve figure.

The response function of the difference VAR(2) model including foreign shock variables after period 0 is determined by the parameters Γ_0^{-1} and the reduced form parameters B_1 and B_2. According to the estimation results, tight monetary policy, measured by the rise in the (*call-ffr*), lowers the won/dollar exchange rate during the same period. Furthermore, tight monetary policy lowers prices in the next period, and falling prices cause the won/dollar exchange rate to decline. However, tight monetary policy reduces industrial production in the next period, causing the won/dollar exchange rate to rise. Therefore, contrary to the empirical results of previous studies, the liquidity demand effect and the deflation effect are larger than the industrial

Table 7. Γ_0 and Γ_0^{-1} Estimates

Period		Γ_0				Γ_0^{-1}			
		IP	*CPI*	*call-ffr*	*won/\$*	*IP*	*CPI*	*call-ffr*	*won/\$*
’90.04–’97.10	*IP*	1.000	0.000	0.000	0.000	1.000	0.000	0.000	0.000
	CPI	–0.007	1.000	0.000	0.000	0.007	1.000	0.000	0.000
	call-ffr	0.052	0.002	1.000	0.000	–0.052	–0.002	1.000	0.000
	won/\$	0.067	0.014	–0.030	1.000	–0.068	–0.014	0.030	1.000
’99.01–’08.08	*IP*	1.000	0.000	0.000	0.000	1.000	0.000	0.000	0.000
	CPI	0.002	1.000	0.000	0.000	–0.022	1.000	0.000	0.000
	call-ffr	–0.001	0.009	1.000	0.000	0.002	–0.009	1.000	0.000
	won/\$	–0.043	–1.356	–0.323	1.000	0.014	1.353	0.323	1.000
‘14.01–’25.05	*IP*	1.000	0.000	0.000	0.000	1.000	0.000	0.000	0.000
	CPI	–0.017	1.000	0.000	0.000	0.017	1.000	0.000	0.000
	call-ffr	–0.008	0.034	1.000	0.000	0.008	–0.034	1.000	0.000
	won/\$	–0.002	–0.820	1.240	1.000	0.006	0.862	–1.240	1.000

Notes: 1) The estimated models by period are identical to those in Table 6.
2) Γ_0 (Γ_0^{-1}) indicates direct (overall) causality within the same period under the Cholesky decomposition assumption.

production reduction effect, leading to the emergence of an "advanced-economy-type forward premium puzzle" phenomenon.[9][10] The results remain similar even when

Table 8. Γ_0 and Γ_0^{-1} Estimates (Period: January 2014−May 2025)

Model		Variable	*$index*	*WTI*	*IP*	*CPI*	*call-ffr*	*Won/$*
Difference VAR (2) Excluding Foreign Variables	Γ_0	*IP*	–	–	1.000	0.000	0.000	0.000
		CPI	–	–	–0.015	1.000	0.000	0.000
		call-ffr	–	–	–0.010	0.040	1.000	0.000
		Won/$	–	–	0.031	–1.041	–1.306	1.000
	Γ_0^{-1}	*IP*	–	–	1.000	0.000	0.000	0.000
		CPI	–	–	0.015	1.000	0.000	0.000
		call-ffr	–	–	0.010	–0.040	1.000	0.000
		Won/$	–	–	–0.002	0.988	1.306	1.000
Lastrapes (2005) Difference VAR(2)	Γ_0	*$index*	1.000	0.000	0.000	0.000	0.000	0.000
		WTI	3.058	1.000	0.000	0.000	0.000	0.000
		IP	0.018	0.018	1.000	0.000	0.000	0.000
		CPI	–0.017	–0.003	–0.007	1.000	0.000	0.000
		call-ffr	–0.021	0.003	–0.011	0.020	1.000	0.000
		Won/$	–1.126	–0.024	–0.023	–0.901	1.405	1.000
	Γ_0^{-1}	*$index*	1.000	0.000	0.000	0.000	0.000	0.000
		WTI	–3.058	1.000	0.000	0.000	0.000	0.000
		IP	0.036	–0.018	1.000	0.000	0.000	0.000
		CPI	0.008	0.003	0.007	1.000	0.000	0.000
		call-ffr	0.030	–0.003	0.011	–0.020	1.000	0.000
		Won/$	1.020	0.030	0.015	0.929	–1.405	1.000

Notes: 1) The estimated models by period are identical to those in Table 6.
2) Γ_0 (Γ_0^{-1}) indicates direct (overall) causality within the same period under the Cholesky decomposition assumption.

[9] In their conclusion, Bansal and Dahlquist (2000, p. 140) stated that the forward premium puzzle—the negative correlation between expected exchange rates and interest rate differentials—had implications which seemed anomalous from the perspective of economic models. However, post-2008 CIP deviations or cross-currency basis movements may weaken the mapping from interest differentials to forward premia.

[10] When using the available new series of monetary aggregates and nominal GDP, the average ratio of M2/GDP (MB/GDP, M1/GDP) ratio was 3.786 (0.163, 1.082) during the period '03.10–'08.09, but increased to 5.441 (0.372, 1.856) during the period '14.01–'25.06. The strength of the liquidity effect is an increasing function of the money-to-GDP ratio.

using a 6-variable difference VAR(2) model (Lastrapes, 2005) with block exogeneity constraints—where domestic variables cannot affect foreign variables—instead of the foreign shock variable estimated from the ARMA model.

According to the estimation results of the difference VAR(2) model excluding external shock variables, the tightening monetary policy measured by the rise in the (*call-ffr*) actually increases the won/dollar exchange rate. In addition, tightening monetary policy reduces industrial production in the next period, thereby increasing the won/dollar exchange rate. However, tight monetary policy lowers prices in the subsequent period, and falling prices cause the won/dollar exchange rate to decline. In this case, the expected inflation effect and the industrial production decline effect outweigh the deflationary effect, leading to a "developing-country-type forward premium puzzle". However, this result arises because it does not account for the influence of third factors—external variables, particularly the dollar index—on the won/dollar exchange rate. Table 6 and Table 8 show that a 1%p increase in the dollar index leads to a respective 1.111%p and 1.020%p rise in the won/dollar exchange rate during the same period.[11]

VII. Economic Implications

From a theoretical perspective, demand for won-denominated assets depends on the relatively expected return on won-denominated assets $(i - i^* - E_t \Delta S_{t+1}/S_t)$. Assuming free capital mobility and constant won/dollar exchange rates and risk premiums, an increase in domestic interest rates boosts demand for won-denominated assets, causing the won/dollar exchange rate to fall. In the Keynesian model, prices are assumed constant in the short run, so changes in nominal and real interest rates are identical. However, as in the Chicago School theory, if domestic expected inflation rises significantly, leading to a larger decline in the expected value of the won than the increase in domestic interest rates, the relatively expected return and demand for won assets decrease. Consequently, the won/dollar exchange rate rises. In other words,

[11] Table 6 shows that, using a difference VAR(1) model incorporating foreign shock variables, a 1%p increase in the dollar index during the period April 1990 to October 1997 leads to a 0.194%p increase in the won/dollar exchange rate during the same period. However, during this period, even when considering foreign shock variables, the effects of expected inflation and reduced industrial production outweigh the deflationary effect, resulting in an "emerging-market-economy-type forward premium puzzle" phenomenon.

when domestic interest rates rise due to increased expected inflation, the value of the won falls.

Empirically, existing studies using panel data included the won/dollar exchange rate in developing country exchange rates, arguing that when domestic interest rates rose, the domestic currency depreciated unlike in advanced economies. While the results of this existing research are similar to the empirical findings of this study for the period before the global financial crisis, they differ in the recent period. Instead, the response of the won/dollar exchange rate to domestic interest rate increases in the recent period is similar to that observed in advanced economies.

The claim in existing studies that the won/dollar exchange rate reacted similarly to developing country exchange rates might stem from the fact that these studies analyzed periods with relatively high expected inflation, as shown by the sample statistics in Table 3. In addition, these studies might also have failed to adequately account for extreme values occurring during periods such as the currency crisis or the global financial crisis. Furthermore, it cannot be ruled out that the panel analysis failed to adequately account for the significant influence of global variables on the small open economy of Korea. Recently, since Korea became a net external financial asset holder after the third quarter of 2014, the major participants in the won/dollar foreign exchange market consist of large domestic pension funds seeking rational investment returns rather than speculative forces. Therefore, the won/dollar exchange rate typically follows the patterns of advanced economies during normal times.

VIII. Summary and Conclusion

This study analyzes the effect of positive interest rate differential shocks on the won/dollar exchange rate using monthly data from March 1990, when the market-average exchange rate system was introduced, through May 2025. Given that the analysis period encompasses events such as the currency crisis, the global financial crisis, and changes in economic structure, the data are examined not only for the entire period (April 1990 to May 2025) but also divided into the pre-currency crisis period (April 1990 to October 1997), the post-currency crisis to pre-global financial crisis period (January 1999–August 2008), and the period from when Korea became a net external financial asset holder to the present (January 2014–May 2025). Based on the cointegration test results, level-variable VAR are used for the entire period (April 1990–May 2025), difference variable VAR are used for the period April 1990–

October 1997 and January 1999–August 2008, and difference-variable VAR model or error correction model are used for the period January 2014–May 2025 for estimation and impulse response analysis. In addition, to account for the influence of global variables, I estimates not only VAR models using *$index* and *WTI* error terms obtained from ARMA(1,1) or MA(1) models as exogenous variables, but also Lastrapes (2005, 2006) VAR models that incorporate global variables as endogenous variables while assuming block exogeneity.

For the entire period (April 1990–May 2025), an upward shock to the call rate minus the federal funds rate causes industrial production to decline in a U-shaped pattern, while prices exhibit the opposite, inverted U-shaped response, initially presenting a price puzzle phenomenon. The won/dollar exchange rate and prices change similarly, showing a negative reaction after the eighth quarter. However, since the impulse response analysis results for the entire period are unclear, I examine the impulse response curves for sub-periods, excluding the currency crisis and global financial crisis periods, which contain extreme values.

During the pre-currency crisis (April 1990 to October 1997), upward shocks to the call rate minus federal funds rate causes declines in industrial production and prices. The initial price puzzle phenomenon observed in the impulse response curve for the overall period is not visible. Meanwhile, the won/dollar exchange rate rises in response to relative domestic interest rate hikes. This result is largely consistent with existing international research showing that tight monetary policy raised exchange rates in developing countries, including the won/dollar rate.

For the period from the post-currency crisis to the pre-global financial crisis (January 1999–August 2008), an upward shock in call rate minus federal funds rate reduces industrial production while increasing prices. Unlike the pre-currency crisis period, a price puzzle phenomenon emerges. The won/dollar exchange rate exhibits an inverse U-shaped response, similar to the overall period. The overshooting phenomenon gradually emerges over eight quarters before declining, consistent with existing research.

From the period when Korea became a net external financial asset holder until recently (January 2014–May 2025), an upward shock in call rate minus federal funds rate reduces not only industrial production but also prices, indicating the absence of the price puzzle phenomenon. Furthermore, unlike previous periods, tight monetary policy causes the won/dollar exchange rate to decline. Because the effects of increased liquidity demand and deflation outweigh the effect of reduced industrial production, a so-called “forward premium puzzle” phenomenon similar to that seen in advanced economies emerges, contrary to claims in existing foreign studies.

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APPENDIX A. Qu and Perron (2007) Tests

In Appendix A, I aim to identify endogenous breakpoints present throughout the entire analysis period (April 1990–May 2025) using the Qu and Perron (2007) test for VAR models. The Qu and Perron (2007) test is a statistical method used to examine whether there are structural changes in the regression coefficients and the covariance matrix. The Qu and Perron (2007) test not only identifies multiple breakpoints but also estimates the dates on which structural changes occur and provides confidence intervals for these dates.

Table A1 presents the results of tests for the presence of two breakpoints in the regression coefficients and covariance matrix in level and difference 4-variable VAR models composed of [*IP*, *CPI*, (*call-ffr*), *won/$*] and [Δ*IP*, Δ*CPI*, (*call-ffr*), Δ*won/$*], respectively. Lags of 1, 2, and 4 are used. When a lag of 1 and difference variables are used, the two breakpoints occur in March 1999 and June 2009, respectively. The 95% confidence interval for the first break point spans from June 1998 to April 1999, and the 95% confidence interval for the second break point spans from May 2008 to January 2010. The sup LR_T is a test statistic for a fixed number of break points; the test statistic for the null hypothesis that there are 0, rather than 1, break points in the coefficients and covariance matrix is 906.054, which is rejected at the 1% significance level. Similarly, the test statistic for the null hypothesis that there are 0, rather than 2, break points is 1,050.946, which is rejected at the 1% significance level. Furthermore, as shown by Seq(2|1), when one break point is added to the existing single break point, the test statistic is 144.892, and the null hypothesis that there is one break point is rejected at the 1% significance level. Even when using level variables or increasing the lag number, it can be observed that structural breaks occur at similar points in time due to the effects of the currency crisis and the global financial crisis.

Table A1. Qu and Perron (2007) Test Results

Lag	Variables		*IP*, *CPI*, (*call-ffr*), *won/$*	Δ*IP*, Δ*CPI*, (*call-ffr*), Δ*won/$*
1	Break Point	1	’99.04(’99.03–’99.05)	’99.03(’98.06–’99.04)
		2	’09.01(’08.12–’09.03)	’09.06(’08.05–’10.01)
	Sup LR_T Test	0 vs. 1	981.311**	906.054**
		0 vs. 2	1,179.164**	1,050.946**
	Seq(2|1) Test		190.321**	144.892**
2	Break Point	1	’99.02(’98.12–’99.03)	’99.03(’99.02–’99.04)
		2	’09.01(’08.12–’09.02)	’08.10(’08.03–’09.06)
	Sup LR_T Test	0 vs. 1	1,042.264**	1,026.404**
		0 vs. 2	1,224.118**	1,153.043**
	Seq(2|1) Test		181.854**	126.038**
4	Break Point	1	’99.04(’99.03–’99.05)	’99.04(’99.01–’99.05)
		2	’09.01(’08.12–’09.02)	’08.10(’08.05–’09.03)
	Sup LR_T Test	0 vs. 1	1,079.516**	1,062.839**
		0 vs. 2	1,293.246**	1,238.668**
	Seq(2|1) Test		213.729**	175.830**

Notes: 1) The Table displays the test result when two structural break points exist in the coefficient and covariance matrix.

2) Periods in parentheses indicate confidence intervals at the 95% level.

3) ** denotes significant at the 1% level.

Appendix B. The VAR Model of Lastrapes (2005)

For a small open economy like Korea, where domestic variables such as industrial production or prices have a negligible impact on the dollar index or international oil prices, we can estimate the following VAR model assuming block exogeneity instead of Equation (6) (Hamilton, 1994; Lastrapes, 2005).

$$\begin{pmatrix} Z_{F,t} \\ Z_{D,t} \end{pmatrix} = \begin{pmatrix} b_{F,0} \\ b_{D,0} \end{pmatrix} + \sum_{i=1}^{p} \begin{pmatrix} B_{FF}^{i} & 0 \\ B_{DF}^{i} & B_{DD}^{i} \end{pmatrix} \begin{pmatrix} Z_{F,t-i} \\ Z_{D,t-i} \end{pmatrix} + \begin{pmatrix} \varepsilon_{F,t} \\ \varepsilon_{D,t} \end{pmatrix} \qquad \text{(B1)}$$

In Equation (B1), B_{FD}^{i}=0 and Z_F is block exogenous with respect to Z_D. Z_F is a 2×1 vector consisting of the dollar index and WTI oil price, while Z_D is a 4×1 vector comprising industrial production, consumer prices, the Korea-US interest rate differential, and the won/dollar exchange rate. For the full period, dummy variables are added to Equation (B1). Equation (B1) can be expressed as the following OLS estimation equation.

$$Z_{F,t} = b_{F,0} + \sum_{i=1}^{p} B_{FF}^{i}\, Z_{F,t-i} + \varepsilon_{F,t},\ E(\varepsilon_F \varepsilon_F') = \Omega_{FF} \qquad \text{(B2)}$$

$$Z_{D,t} = c_{D,0} + \sum_{i=0}^{p} C_i\, Z_{F,t-i} + \sum_{i=1}^{p} B_{FF}^{i}\, Z_{D,t-i} + \nu_{D,t},\ E(\nu_D \nu_D') = H \qquad \text{(B3)}$$

$c_{D,0}$, C_0, C_i, and $E(\nu_D \nu_D')$ in Equation (B3) can be expressed as follows.

$$c_{D,0} = b_{D,0} - C_0 b_{F,0} \qquad \text{(B4)}$$

$$C_0 = \Omega_{DF}\Omega_{FF}^{-1} \qquad \text{(B5)}$$

$$C_i = B_{DF}^{i} - c_0 B_{FF}^{i},\ i=1,\ \ldots,\ p \qquad \text{(B6)}$$

$$H = \Omega_{DD} - \Omega_{DF}\Omega_{FF}^{-1}\Omega_{FD} \qquad \text{(B7)}$$

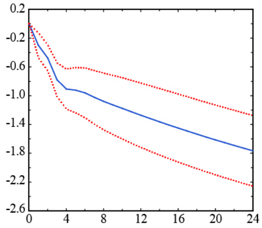
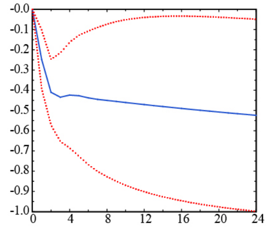
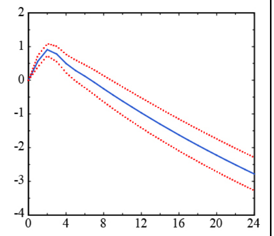
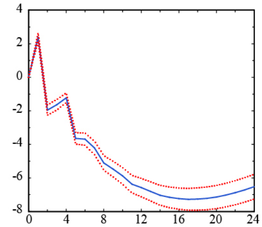
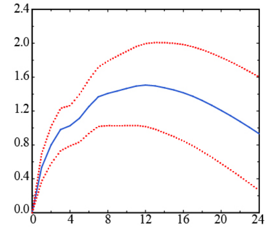
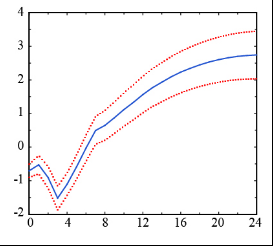
The parameters $B(L)$ and Ω in Equation (B1) can be derived using Equations (B4) to (B7) after estimating Equations (B2) and (B3) using OLS.

Appendix C. Impulse Response Analysis with Changed Variable Order or Use of $call_t$

In this study, as in previous studies, the variables are listed in the order, $\Delta IP_t, \Delta CPI_t,$ (*call-ffr*)$_t$, $\Delta won/\$_t$, and impulse response analysis is conducted using Cholesky decomposition. As suggested by the Taylor rule in VAR analysis models, the policy interest rate is determined by considering the growth gap and the inflation gap. Due to the assumptions of the Cholesky decomposition, ΔIP_t and ΔCPI_t influence (*call-ffr*)$_t$ in the current period, while (*call-ffr*)$_t$ influences ΔIP_{t+1} and ΔCPI_{t+1} in the next period. Since neither Korea nor the United States has adopted an exchange rate targeting regime, I assume that $\Delta won/\$_t$ influences the next period's ΔIP_{t+1}, ΔCPI_{t+1}, and (*call-ffr*)$_{t+1}$. However, even when this assumption is relaxed, the result that a tightening monetary policy shock causes the won/dollar exchange rate to decline during the period from January 2014 to May 2025 remains unchanged. The first row of Figure C1 shows the case where the variables are listed in the order $\Delta won/\$_t$, $\Delta IP_t, \Delta CPI_t,$ (*call-ffr*)$_t$, and it demonstrates that the won/dollar exchange rate initially rises but soon falls. Industrial production and prices also decline accordingly. The same result is obtained even when assuming sign restrictions rather than a Cholesky decomposition (see Appendix D).

The second row of Figure C1 shows the case where $call_t$ is used instead of (*call-ffr*)$_t$. Unlike in the case of (*call-ffr*)$_t$, the won/dollar exchange rate initially declines during the early stages of monetary tightening and then rises. In addition, the price puzzle where prices rise initially despite monetary tightening occurs. These results are similar when $\Delta call_t$ is used instead of $call_t$, or when $\eta_{ffr,t}$ is used as a global shock with $call_t$. But it is difficult to understand the finding that consumer prices rise when the call rate increases for no particular reason. Since Korea is a small open economy with free capital mobility under a floating exchange rate regime, the impact of relative interest rate differentials on the exchange rate and inflation is judged to be more consistent with both theory and reality.

Figure C1. Cumulative Impulse Response Results (Period: January 2014–May 2025)

Var. Order & Model	$(call\text{-}ffr) \rightarrow \Delta IP$	$(call\text{-}ffr) \rightarrow \Delta CPI$	$(call\text{-}ffr) \rightarrow \Delta won/\$$
$\Delta won/\$$, ΔIP, ΔCPI, $(call\text{-}ffr)$, Difference VAR(2)			
Var. Order & Model	$call \rightarrow \Delta IP$	$call \rightarrow \Delta CPI$	$call \rightarrow \Delta won/\$$
ΔIP, ΔCPI, $call$, $\Delta won/\$$, Difference VAR(4)			

Notes: 1) For difference variables, the cumulative impulse response curve is displayed.

2) The dotted lines above and below the solid lines indicate the 95% confidence bands (response estimate ± 1.96×one standard deviation) obtained through 1,000 Monte Carlo simulations.

APPENDIX D. Impulse Response Analysis with Contemporary Sign Restrictions

In Appendix D, I examine the impulse response analysis using contemporary sign restrictions based on economic theory, with the aim of verifying the reliability of the impulse response results obtained via Cholesky decomposition. Existing methods that derive structural shocks from reduced-form shocks through Cholesky decomposition or long-run constraints have the drawback that these shocks may or may not have signs consistent with economic theory. In contrast, impulse response analysis using sign restrictions focuses on estimating structural shocks that not only have no correlation with other reduced-form shocks or their own past shocks from an agnostic perspective but also possess signs restrictions with economic theory.

As seen in Equations (5) and (6), the relationship between the reduced-form shock (ε_t) and the structural shock (u_t) is given by $\varepsilon_t = \Gamma_0^{-1} u_t$. When the structural covariance matrix and the reduced-form covariance matrix are denoted by $E(u_t u_t') = \Omega$ and $E(\varepsilon_t \varepsilon_t') = \Sigma$, respectively, and we assume that $P = \Gamma_0^{-1}\Omega^{1/2}$, then the reduced-form covariance matrix $\Sigma (= \Gamma_0^{-1}\Omega\Gamma_0^{-1\prime})$ is identical to PP'. Using this, ζ_t, which is the structural shock u_t divided by the standard deviation $\Omega^{1/2}$, can be expressed as follows.

$$\zeta_t = \Omega^{-1/2} u_t = \Omega^{-1/2}\Gamma_0 \varepsilon_t = P^{-1}\varepsilon_t \tag{D1}$$

In equation (D1), the covariance matrix $E(\zeta_t \zeta_t')$ of ζ_t becomes the identity matrix ($E(P^{-1}\varepsilon_t \varepsilon_t' P^{-1\prime}) = P^{-1}PP'P^{-1\prime} =$ I). This equation can be transformed into a structural shock using a square matrix with the property of $Q'Q = QQ' =$ I, i.e., an orthogonal matrix, as follows.

$$\varepsilon_t = P\zeta_t = PQ'Q\zeta_t \ = \ P^*\zeta_t^* \tag{D2}$$

In Equation (D2), since the covariance matrix $E(\zeta_t^* \zeta_t^{*\prime})$ of ζ_t^* is the identity matrix—just like the covariance matrix of ζ_t, the covariance matrices of ζ_t and ζ_t^* are identical; however, the effects of ζ_t and ζ_t^* on the reduced-form shock ε_t and the endogenous variable Z_t differ. In this study, I use the Householder transformation method to obtain an orthogonal matrix. First, a 4×4 random variable Φ is drawn from

$N(0, I_4)$, and then decomposed into $\Phi = Q_R R$. I extract the structural shocks ζ_t^* using Q_R generated by QR decomposition, and then calculate the responses of each variable to the structural shocks. I repeatedly calculate the responses to the shocks by selecting those that satisfy the sign constraints consistent with theory and discarding those that do not. In this case, the impulse response functions are responses obtained from each unknown model, unlike confidence intervals.

Figure D1 shows the responses of industrial production, the consumer price index, and the won/dollar exchange rate to monetary tightening shocks over the two periods. The solid line represents the median, while the dotted lines above and below the median indicate the 2.5th and 97.5th percentiles. The sign-restricted impulse response functions are obtained by using the Householder transformation method to select impulse response functions from one million random samples that satisfy the constraints in Table D1, while rejecting those that do not meet these conditions. Therefore, the individual results of these impulse response functions are all agnostic results obtained based on their respective models, and the dotted lines in Figure D1 representing the 2.5th and 97.5th percentiles are a concept distinct from confidence intervals. However, the median, 2.5th, and 97.5th percentile impulse response functions shown in the figure are derived by selecting the median, 2.5th percentile, and 97.5th percentile values from these results at each response point; therefore, it is difficult to view them as results obtainable from any specific model. It is reasonable to interpret that, in cases where the results based on the Cholesky decomposition do not deviate significantly from this range, these results do not pose major problems. The dot-and-dash line shows the estimates from the Cholesky decomposition.

The figures in the first row show the cumulative responses of *IP*, *CPI*, and *won/$* to a 1%p positive shock to (*call-ffr*), estimated using a two-lag difference VAR model with the sign restrictions specified in Table D1, covering the period from the aftermath of the currency crisis to the onset of the global financial crisis (January 1999–August 2008). An upward shock to (*call-ffr*) causes *IP* to decline, while simultaneously increasing *CPI* despite the negative sign restriction during the same period. The won/dollar exchange rate exhibits an inverted U-shaped response. Although the response estimates derived from the Cholesky decomposition are closer to those at the 2.5th and 97.5th percentiles rather than the median, they follow a similar pattern of movement.

The figures in the second row show the case for January 2014 to May 2025, where an upward shock to (*call-ffr*) reduces not only *IP* but also *CPI*, indicating that the price puzzle phenomenon does not occur. Furthermore, unlike in previous periods, tight monetary policy causes *won/$* to decline. The response results do not differ significantly from those assuming Cholesky decomposition.

The figures in the third and fourth rows show the results for the periods from January 2010 to May 2025 and January 2011 to May 2025, respectively. In both cases—whether assuming Cholesky decomposition or sign restrictions—an upward shock to the (*call-ffr*) causes the won/dollar exchange rate to decline. However, the more the analysis period includes the time immediately following the end of the global financial crisis, the more the decline in *CPI* and *IP* is significantly reduced. In particular, during the Lee Myung-bak administration—a period plagued by low growth and high inflation in the aftermath of the global financial crisis—a controversy arose in November 2011 when Statistics Korea revised the price index. The revision involved removing items that had previously driven up inflation or significantly reducing their weight, leading to accusations that the revision was a ploy to artificially lower prices. As a result of this revision, the consumer price inflation rate from January to October 2011 fell from 4.4% to 4.0%—a drop of 0.4 percentage points, the largest decline recorded following an index revision to date.

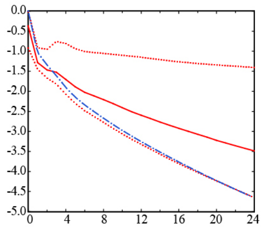
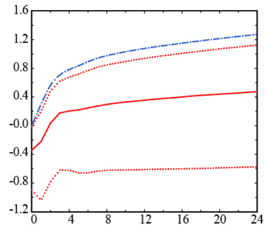
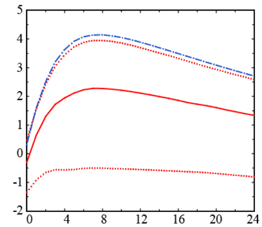
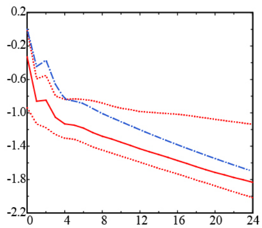
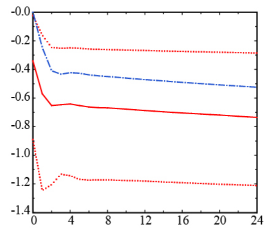
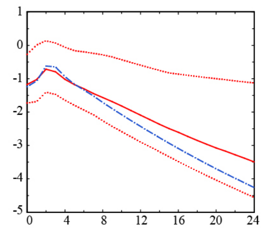
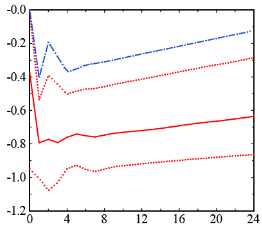
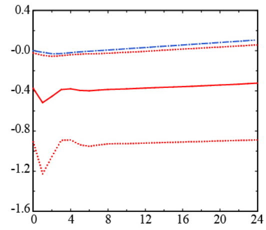
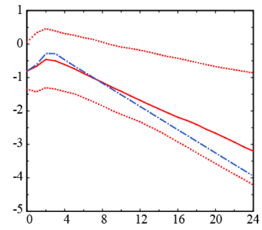
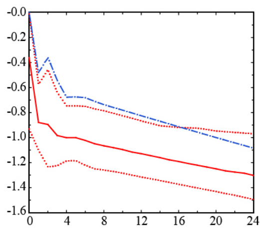
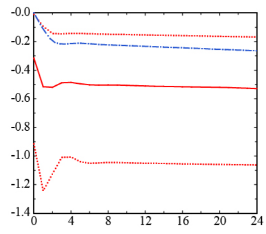
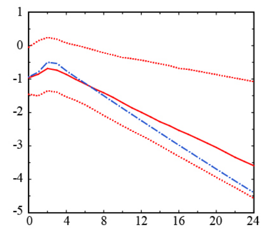
Table D1. Contemporary Sign Restrictions for Each Shock

	IP	*CPI*	*call-ffr*	*won/$*
$u_{IP,t}$	$\frac{\partial Z_t}{\partial u_{IP,t}} \geq 0$	$\frac{\partial Z_t}{\partial u_{IP,t}} \leq 0$	$\frac{\partial Z_t}{\partial u_{IP,t}}$?	$\frac{\partial Z_t}{\partial u_{IP,t}}$?
$u_{CPI,t}$	$\frac{\partial Z_t}{\partial u_{CPI,t}} \geq 0$	$\frac{\partial Z_t}{\partial u_{CPI,t}} \geq 0$	$\frac{\partial Z_t}{\partial u_{CPI,t}} \geq 0$	$\frac{\partial Z_t}{\partial u_{CPI,t}}$? (≤ 0)
$u_{call-ffr,t}$	$\frac{\partial Z_t}{\partial u_{call-ffr,t}} \leq 0$	$\frac{\partial Z_t}{\partial u_{call-ffr,t}} \leq 0$	$\frac{\partial Z_t}{\partial u_{call-ffr,t}} \geq 0$	$\frac{\partial Z_t}{\partial u_{call-ffr,t}}$? (≤ 0)
$u_{won/\$,t}$	$\frac{\partial Z_t}{\partial u_{won/\$,t}} \geq 0$	$\frac{\partial Z_t}{\partial u_{won/\$,t}} \geq 0$	$\frac{\partial Z_t}{\partial u_{won/\$,t}} \geq 0$	$\frac{\partial Z_t}{\partial u_{won/\$,t}} \geq 0$

Notes: 1) As indicated in parentheses in Table, Farrant and Peersman (2006) additionally use sign restrictions $\partial Z_{won/\$,t}/\ \partial u_{call-ffr,t} \leq 0$ (monetary tightening shock ↑ ⇒ real exchange rate ↓) and $\partial Z_{won/\$,t}/\ \partial u_{CPI,t} \leq 0$ (demand shock ↑ ⇒ real exchange rate ↓). However, this study uses the nominal exchange rate rather than the real exchange rate, and since the impact of a tightening policy shock on the exchange rate differs both theoretically and empirically, no sign constraints were imposed.

2) Farrant and Peersman (2006) classify $u_{IP,t}$, $u_{CPI,t}$, $u_{call-ffr,t}$, $u_{won/\$,t}$ as supply, demand, monetary, and exchange shocks, respectively, and use the real exchange rate.

Figure D1. Cumulative Impulse Response Results (Sign Restrictions Case)

Period & Model	*(call-ffr)* → *ΔIP*	*(call-ffr)* → *ΔCPI*	*(call-ffr)* → *Δwon/$*
’99.01–’08.08 Difference VAR(2)			
’14.01–’25.05 Difference VAR(2)			
’10.01–’25.05 Difference VAR(2)			
’11.01–’25.05 Difference VAR(2)			

Notes: 1) The solid line represents the median, while the dotted lines above and below the median indicate the 2.5th and 97.5th percentiles which are a concept distinct from confidence intervals.

2) The dot-and-dash line shows the estimates from the Cholesky decomposition.

APPENDIX E. Impulse Response Analysis with Additional Global Variables

In Appendix E, I examine the impulse response functions when global variables are added in addition to the dollar index and oil prices. The additional global variables include *ffr*, the Equity Market Volatility Tracker (EMVOVERALLEMV), 10-Year Treasury Constant Maturity Minus 2-Year Treasury Constant Maturity (T10Y2YM), 5-Year Treasury Constant Maturity Minus Federal Funds Rate (T5YFFM), 1-Year Treasury Constant Maturity Minus Federal Funds Rate (T1YFFM), 6-Month Treasury Constant Maturity Minus Federal Funds Rate (T6MFFM), and 3-Month Treasury Constant Maturity Minus Federal Funds Rate (T3MFFM). The Equity Market Volatility tracker tracks the VIX and the realized volatility of returns on the S&P 500. These volatility and term spread data are obtained from the Federal Reserve Economic Data (St. Louis Fed).

Figure E1 shows the impulse response functions when various variables, in addition to the dollar index and oil prices, are included as global shocks. To save space, only the period from January 2014 to May 2025 is shown. In all cases, an upward shock to (*call-ffr*) causes *IP*, *CPI*, and *won/$* to decline.

Figure E1. Cumulative Impulse Response Results
(Period: January 2014–May 2025, Model: Difference VAR(2))

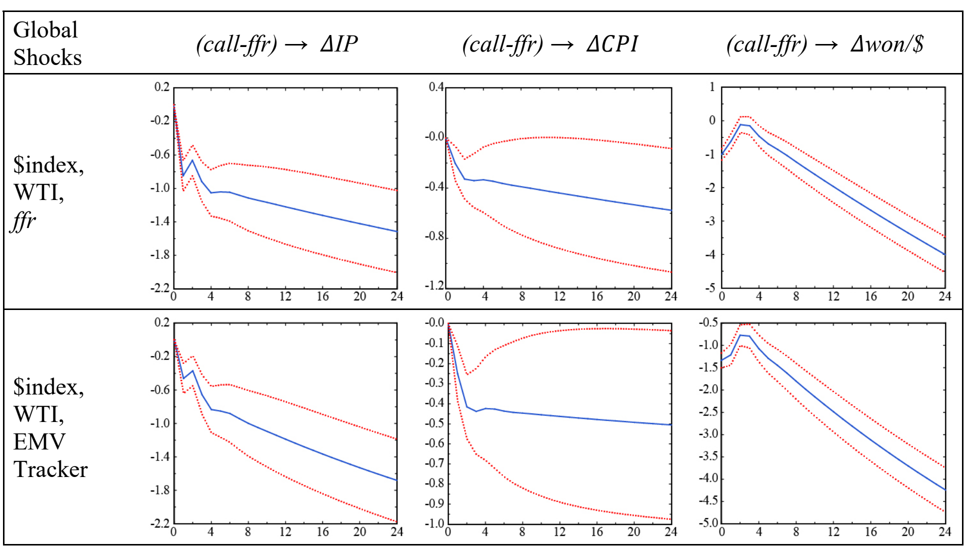
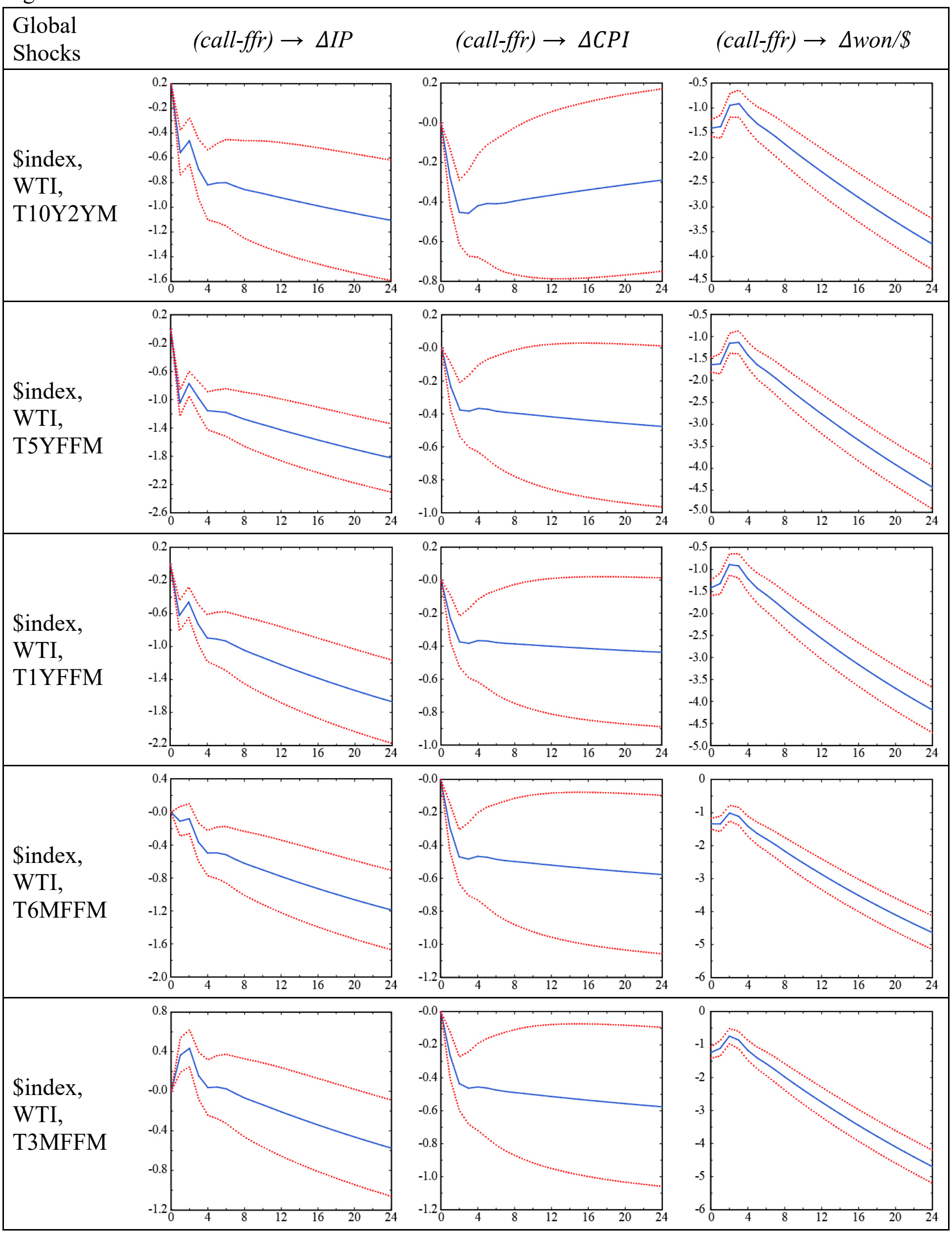


Figure E1. Continued



Notes: 1) EMV Tracker: Equity Market Volatility Tracker (EMVOVERALLEMV)

2) The dotted lines above and below the solid lines indicate the 95% confidence bands (response estimate ± 1.96×one standard deviation) obtained through 1,000 Monte Carlo simulations.

APPENDIX F. Comparison of Direct and Overall Effects

I consider the following simple structural two-variable VAR(1) model.

$$\Gamma_0 Z_t = a_0 + A_1 Z_{t-1} + u_t \tag{F1}$$

Since $Z_t = [Z_{1,t}, Z_{2,t}]'$ in Equation (F1) and the error vector $u_t = [u_{1,t}, u_{2,t}]'$ represent structural error vectors, I assume that the conditional expectations of their cross-products are also zero, regardless of the lag number. Matrix Γ_0 represents the contemporary relationship between the two variables and its diagonal elements are normalized to 1 as follows.

$$\Gamma_0 \equiv \begin{bmatrix} 1 & \gamma_{12} \\ \gamma_{21} & 1 \end{bmatrix} \tag{F2}$$

To estimate the model, I first derive a reduced form of Equation (F1), which can be obtained by multiplying Equation (F1) by Γ_0^{-1}, as follows.

$$\mathrm{Z}_t = b_0 + B_1 Z_{t-1} + \varepsilon_t \tag{F3}$$

In Equation (F3), $b_0 = \Gamma_0^{-1} a_0$ and $B_1 = \Gamma_0^{-1} A_1$ is given. The reduced-form error vector ε_t is expressed as follows.

$$\varepsilon_t = \begin{bmatrix} \varepsilon_{1,t} \\ \varepsilon_{2,t} \end{bmatrix} = \begin{bmatrix} 1 & \gamma_{12} \\ \gamma_{21} & 1 \end{bmatrix}^{-1} \begin{bmatrix} u_{1,t} \\ u_{2,t} \end{bmatrix} = \begin{bmatrix} \frac{1}{1-\gamma_{12}\gamma_{21}} & \frac{-\gamma_{12}}{1-\gamma_{12}\gamma_{21}} \\ \frac{-\gamma_{21}}{1-\gamma_{12}\gamma_{21}} & \frac{1}{1-\gamma_{12}\gamma_{21}} \end{bmatrix} \begin{bmatrix} u_{1,t} \\ u_{2,t} \end{bmatrix} = \Gamma_0^{-1} u_t \tag{F4}$$

Therefore, when the reduced-form covariance matrix and the structural covariance matrix are denoted by Σ and Ω, respectively, Σ and Ω have the following relationship.

$$\Sigma = \mathrm{E}(\varepsilon_t \varepsilon_t') = \mathrm{E}\left(\Gamma_0^{-1} u_t u_t' \Gamma_0^{-1\prime}\right) = \Gamma_0^{-1} \Omega \Gamma_0^{-1\prime} \tag{F5}$$

The covariance matrix in Equation (F5) can be expressed specifically as follows.

$$\begin{bmatrix} \Sigma_{11} & \Sigma_{12} \\ \Sigma_{21} & \Sigma_{22} \end{bmatrix} = \frac{1}{(1-\gamma_{12}\gamma_{21})^2} \begin{bmatrix} \Omega_{11} + \gamma_{12}^2 \Omega_{22} & -(\gamma_{21}\Omega_{11} + \gamma_{12}\Omega_{22}) \\ -(\gamma_{21}\Omega_{11} + \gamma_{12}\Omega_{22}) & \gamma_{21}^2 \Omega_{11} + \Omega_{22} \end{bmatrix} \quad \text{(F6)}$$

In the case of Equation (F6), since there are four unknowns $(\gamma_{12}, \gamma_{21}, \Omega_{11}, \Omega_{22})$ but only three equations, it is impossible to identify the parameters of the structural model from the reduced-form model. Therefore, as is well known, additional assumptions or constraints are required. For example, by assuming γ_{12}=0 in (F4) using the Cholesky decomposition—a triangular matrix decomposition method—the impulse response function can be easily obtained by identifying γ_{21}, Ω_{11}, and Ω_{22}. Another example, as discussed in Appendix D, involves imposing a constraint based on the theoretical model on the sign of the parameter Γ_0 at the same time period to identify the range within which the parameter lies. To examine the direct and overall response effects of the shock, Equation (F1) can be expressed more specifically as follows.

$$Z_{1,t} = -\gamma_{12} Z_{2.t} + a_{01} + a_{11} Z_{1,t-1} + a_{12} Z_{2,t-1} + u_{1,t} \quad \text{(F7)}$$

$$Z_{2,t} = -\gamma_{21} Z_{1.t} + a_{02} + a_{21} Z_{1,t-1} + a_{22} Z_{2,t-1} + u_{2,t} \quad \text{(F8)}$$

In the above equation, a 1%p increase in Z_2 results in a direct change of $-\gamma_{12}$%p in Z_1, while a 1%p increase in Z_1 results in a direct change of $-\gamma_{21}$%p in Z_2. Meanwhile, parameter Γ_0 in the structural model indicates a direct causal relationship between Z_1 and Z_2, whereas parameter Γ_0^{-1} in the reduced-form model reflects the overall effect, including not only direct causality but also indirect causality mediated through other variables.

$$Z_{1,t} = b_{01} + b_{11} Z_{1,t-1} + b_{12} Z_{2,t-1} + \frac{1}{1-\gamma_{12}\gamma_{21}} u_{1,t} - \frac{\gamma_{12}}{1-\gamma_{12}\gamma_{21}} u_{2,t} \quad \text{(F9)}$$

$$Z_{2,t} = b_{02} + b_{21} Z_{1,t-1} + b_{22} Z_{2,t-1} - \frac{\gamma_{21}}{1-\gamma_{12}\gamma_{21}} u_{1,t} + \frac{1}{1-\gamma_{12}\gamma_{21}} u_{2,t} \quad \text{(F10)}$$

As can be seen in Equation (F10), a 1%p increase in Z_1 generally causes Z_2 to change by $-\frac{\gamma_{21}}{1-\gamma_{12}\gamma_{21}}$ %p. To be more specific, in Equation (F8), a 1%p increase in Z_1 causes Z_2 to change by $-\gamma_{21}$%p, whereas a $-\gamma_{21}$%p in Z_2 causes Z_1 to change by $(-\gamma_{12}) \times (-\gamma_{21})$%p, as shown in Equation (F7). This, in turn, causes Z_2 to change by $(-\gamma_{21}) \times (-\gamma_{12}) \times (-\gamma_{21})$ %p, as shown in Equation (F8). If this process is repeated, the final result is that when Z_1 increases by 1%p, Z_2 changes by $-\frac{\gamma_{21}}{1-\gamma_{12}\gamma_{21}}$ %p. Meanwhile, Equation (F9) shows that a 1%p upward shock to Z_2 generally causes Z_1 to change by $-\frac{\gamma_{12}}{1-\gamma_{12}\gamma_{21}}$ %p (the case of sign restrictions). However, assuming Cholesky decomposition, γ_{12}=0, so in this case, it does not affect Z_1 at the same time, and a 1%p upward shock to Z_1 also causes Z_2 to change by only $-\gamma_{12}$ %p overall. Although the case with four variables is somewhat more complex, it can be derived using the same principle.