Output growth volatility and inflation uncertainty:
Empirical evidence from East Asian Economies

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Abstract

This paper examines the dynamic linkages between output growth volatility and inflation uncertainty for three East Asian economies by employing different originated time span and monthly time series data from January 1990 onwards for China, from January 1957 for Japan, and from January 1975 for South Korea, through December 2016. The empirical evidence relies on a bivariate version of an asymmetric and unrestricted VAR(p)–GARCH–M–BEKK econometric approach and reveals that there are strong and noteworthy relationships between the series of concern. First, output growth volatility has a significant adverse impact on the inflation uncertainty of China's economy but a significant and positive impact in the economies of Japan and South Korea. Furthermore, our series of concern are also applied for variance non–causality test and generalized impulse response analysis of output growth volatility to inflation uncertainty under a vector autoregression process as well. Here, the variance non–causality test results support that there is strong evidence of a bi–directional causal effect from the real output growth volatility to inflation uncertainty of China's economy; on the other hand, a uni–directional impact is found for Japan and South Korea's economies. The analytical framework of impulse response analysis shows that output growth volatility has significant affirmative influence in response to the shocks of inflation uncertainty on China and South Korea's economies but only an ambiguous impact on Japan's economy.

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1. Introduction

Almost a decade has passed since the global financial crisis of 2008, and the nexus between output growth volatility and inflation uncertainty performances is controversial to monetarists and academics. In fact, since the 1980's, this relationship has underpinned growing research interests among many others. Following the Friedman (1977) Nobel lecture, a number of theoretical and empirical works have proposed sound effects of output growth volatility on inflation uncertainty (see, e.g., Devereux, 1989; Blackburn, 1999; Fountas et al., 2002; Cuikerman and Gelrach, 2003; Grier et al., 2004; Bredin and Fountas, 2009; Li and Kwok, 2009; inter alia). Focusing on the output growth volatility three economically–giant and crisis–affected East Asian economies, the recent relative paper by Li and Kwok (2009) argues that declines in output growth volatility, especially in developing countries, is due to the recent decade of declining inflation and several structural reforms. Also, the outward–oriented economy tends to cause more external shock spillover and impulsiveness in output growth. However, Easterly et al. (2001) argue that output growth volatility can fall at both the micro and macro level.

Evidently, other world economies have felt the impact of the recent economic and financial slumps, such as Asian financial crisis (hereinafter, AFC) during 1997–1998 and global financial crisis (henceforth, GFC) of 2008 (the largest financial slump since the Great Depression of 1929–1933). East Asian economies, namely China, Japan and South Korea, are also adequately stung by these financial slumps. Focusing on these crisis–affected economies has the foremost impetus in that they have expressed cognate economic structure and performance over the past few decades, such as a rapid industrialization in the Post–World War II era. They have also suffered from different degrees of setback, especially during the AFC and GFC (Li and Kwok, 2009). Given their structural transformation over the last few decades from the secondary to tertiary industry, their economic performance has become more stable. However, their economic openness and strengths in the financial market have incited them from this steady growth. Throughout the following subsection, the distinct features of the economies of concern responses on the recent decade of economic and financial downturns will be concisely observed.

Mainland China

Prior to 1979, mainland China was known as a centrally–planned economy that was not highly integrated with the rest of the world. After starting economic reforms and adapting policy to open up trade, China has successfully transformed its own economy into an outward–oriented market economy within a few decades and is on track to become the second largest economy following that of the US. Meanwhile, its admission into the World Trade Organization has also enhanced its integration within the multinational trade system and led to considerable growth in industrial output (Caporale et al., 2015). On the other hand, its multilateral economic integration contributed to increase in output growth and employment, which facilitated other comparative advantages and an increasing specialization level. A relatively recent paper by Tseng and Zebregs (2002) links China's thriving output growth with economic openness and foreign direct investment (henceforth, FDI). At the outset of opening–up trade policy adaptation, the government massively orientated external investment towards agriculture and industrial output as well as service sectors in coastal areas in order to boost the establishment of a modern enterprise system. As a result, the economic growth of the country was prophesied at around ten percent per annum.

Since China was highly integrated through its financial activities, the recent decade of shocks are often attributed to economic and financial crunches that were suitably ensued. In turn, a number of relatively recent papers embody that notwithstanding unprecedented economic growth and structural modifications, China performed rapid fluctuations in its inflation trend. Early reforms in 1979 included administrative price adjustments that were continued until 1990, and the government took advantage of the relative stability of the overall price level to undertake major price reforms. In other sectors, the government liberalized industrial output prices from the state–determined levels. Consequently, two double–digit inflationary spells occurred, and the inflation rate has exhibited conspicuous patterns since then. The price controls were also reduced significantly in light and heavy industries, whereas they remained in the 1990's, when industrial output mainly geared towards export. In the end, it has exhibited more stability than other East Asian economies with regard to the recent decade's economic slumps.

Japan

Indubitably, among the East Asian economies, Japan has unique originality in its growth in that so far it has successfully managed to become the only great nation from whole Asia. Overall, the successes and failures of Japan's growth can be split into three stages: (1) the catch–up process during the post–war era from 1945 to 1960's, (2) failure to reform and a bubble economy during the 1970's and 1980's, and (3) long stagnation periods and beyond from the 1990's to the present day. After WWII and in the 1960s, the common shared purpose was to catch–up with North America and other highly industrialized Western European economies' experience with the governmental sectors and households. During this period, the government created a system that was able to mobilize and direct funds to key industries for rapid economic growth. After achievement of the common goal, business and household sectors should have changed their behavior from collective actions into autonomous under a more competitive environment.

Moreover, during the recent decade of economic and financial instabilities, Japan's economy experienced a structural break in the 1980's, where the financial sector fell into a liquidity trap in the early 1990's, which caused an increase in output growth volatility. During the liquidity trap period, the GDP growth of the economy was negative, and it led to increase in government spending and debts in order to speed up the economy. In the 1990's, Japan also suffered from a deflationary series which was associated with low or zero inflation and low output growth rates. The AFC could not cause an increase in output volatility of the economy; the main reason was the high accessibility of the cyclical movement between investment and output. However, the stock market volatility increased faintly, but remained the lowest compared to other crisis–effected East Asian economies under study. In a similar vein, focusing on the linkages between output growth volatility and inflation uncertainty of Japanese economy, Fountas et al. (2002) found empirical evidence that the higher inflation rate and inflation uncertainty lead to lower output growth in the economy. However, relying on Friedman's (1977) postulation, we are interested in finding out whether the volatility of output growth can be associated with corresponding inflation uncertainty of the economy and its causal effects, as well as impulse response to inflation uncertainty.

South Korea

Similar to other East Asian economies, South Korea has indicated remarkable performance since the early 1960's, growing from a lower–middle income economy into a top ten nation. Until the end of 1970's, the government massively intervened in domestic investments along the lines of comparative advantage in output industry and foreign trade. In particular, tax exemptions were provided to labor–intensive manufacturing industries to enhance export, and high tariffs were imposed to preserve local manufacturers. Besides, South Korea's economy has also exhibited impressive growth in industrial output, surpassing around seven percent per annum. However, this progress was interrupted by the shocks brought with the recent decade's economic and financial slowdowns. Since the government has a crucial position in promoting export–oriented output, the economy has been recovering rapidly and much faster than was expected. Overall, South Korea's economic growth originality resembles that of Japan. Its historical growth and catch–up to the advanced economies have enormously attributed to favorable growth factors.

Relatively, maintenance of good institutions, high investment rates and trade openness as well as financial market activities and strong human capital supply, especially in technological process and adoption of inflation targeting policy, have contributed to strong economic growth over the few decades. However, it should be accentuated that owing to shockwaves from the AFC in 1998, the South Korea economy has adequately smarted and annual growth change was measured negatively (–5.5 % in 1998). Around the catch–up policy, it launched an inflation targeting regime in 1999 for rapid recovering the economy. During this period, the optimal policy geared towards setting more stability in domestic prices. The launch of the regime was preceded by the significant shocks of macroeconomic and financial instability. Notwithstanding, during the inflation targeting period, the persistence of inflation remained relatively high while that of output measures declined considerably. As noted by Li and Kwok (2009) for the post–AFC era, the South Korea shows negative output volatility; however, it experienced a relatively large decline in export volatility while import growth and its stock price volatility have drastically increased. In reducing volatilities relying on foreign direct investment, the budget was restructured.

Output and inflation growth performances of three East Asian economies

Li and Kwok (2009) further stress that after ensuring a structural break in the 1980's and a decline in the early 1990's, Japan fell into the liquidity trap and output growth volatility sufficiently amplified. A similar scenario also supervened in South Korea's economy, and so far, the largest negative annual growth rate aftershocks were brought about by the AFC from 1997–1998.

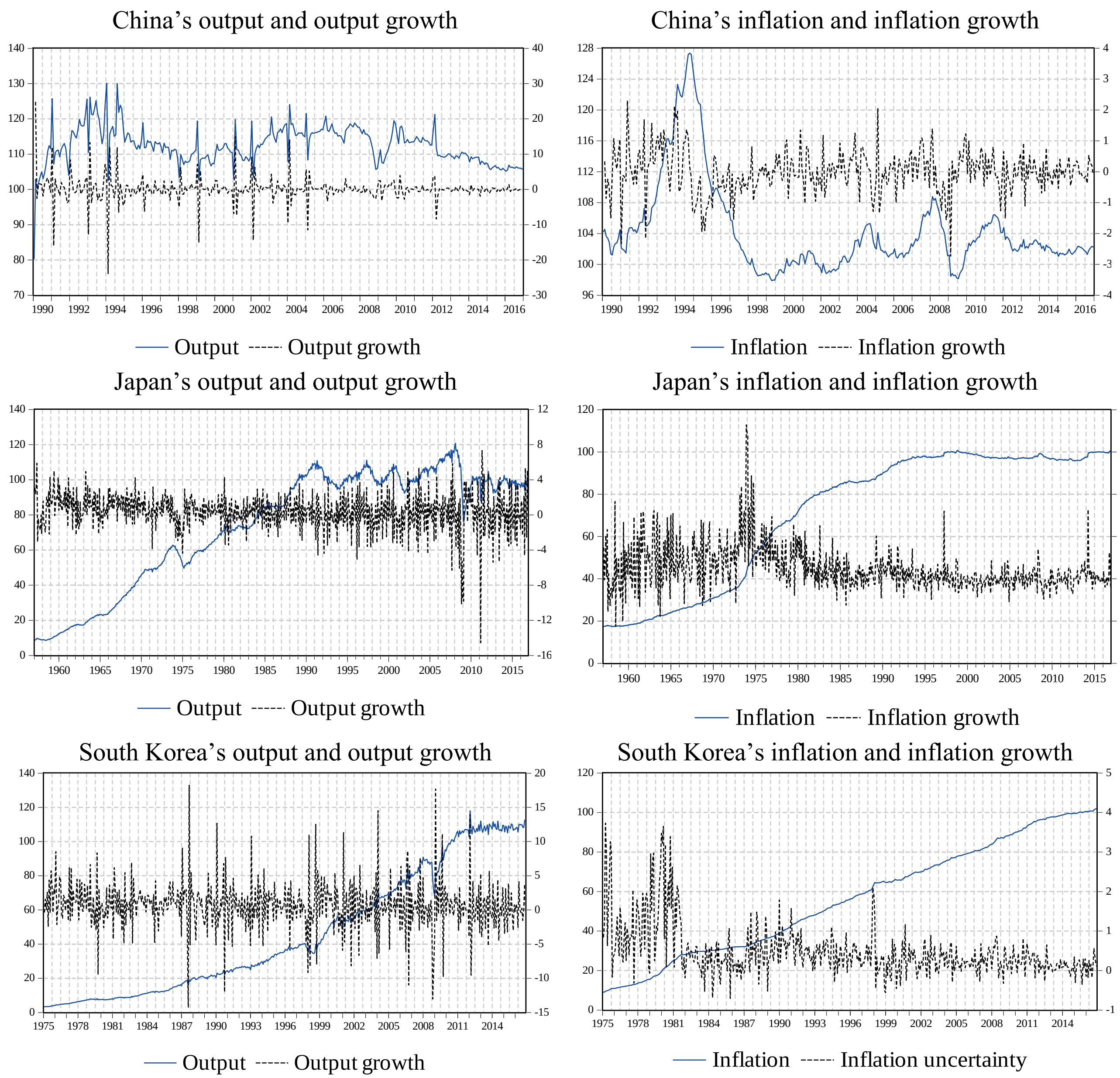

Fig. 1. The data and growth of the series for the sample periods.

Fig. 1 depicts graphical illustrations of seasonally adjusted monthly time series and log–transformed data on the index of industrial production (henceforward, IIP) and consumer price index (hereinafter, CPI) as a proxies of output volatility and performance of inflation uncertainty for three East Asian economies. As plotted, the output growth of Japan is highly volatile for the sample period, and a structural break is also ensured in 2009 and 2011, whereas in China and South Korea economies the output growth series exhibit relatively high volatility for the sample periods of 1985–1990, 1990–1994, 1998–2004 and 2008–2012. There is a high possibility that these volatile sample periods are relatively associated with the recent decade's economic and financial downturns.

In the instance of China's economy, after the 1990's the real industrial output growth created a chance for high integration with the rest of the globe, and international trade improved its economic potential. This led to faster technological progress in industrial output sector. Understanding the prominence of China's recent output growth volatility and its impact on other macroeconomic variables have attracted considerable attention (see, e.g., Laurenceson and Rodgers, 2010; He and Chen, 2014; Caporale et al., 2015; inter alia). The study by Caporale et al. (2015) also argues that similar to other world economies, the recent decade of economic unsteadiness also distressed China's economy. As a result, the volatilities through international trade spilled over to the industrial sector. An exception is He and Chen (2014), who provide the potential sources of China's macroeconomic moderation through three possible explanations: good policy, good practice and good luck. Since the global economic condition continued to deteriorate at the end of 2014, GDP growth was computed at more than seven percent, and hitherto the low levels have been continued.

Undoubtedly, the effects of output growth volatility on inflation uncertainty have been widely studied, and its upshots are different. Following Grier et al. (2004) this paper specifies and estimates an extremely general model of output growth volatility on inflation uncertainty. Unlike earlier studies in this context, our model allows for the possibilities of spillover effects and asymmetries in the variance–covariance structure of output and inflation growth series. By extending the Barro–Gordon model to introduce wage indexation endogenously, Devereux (1989) examined the precise relationship between output growth volatility and inflation uncertainty nexus. According to his research, an increase in the real output growth volatility exogenously impacted the degree of wage indexation and, consequently, the optimal inflation rate delivered by policymakers. Similar to the theoretical approach, Devereux's hypothesis argues that more output volatility reduces the inflation rate as well as inflation uncertainty.

In a similar way, Bredin and Fountas (2009) also show that relying on the Cukierman–Meltzer hypothesis, higher real output growth volatility reduces the nominal uncertainty and inflation rate. Based on the stabilization motives of the monetary authority, Holland (1995) argues that when inflation uncertainty emerges owing to increasing inflation, the monetary authority should construct money supply growth to eliminate it. As noted by Karahan (2012), normally Central Banks prefer to keep inflation at low levels and to increase output by making monetary surprises. Thus, due to vague monetary control mechanisms, the money supply process is considered to be random. Friedman (1977) explains that the emergence of higher inflation uncertainty distorts the effectiveness of the price mechanism in allocating resources efficiently, and thus it adversely affects output growth. According to his hypothesis, as higher output growth leads more inflation, inflation uncertainty will also increase.

By focusing on the effects of the recent decade economic and financial recessions on five crisis–affected East Asian economies, the study by Li and Kwok (2009) sum up that in the post–AFC era, the output growth volatility increased in South Korea. Other relative studies on this context argue that a strong performance in financial market activities and the degree of openness of the economy tend to face more external shocks, and output volatility can remain high. However, over the past few decades, sufficient attention has been concentrated on output volatility and stability of the economies under investigation, but recently the output volatility still remains a major factor of uncertainty. Due to the divergent features on the financial crisis, every economy responds differently in each economic and financial crisis series. Specifically, since the AFC in 1997–1998, the economies of Japan and South Korea have been suffering from banking constraints and with increasing size of non–performing loans and debts. Cherry (2007) argues that South Korea, after devaluation of the national currency, followed by the massive injection of foreign direct investment, adopted financial and budgetary restructuring, while Japan instead looked for a domestic solution, tried to minimize external influence, and ensured a decline in output volatility.

To elucidate the nexus of concern, the current study proposes three research questions. First, we inspect the impact of output growth volatility on performance of inflation uncertainty for selected economies of concern. Next, we conduct variance non–causality test if there is any shock or volatility spillover influence between the variables and if the estimated growth series react to each other's on cross–variables correlation analysis. Lastly, we check for an analytical framework of generalized impulse responses of output growth volatility to inflation uncertainty shocks under a vector autoregression process to inspect a unit (one standard deviation) shock of real volatility on nominal uncertainty.

The reminder of this paper is organized as follows. Section 2 introduces a concise depiction of output and inflation growth statistical trends and preliminary data analysis of the respective economies for the sample periods, paying particular attention to cyclical macroeconomic performance. Section 3 outlines econometric approaches that rely on concerned specific research questions, namely multivariate GARCH–in–mean model formulation, variance non–causality analysis and an analytical framework of impulsive response of output growth volatility to inflation uncertainty under a vector autoregression process. Section 4 discusses the empirical results, which are drawn from the model assessment, specifically the exposure of nominal uncertainty owing to real output volatility, asymmetry and cross–correlation, mutual causality relationships and generalized impulse response function analysis of the series. Finally, section 5 concludes the study and provides with some policy implications for the sample periods of the economies.

2. Data and summary statistics

We employ monthly historical time series data on IIP and CPI of the People's Republic of China (hereinafter, China) excluding Hong Kong and Macau, Japan and the Republic of Korea (henceforward, South Korea) that are retrieved from their respective government sources through DataStream online services[2]. The lattermost months of both series are set on December 2016, while the beginning period started from January 1990 onwards for China, from January 1957 onwards for Japan, and from January 1975 onwards for South Korea, respectively. The data scale of China's indices are defined in a corresponding period of the previous year that equals to one hundred (CPPY=100), while Japan and South Korea's growth series are a constant one hundred for 2010 and 2015 (2010=100 and 2015=100). Following recent uncertainty and volatility related works (see, e.g., Fountas et al., 2002; Grier et al., 2004; Bredin and Fountas, 2009; inter alia), the employed data was seasonally adjusted to allow for smoothening of any seasonal impacts and to lessen the noise in model. However, it should be pointed out that the real industrial output series reflects only manufacturing, mining and utilities of the respective economy.

Table 1
Summary statistics for output and inflation growth.

	China		Japan		South Korea	
	$R_{y,t}$	$R_{\pi,t}$	$R_{y,t}$	$R_{\pi,t}$	$R_{y,t}$	$R_{\pi,t}$
Panel A: Descriptive statistics						
Mean	0.0852	– 0.0058	0.3434	0.2438	0.6947	0.4833
Std. Dev.	3.7872	0.7081	2.1105	0.5203	3.6412	0.6645
Skewness	0.0212	– 0.1355	– 1.0239	1.8395	0.2976	2.1748
Kurtosis	17.261	4.4275	8.2524	9.8237	7.0219	9.1052
B–J	2737.4***	28.415***	952.13***	1800.4***	346.44***	1177.7***
Panel B: Serial correlation and ARCH test						
Q(4)	57.120***	44.945***	85.913***	325.64***	107.50***	410.08***
Q(8)	59.717***	89.387***	133.89***	640.68***	127.37***	745.24***
$Q^2(4)$	15.117***	21.133***	95.843***	391.63***	54.257***	380.36***
$Q^2(8)$	15.966**	26.681***	109.24***	626.14***	69.186***	631.05***
BDS(8)	0.1183***	0.0599***	0.0514***	0.1495***	0.0536***	0.1787***
ARCH(4)	10.041***	4.5780***	19.337***	77.878***	15.838***	68.525***
Panel C: Unit root and stationarity tests						
PP(μ)	– 44.051***	– 15.003***	– 31.730***	– 26.778***	– 36.397***	– 15.067***
PP(τ)	– 45.939***	– 14.984***	– 31.847***	– 26.998***	– 36.567***	– 16.936***
ADF(μ)	– 16.757***	– 5.6960***	– 6.5596***	– 3.4899***	– 21.346***	– 3.2415**
ADF(τ)	– 7.6446***	– 5.6973***	– 7.8686***	– 4.5984***	– 21.884***	– 3.5084**
KPSS(μ)	0.3668	0.0621	1.3984	1.5827	0.7948	1.4603
KPSS(τ)	0.1169	0.0628	0.0874	0.2067	0.0329	0.2016
N obs.	324	324	720	720	504	504

Notes: ***, **, * indicate 1%, 5% and 10% significance level, respectively. The data are presented as seasonally adjusted. Here, $R_{y,t}$ and $R_{\pi,t}$ denote log changes of output and inflation growth series, respectively. μ is the inclusion of an intercept without time trend, while τ refers to an intercept with time trend at this juncture for unit root tests. The critical values of the KPSS(μ) and KPSS(τ) unit root tests at 5% significance level are 0.463 and 0.146, respectively.

[2]The output and inflation series were retrieved for China from the National Bureau of Statistics of China database, for Japan from he Statistics Bureau of Japan database and for South Korea from the Statistics Korea database, without any seasonal adjustments through DataStream online services. Note here that the statistical data series on index of China's real industrial production are repeatedly not published for the first months of 2006 through 2013 in National Bureau of Statistics of China database, and so the series are amended for estimation purposes. However, the derivation of data is available from the author upon request.

Table 1 reports summary statistics of the series under concern, from which $R_{i,t}$ is constructed by using monthly logarithmic first differences of $r_{i,t}$ as follows

$$R_{i,t} = log\left(\frac{r_{i,t}}{r_{i,t-1}}\right) \times 100 \qquad (1)$$

where, subscript i denotes real IIP ($R_{y,t}$) and nominal CPI ($R_{\pi,t}$) growth series of the respective economies under concern. According to the entries of Panel A, the means of monthly returns are smaller compared to their computed standard deviations (variances) in all cases. As reported, all returns of the growth series take have positive mean values (except the returns of China's inflation growth series). It expresses that the level of indices of real industrial output growth series in focusing on selected three East Asian economies reflect the corroborating evidence of higher varying degrees of volatility rather than their levels. On the contrary, Japan's real output growth series show evidence of being the steadiest among the three, followed by the economies of South Korea (3.6412 percent) and China (3.7872 percent). The expression of highest volatility on China's real output growth is probably due to recent decade fluctuations in international trade of the economy.

In the case of Japan's economy, the fact that it has the lowest standard deviation (2.1105 percent) for the sample period suggests that the impacts of output growth do less to impact inflation growth series. However, Japan's economy has experienced a relatively high decrease in inflation growth (0.5203 percent) compared to the economies of South Korea (0.6645 percent) and China (0.7081 percent), despite the severe impacts of the recent decade's economic and financial downturns. The unconditional distributions of China's nominal inflation growth and the real output growth of Japan are negatively skewed, while other series under investigation are positively skewed and display significant amounts of excess kurtosis. Here, the existence of excess kurtosis indicates a leptokurtic distribution of the return of the series. Besides, the lacking of symmetry and the significant amounts of excess kurtosis (i.e., fat–tailedness) suggests that the growth series are conditionally heteroscedastic with failing into consent of the null hypothesis of Bera–Jarque (henceforward, B–J) test from Bera and Jarque (1980) for normality.

In Panel B of Table 1, the Ljung–Box Q test statistics from Ljung and Box (1979) for serial correlation of the growth and the squared growth series of $R_{y,t}$ and $R_{\pi,t}$ are thoroughly detailed. Here, the statistics Q and Q^2 are asymptotically distributed as Chi–squared (χ^2) values, with their degree of freedom on the null hypothesis having no autocorrelation. Apparently, the underlying growth series exhibit a significant amount of linear dependence and strong evidence of conditional heteroscedastic effects in growth series, suggesting the existence of an autoregressive structure in both mean and volatility of the returns. Besides, Brock, Dechert and Scheinkmen (hereinafter, BDS) test statistics for independence proposed by Brock et al. (1996) indicate that all growth series are not independently and identically distributed (iid) at the one percent significance level. For the sample size considered in the study, both output and inflation growth series display strong conditional heteroscedasticity; thus, the multivariate GARCH–in–mean model is exceedingly suitable for model running. However, the GARCH family models often require that the underlining data be applied for stationarity properties in order to avoid from the non–sense regression problem in model assessment of underlying data.

As mentioned above, prior to running the model, some usual unit root and stationarity tests, such as the augmented Dickey–Fuller (hereinafter, ADF) test from Dickey and Fuller (1979), the Phillips–Perron (henceforth, PP) test from Phillips and Perron (1988) and the Kwiatkowski–Phillips–Schmidt–Shin (henceforward, KPSS) test from Kwiatkowski et al. (1992) are conducted to test for the existence of unit root and stationarity in the series of concern. The results are reported in Panel C of Table 1. The optimal lags for the ADF test are selected using Schwarz information criterion (hereinafter, SIC), and the bandwidth for PP and KPSS tests are selected with Newey–West by using the Bartlett kernel. To select the optimal lag length in ADF and PP tests, the Schwarz Bayesian criterion is used. Likewise, in Grier et al. (2004) and Mahadevan and Suardi (2011), the null hypothesis of the latter test for stationarity is opposed to the null of a unit root in ADF and PP tests. The robustness of the sample series is established by holding it in level with incorporating an intercept, μ, and an intercept and time trend, τ. Referring to the unit root and stationarity tests' results, the ADF and PP tests display that all the series under concern are stationary.

Meanwhile, the KPSS test statistics also support this inference. All in all, the tests reject the null hypothesis of the existence of unit root at the one percent significance level, and thus the returns follow a stationary process regardless of whether a trend variable or/and incorporated in the model. To sum up, all the growth series under concern are treated as $I(0)$ process.

3. A bivariate GARCH–in–mean model of output and inflation growth

Prior to running the model, it is expediently proposed to elucidate a general multivariate GARCH–in–mean model that employs recent data information on uncertainty and volatility related works. Indeed, the GARCH family models are highly functional in financial issue related works owing to their high volatility. As noted by Bauwens et al. (2006), these models are conventionally applied to describe and forecast the correlations between the volatilities and co–volatilities as well as spillover effects directly (through the conditional variances) or indirectly (through the conditional covariances). Moreover, a number of definitions in the literature make standard multivariate GARCH models possible. Specifically, in this work, we use the model depiction, which is explained by Bauwens et al. (2006) as follows

$$y_t = x_t + u_t \tag{2}$$

where, y_t is a vector stochastic process with dimension $N \times 1$ such that $E(y_t) = 0$. The information set is denoted by Ω_{t-1} and generated with the past information of y_t at the time $t-1$. x_t is the conditional mean vector and the residual terms, u_t is specified as

$$u_t = \sqrt{H_t}\eta_t \tag{3}$$

where, Ω_{t-1} is the given information set at the time, $t-1$, $H_t = h_{ij,t}$ is the conditional covariance matrix of y_t with dimension $N \times N$, and η_t is an iid vector error process such that $E\eta_t\eta_t' = I$. There is no linear dependence structure in a vector stochastic error process, whereas in multivariate time series applications, the residual term, u_t is normally considered as a vector of logarithmic returns of N and denotes indices, prices and assets, etc.

Despite the existence of numerous definitions of the standard multivariate GARCH–in–mean model in the literature, this study closely follows several recent papers (see, e.g., Kim, 2000; Grier et al., 2004; Bauwens et al., 2006; Li and Kwok, 2009; Bredin and Fountas, 2009; Mahadevan and Suardi, 2011; inter alia). In a model assessment, a bivariate version of an asymmetric and unrestricted econometric approach labeled as VAR(p)–BGARCH–M–BEKK is applied on log–transformed output and inflation growth series of the selected three East Asian economies. The proxies of real output growth volatility and inflation (nominal) uncertainty are measured by the conditional variance of output $R_{y,t}$ and inflation $R_{\pi,t}$ growth, which are generated by a fitted bivariate GARCH–in–mean process. The conditional standard deviations of growth series are included as the explanatory variables in the conditional mean equation. Thus, the conditional mean and conditional variance–covariances matrices in the current work are being estimated jointly within a single system.

Moreover, we exploit five widely–used criteria which are given in Lütkepohl (2005), namely: the Akaike information criterion of Akaike (1973) (hereinafter, AIC), the Schwarz Bayesian criterion (henceforth, SBC), the Hannan–Quinn criterion (henceforward, HQC), the final prediction error (hereafter, FPE), and the log–likelihood (henceforth, LL) value. These criteria are used to select the optimal lag length in the employed model. For checking out the adequacy of the estimated model, the diagnostic checks and specification tests no GARCH, no GARCH–in–mean, diagonal GARCH, diagonal VAR–in–mean and non–asymmetry are considered. However, note here that the non–diagonal covariance process is allowed in model estimation to inspect whether the volatility in one series spills over into the volatility of another growth series of concern. By setting the number of explanatory variables as equal to two $(n = 2)$ and with strong evidence of conditional heteroscedasticity in output $R_{y,t}$ and inflation $R_{\pi,t}$ growths of selected economies, the conditional mean process of the study is specified as follows

$$Y_t = M + \sum_{i=1}^{p} \Gamma^{(i)} Y_{t-i} + \Psi\sqrt{H_t} + U_t; \tag{4}$$

$$U_t|\Omega_{t-1} \sim (0, H_t), \;\; H_t = \begin{bmatrix} h_{y,t} & h_{y\pi,t} \\ h_{\pi y,t} & h_{\pi,t} \end{bmatrix};$$

where, Ω_{t-1} is an available information set at the time $t-1$, 0 is the null vector, and

$$Y_t = \begin{bmatrix} y_t \\ \pi_t \end{bmatrix}; \; M = \begin{bmatrix} \mu_y \\ \mu_\pi \end{bmatrix}; \; \Gamma^{(i)} = \begin{bmatrix} \gamma_{11}^{(i)} & \gamma_{12}^{(i)} \\ \gamma_{21}^{(i)} & \gamma_{22}^{(i)} \end{bmatrix}; \; Y_{t-i} = \begin{bmatrix} y_{t-i} \\ \pi_{t-i} \end{bmatrix};$$

$$\Psi = \begin{bmatrix} \psi_{11} & \psi_{12} \\ \psi_{21} & \psi_{22} \end{bmatrix}; \; H_t = \begin{bmatrix} h_{y,t} \\ h_{\pi,t} \end{bmatrix}; \; U_t = \begin{bmatrix} u_{y,t} \\ u_{\pi,t} \end{bmatrix};$$

or as a two–dimensional matrix form, Eq. (4) can be rewritten as follows

$$\begin{bmatrix} y_t \\ \pi_t \end{bmatrix} = \begin{bmatrix} \mu_y \\ \mu_\pi \end{bmatrix} + \begin{bmatrix} \gamma_{11}^{(i)} & \gamma_{12}^{(i)} \\ \gamma_{21}^{(i)} & \gamma_{22}^{(i)} \end{bmatrix} \begin{bmatrix} y_{t-i} \\ \pi_{t-i} \end{bmatrix} + \begin{bmatrix} \psi_{11} & \psi_{12} \\ \psi_{21} & \psi_{22} \end{bmatrix} \begin{bmatrix} h_{y,t} \\ h_{\pi,t} \end{bmatrix} + \begin{bmatrix} u_{y,t} \\ u_{\pi,t} \end{bmatrix} \tag{5}$$

where, from Eq. (5), the discrete equations can be drawn as follows

$$\begin{cases} y_t = \mu_y + \gamma_{11}^{(i)} y_{t-i} + \gamma_{12}^{(i)} \pi_{t-i} + \psi_{11} h_{y,t} + \psi_{12} h_{\pi,t} + u_{y,t} \\ \pi_t = \mu_\pi + \gamma_{21}^{(i)} y_{t-i} + \gamma_{22}^{(i)} \pi_{t-i} + \psi_{21} h_{y,t} + \psi_{22} h_{\pi,t} + u_{\pi,t} \end{cases} \tag{6}$$

Eq. (4) expresses the first order moment given according to an i–variable VAR process, where the stochastic error term, U_t, is assumed to follow a normal distribution with mean zero and variance H_t, while Ω_{t-1} is the available information set at the time, $t-1$. The conditional variance of output and inflation growth are generated from second moment equations that have been simultaneously incorporated in the conditional mean to assess the sound effects of output growth volatility (real) on inflation (nominal) uncertainty for the selected economies under concern. Despite the existence of many proposed model specifications, an asymmetric version of BEKK variance–covariance parameterization of Grier et al. (2004) is used in a quadratic form to ensure for the positive definiteness as follows[3]

$$H_t = C'C + A'u_{t-1}u'_{t-1}A + B'H_{t-1}B + D'\zeta_{t-1}\zeta'_{t-1}D \tag{7}$$

[3] The acronym BEKK stands for the scholars Yoshi Baba, Robert Engle, Dennis Kraft, and Ken Kroner.

or the matrix form of the variance–covariance specification given as follows

$$C = \begin{bmatrix} c_{11} & 0 \\ c_{21} & c_{22} \end{bmatrix};\ A = \begin{bmatrix} a_{11} & a_{12} \\ a_{21} & a_{22} \end{bmatrix};\ B = \begin{bmatrix} b_{11} & b_{12} \\ b_{21} & b_{22} \end{bmatrix}; D = \begin{bmatrix} d_{11} & d_{12} \\ d_{21} & d_{22} \end{bmatrix};\ \zeta_{t-1} = \begin{bmatrix} \zeta_{y,t-1} \\ \zeta_{\pi,t-1} \end{bmatrix};$$

where, C is an upper triangular matrix to ensure for the positive definiteness of H_t, whereas A, B and D are 2×2 dimensional matrices that capture the lagged conditional standard deviations and covariances, H_{t-1}, as well as the past values of $u_{t-1}u'_{t-1}$ and $\zeta_{t-1}\zeta'_{t-1}$ in joint estimations of contemporaneous volatility of output growth and inflation uncertainty. Importantly, this parameterization contains a possible asymmetry, which is captured by the matrix D.

As proposed by Grier et al. (2004), a term $\zeta_{t-1}\zeta'_{t-1}$ is also used in this study to account for the potential asymmetric responses. More specifically, if the output growth volatility and inflation uncertainty are lower than their expected levels, it is generally treated as negative innovations or bad news regarding the changes of the growth. Hence, the variables $\zeta_{y,t}$, and $\zeta_{\pi,t}$ are defined as min $\{u_{y,t}, 0\}$ and min $\{u_{\pi,t}, 0\}$, which are negative residuals or bad news about the output and inflation growth, respectively. The bivariate version of an asymmetric and unrestricted VAR(p)–GARCH–M–BEKK econometric approach is estimated by using maximum log–likelihood function, $L_T(\theta, \eta)$, for T observations (from Table 1, for the economies in the current work, $T = 324$ for both series of China's, $T = 720$ for Japan and $T = 504$ for South Korea). This is explained in Bauwens et al. (2006) as follows

$$L_T(\theta, \eta) = \sum_{t=1}^{T} log\ f(y_t|\theta, \eta, I_{t-1}) \tag{8}$$

with

$$f(y_t|\theta, \eta, I_{t-1}) = |H_t|^{-1/2} g(|H_t|^{-1/2}(y_t - \mu_t)|\eta) \tag{9}$$

where the dependence with respect to θ occurs through μ_t and H_t. The term $|H_t|^{-1/2}$ in the equation is the Jacobian that arises in the transformation from the innovations to the observables. It should be emphasized that if the g function belongs to the class of elliptical distributions, i.e., if it is a function of $z_t' z_t$, the maximum likelihood estimator depends on the choice of decomposition of $|H_t|^{-1/2}$, since $z_t' z_t = (y_t - \mu_t)' H_t^{-1} (y_t - \mu_t)$ and the maximum log–likelihood function of the study are optimized assuming the multivariate Student's t distribution. This is explained for the disturbance term process, (U_t), as follows

$$g(z_t|\theta,\nu) = \frac{\Gamma(\frac{\nu+N}{2})}{\Gamma(\frac{\nu}{2})[\pi(\nu-2)]^{\frac{N}{2}}}[1+\frac{z_t' z_t}{\nu-2}]^{-\frac{N+\nu}{2}} \tag{10}$$

Here, Γ is the Gamma function, and η is equal to ν. The density function of y_t for $(t = 1, \dots, T)$ is a stochastic process which is obtained by applying Eq. (10). As revealed by Bauwens et al. (2006), the Student's t density has an extra scalar parameter, with the degrees of freedom denoted by ν. Furthermore, a distinct feature of the Student's t density is that when ν trends towards infinity, the density trends towards the normal, whereas when it trends towards zero, the tail of the density becomes increasingly thicker. The parameter value indicates the order of existence of the moments. For example, if $\nu = 2$, only the first–order moments exist, and the second–order moments do not exist. Hence, for an advanced account of the conditional variance–covariance parameterization, it is convenient to assume that $\nu > 2$ (see, for details, Harvey et al., 1992; Fiorentini et al., 2003; inter alia). As aforementioned, all the parameters of this study are assessed simultaneously, rather than estimating mean and standard deviation parameters separately. To determine the statistical inference in model assessment, the robust standard errors of Bollerslev and Wooldridge (1992) are considered. To conduct optimization, the numerical algorithm named the BFGS quasi–Newton method from Press et al. (2007) is employed[4].

4. Empirical results and discussion

Throughout this section, the empirical results from model estimation will be exhaustively discussed. For broad discussion purposes, the section is split into several subsections that include: (1) interpretation

[4] BFGS is an acronym stands for Broyden–Fletcher–Goldfarb–Shanno algorithm.

of the results of the bivariate version of an asymmetric and unrestricted VAR(p)–GARCH–M–BEKK model, (2) the diagnostic checks and model specification tests which relied on uni– and multivariate serial correlations, (3) heteroscedasticity as well as variance non–causality analysis, and (4) an analytical framework of impulse responses of real output growth volatility to inflation uncertainty under a vector autoregression process. More specifically, the parameter estimates applied diagnostic checks and model specification tests for selecting the most favorable assessed model. As stated, the shocks and volatility spillovers between the growth series are potentially related. Here, variance non–causality tests of the real output growth volatility to inflation uncertainty is inspected to achieve the second specific objective of the study. Finally, we held the generalized impulse response function analysis of output growth volatility to a one unit of inflation uncertainty of the respective economy under a vector autoregression process.

4.1 Estimation results for unrestricted and asymmetric VAR(p)–BGARCH–M–BEKK model

As mentioned above, to determine the optimal lag length on the conditional mean equations, four famous selection criteria given in Lütkepohl (2005), namely the AIC, SBC, HQC and FPE criterion, are employed, and they include a vector autoregression process in a single equation. As far as the multivariate model is concerned, the used selection criteria show a vector autoregression order of lag two for China, lag five for Japan and lag six for South Korea as the most favored estimated models. In fact, if the chosen p is too small, one might miss the information content at higher lags; if it is too large, the degrees of freedom increase, and the power of an estimated model may be reduced. To select the most preferred model, we further relied on LL values and residual diagnostic checks. In terms of selection criteria and robustness tests as well as with the distribution of the explanatory variables (output and inflation growth) for the available sample sizes, the maximum vector autoregression order is set to ensure sufficient degrees of randomness and to avoid numerical convergence problems. In addition, the model specification and estimation results are utilized in volatility transmission and impulse response analysis between the series of concern. However, the conditional means and variance–covariance specifications of the study are jointly estimated; for convenience and space concerns, their results are reported in discrete tables.

Table 2
Parameter estimates for the unrestricted and asymmetric VAR(p)–BGARCH–M–BEKK model.

China				Japan				South Korea			
y_t		π_t		y_t		π_t		y_t		π_t	
μ_y	1.3013***	μ_π	0.3336*	μ_y	– 0.4767	μ_π	– 0.1963***	μ_y	– 0.6324	μ_π	– 0.0149
$\gamma_{y,y}^{(1)}$	– 0.2665***	$\gamma_{\pi,y}^{(1)}$	0.0516***	$\gamma_{y,y}^{(1)}$	– 0.3740***	$\gamma_{\pi,y}^{(1)}$	– 0.0052	$\gamma_{y,y}^{(1)}$	– 0.4808***	$\gamma_{\pi,y}^{(1)}$	0.0042
$\gamma_{y,\pi}^{(1)}$	– 0.0071	$\gamma_{\pi,\pi}^{(1)}$	0.2629***	$\gamma_{y,\pi}^{(1)}$	– 0.0545	$\gamma_{\pi,\pi}^{(1)}$	0.1096***	$\gamma_{y,\pi}^{(1)}$	– 0.1692	$\gamma_{\pi,\pi}^{(1)}$	0.2447***
$\gamma_{y,y}^{(2)}$	– 0.0810***	$\gamma_{\pi,y}^{(2)}$	0.0303***	$\gamma_{y,y}^{(2)}$	0.0558	$\gamma_{\pi,y}^{(2)}$	0.0122***	$\gamma_{y,y}^{(2)}$	– 0.0765	$\gamma_{\pi,y}^{(2)}$	0.0017
$\gamma_{y,\pi}^{(2)}$	– 0.0467	$\gamma_{\pi,\pi}^{(2)}$	0.0771**	$\gamma_{y,\pi}^{(2)}$	– 0.3050***	$\gamma_{\pi,\pi}^{(2)}$	0.0543	$\gamma_{y,\pi}^{(2)}$	– 0.4370***	$\gamma_{\pi,\pi}^{(2)}$	0.0041
—	—	—	—	$\gamma_{y,y}^{(3)}$	0.3741***	$\gamma_{\pi,y}^{(3)}$	0.0034	$\gamma_{y,y}^{(3)}$	0.1741***	$\gamma_{\pi,y}^{(3)}$	0.0033
—	—	—	—	$\gamma_{y,\pi}^{(3)}$	– 0.0847	$\gamma_{\pi,\pi}^{(3)}$	0.0482	$\gamma_{y,\pi}^{(3)}$	0.1288	$\gamma_{\pi,\pi}^{(3)}$	0.0636
—	—	—	—	$\gamma_{y,y}^{(4)}$	0.1014***	$\gamma_{\pi,y}^{(4)}$	0.0019	$\gamma_{y,y}^{(4)}$	0.2232***	$\gamma_{\pi,y}^{(4)}$	0.0020
—	—	—	—	$\gamma_{y,\pi}^{(4)}$	– 0.2353***	$\gamma_{\pi,\pi}^{(4)}$	0.0523*	$\gamma_{y,\pi}^{(4)}$	– 0.2354	$\gamma_{\pi,\pi}^{(4)}$	0.0664
—	—	—	—	$\gamma_{y,y}^{(5)}$	0.0292	$\gamma_{\pi,y}^{(5)}$	0.0060	$\gamma_{y,y}^{(5)}$	0.1200***	$\gamma_{\pi,y}^{(5)}$	0.0002
—	—	—	—	$\gamma_{y,\pi}^{(5)}$	– 0.1534	$\gamma_{\pi,\pi}^{(5)}$	0.1065***	$\gamma_{y,\pi}^{(5)}$	0.0270	$\gamma_{\pi,\pi}^{(5)}$	0.0482
—	—	—	—	—	—	—	—	$\gamma_{y,y}^{(6)}$	0.0753**	$\gamma_{\pi,y}^{(6)}$	0.0045
—	—	—	—	—	—	—	—	$\gamma_{y,\pi}^{(6)}$	– 0.3700**	$\gamma_{\pi,\pi}^{(6)}$	0.1216***
$\psi_{y,y}$	– 0.0244	$\psi_{\pi,y}$	0.0244**	$\psi_{y,y}$	0.0492	$\psi_{\pi,y}$	– 0.0048	$\psi_{y,y}$	0.5095	$\psi_{\pi,y}$	– 0.0022
$\psi_{y,\pi}$	– 1.4301***	$\psi_{\pi,\pi}$	– 0.4359**	$\psi_{y,\pi}$	2.2834***	$\psi_{\pi,\pi}$	0.8647***	$\psi_{y,\pi}$	3.8846***	$\psi_{\pi,\pi}$	0.4041**
Shape	2.5768***	*LL*	– 924.074	*Shape*	6.6532***	*LL*	– 1622.849	*Shape*	4.9529***	*LL*	– 1426.937
AIC	5.944	HQC	6.085	AIC	4.663	HQC	4.767	AIC	5.927	HQC	6.080
SBC	6.297	FPE	5.945	SBC	4.932	FPE	4.664	SBC	6.317	FPE	5.928
$H_t = C'C + A'u_{t-1}u'_{t-1}A + B'H_{t-1}B + D'\zeta_{t-1}\zeta'_{t-1}D$											
c_{11}	0.0476	c_{12}	—	c_{11}	0.3649***	c_{12}	—	c_{11}	2.2905***	c_{12}	—
c_{21}	0.8429***	c_{22}	– 0.0000	c_{21}	0.0129**	c_{22}	0.0000	c_{21}	– 0.0072	c_{22}	0.0000
a_{11}	1.4555***	a_{12}	– 0.0737***	a_{11}	– 0.0153	a_{12}	0.0034*	a_{11}	0.5529***	a_{12}	0.0038
a_{21}	0.8017***	a_{22}	0.4147**	a_{21}	– 0.0158	a_{22}	0.0167	a_{21}	0.3820	a_{22}	0.3303***
b_{11}	0.0947**	b_{12}	0.0060	b_{11}	0.9394***	b_{12}	– 0.0020*	b_{11}	– 0.0320	b_{12}	– 0.0194***
b_{21}	1.4800***	b_{22}	0.0827	b_{21}	0.0037	b_{22}	0.9912***	b_{21}	0.2322	b_{22}	0.9455***
d_{11}	– 0.6809	d_{12}	0.0263	d_{11}	0.3767***	d_{12}	0.0126***	d_{11}	0.6793***	d_{12}	0.0170**
d_{21}	0.0114	d_{22}	0.4112**	d_{21}	0.0769	d_{22}	– 0.1836***	d_{21}	– 0.9783	d_{22}	– 0.0733

Notes: ***, **, * indicate 1%, 5% and 10% significance level, respectively. AIC, SBC, HQC and FPE are acronyms for the Akaike information criterion, Schwarz Bayesian criterion, Hannan–Quinn criterion and Final prediction errors, respectively. LL stands for log–likelihood value.

In Table 3, the results for estimated conditional mean equations of output and inflation growth for the economies of concern are reported. Here, μ_y and μ_π are the coefficients of intercept that carry the positive values for China's real output and nominal inflation growth equations and holds the negative values for the other remaining intercepts of the equations. We first consider matrices $\Gamma^{(i)}$ $(i = 1,2)$ for China, $\Gamma^{(i)}$ $(i = 1,2,3,4,5)$ for Japan and $\Gamma^{(i)}$ $(i = 1,2,3,4,5,6)$ for South Korea, which are used in the mean equations and captured by the parameters $\gamma_{kj}^{(i)}$ to realize the relationships across the growth series of concern. While all the diagonal parameters of $\gamma_{kj}^{(i)}, k = j$, $\gamma_{y,y}^{(2)}, \gamma_{\pi,\pi}^{(2)}, \gamma_{\pi,\pi}^{(3)}, \gamma_{y,y}^{(5)},$ for Japan and $\gamma_{y,y}^{(2)}, \gamma_{\pi,\pi}^{(2)}, \gamma_{\pi,\pi}^{(3)}, \gamma_{\pi,\pi}^{(4)}, \gamma_{\pi,\pi}^{(5)}$ for South Korea are statistically significant, the diagonal parameters $\gamma_{y,y}^{(2)}, \gamma_{\pi,\pi}^{(2)}, \gamma_{\pi,\pi}^{(3)}, \gamma_{\pi,\pi}^{(4)}, \gamma_{\pi,\pi}^{(5)}$ for China are statistically significant, where $k, j = 1,2$. These equations depend on their first order lag and up to second lag for China, up to fifth lag for Japan, and up to sixth lag for South Korea. It should be noted that the cross–variable log change links between the variables under investigation can be examined by the off–diagonal elements, and the results are noteworthy. First, the off–diagonal parameters $\gamma_{\pi,y}^{(1)}, \gamma_{\pi,y}^{(2)}$ for China, $\gamma_{y,\pi}^{(2)}, \gamma_{\pi,y}^{(2)}, \gamma_{y,\pi}^{(4)}$ for Japan, and $\gamma_{y,\pi}^{(2)}, \gamma_{y,\pi}^{(6)}$ for South Korea are statistically significant. Further, their counterparts $\gamma_{y,\pi}^{(1)}, \gamma_{y,\pi}^{(2)}$ for China, $\gamma_{y,\pi}^{(1)}, \gamma_{\pi,y}^{(1)}, \gamma_{y,\pi}^{(3)}, \gamma_{\pi,y}^{(3)}, \gamma_{\pi,y}^{(4)}, \gamma_{y,\pi}^{(5)}, \gamma_{\pi,y}^{(5)}$ for Japan, and $\gamma_{y,\pi}^{(1)}, \gamma_{\pi,y}^{(1)}, \gamma_{\pi,y}^{(2)}$, $\gamma_{y,\pi}^{(3)}, \gamma_{\pi,y}^{(3)}, \gamma_{y,\pi}^{(4)}, \gamma_{\pi,y}^{(4)}, \gamma_{y,\pi}^{(5)}, \gamma_{\pi,y}^{(5)}, \gamma_{y,\pi}^{(6)}, \gamma_{\pi,y}^{(6)}$ for South Korea are also statistically insignificant. Since the first specific objective of the study is to evaluate the impact of output growth volatility on inflation uncertainty for these three East Asian economies, it can be inferred from the sign and significance of $\psi_{y,\pi}$ that the point estimates of these economies is equal to –1.4301 in China, 2.2834 in Japan, and 3.8846 in South Korea (with one percent significant *p*–values). As a result, relying on model estimation, the conditional standard deviation of China's real output growth volatility has a significant adverse impact on inflation uncertainty of the economy, whereas Japan and South Korea economies' real industrial output growth volatility have significant positive impacts on inflation uncertainty. Additionally, the tail parameters (shape) of all models show that these results are statistically significant values.

Table 2 further informs the estimated parameters of matrices *C, A, B,* and *D,* which are detailed in the conditional second moment equation. In the equation, the diagonal elements of matrix A, a_{11} and a_{22} capture own ARCH effects, while the off–diagonal elements a_{12} and a_{21} evaluate the effects of shock to real output lagged growth on the contemporaneous uncertainty of inflation growth of the respective economies of concern, and vice versa. Referring on the table entries, a set of results are worth mentioning. First, the statistical significant coefficients of a_{11} and a_{22} for China and South Korea's economies imply that the volatilities of output growth and inflation uncertainty of these economies are affected by the shocks from their own growths, while no shock effects are observed in Japan from its own growth. Next, we found an evidence of bi–directional shock transmissions between the real industrial output growth volatility and nominal inflation uncertainty of China's economy, but uni–directional shock transmissions between the growth series of Japan economy. This is because a pair of off–diagonal parameters. a_{12} and a_{21}, are statistically significant for China's economy, whereas only a single parameter, a_{12}, is statistically significant for Japan's economy. Likewise, the parameters a_{21} for Japan and both a_{12} and a_{21} for South Korea are statistically insignificant. The evidence of bi–directional shock transmissions infers a robust linkage between China's real output growth volatility and highly precariousness on its inflation uncertainty, whereas uni–directional shock transmissions express availability of one side linkages only between Japan's real output growth volatility and its nominal inflation uncertainty.

Similar to the interpretation of the elements of matrix A, the diagonal elements, b_{11} and b_{22} in matrix B, capture own GARCH effects, while off–diagonal elements, b_{12} and b_{21} measure the effects of lagged volatility of real output growth on the current volatility of nominal inflation uncertainty for the respective economies under concern, and vice versa. Since the diagonal elements of the matrix B, b_{11} and b_{22} generally express a strong GARCH (1,1) process, which drives from the conditional standard deviations, all these statistical significant elements (except for the inflation growth of China and output growth of South Korea) for the respective economies showing the highly heteroscedasticity in residual terms of the employed model. Moreover, we found the positive one–way spillover effects from China's nominal uncertainty to its real industrial output growth volatility, while uni–directional adverse volatility spillover effect from the real industrial output growth volatility to inflation uncertainty of Japan and South Korea's economies. Indeed, this is due to fact that the point estimate of b_{21} is statistically significant for China economy and that of b_{12} of Japan and South Korea's economies are statistically insignificant. As previously mentioned, the significant point estimates of b_{12} and b_{21} express that the volatility of past growth in China's real industrial output growth volatility does have an impact on the current instability of China's nominal inflation uncertainty, whereas the volatility of past growth in Japan and South Korea's economies nominal inflation uncertainty does have an impact on the current instability of real industrial output growth volatility.

Moreover, as far as asymmetric parameter matrix D is concerned, there is evidence of an asymmetric response to negative shocks for growth, as the diagonal parameters d_{11} and d_{22} are statistically significant. Based on model assessment, the significance of d_{11} and d_{22} implies that real output volatility and nominal inflation uncertainty display their own variance asymmetry to negative shocks. Therefore, a negative growth shock leads to more volatility on growth series, but a positive shock of a similar magnitude does not. Here also several cross–variable asymmetric responses are detected. The statistically significant off–diagonal elements of the matrix D, specifically d_{12} of Japan and South Korea's economies, generally express that the output growth volatility of them responses asymmetrically towards to the shocks of nominal inflation uncertainty of the respective economies of concern.

4.2 Robustness checks and model specification tests

Since the specified first and second moment equations are estimated by using quasi–maximum likelihood estimation, proposed by Bollerslev and Wooldridge (1992) as a way to inspect the robustness of the model, we performed serial dependence and heteroscedasticity tests relying on the standardized residuals of output and inflation growth equations. These standardized residuals are defined as $z_{j,t} = u_{j,t}/\sqrt{h_{j,t}}$ for $j = y$ and π, where y and π represent the returns of output and inflation growths.

Table 3
Results for robustness checks and model specification tests.

	Univariate	$z_{y,t}$	$z_{\pi,t}$	Multivariate	Statistic
Panel A: Uni– and multivariate tests					
China	Ljung–Box Q(4)	3.1936	4.4270	MQ (4)	$\chi^2(16)$ = 20.329
	McLeod–Li (4)	0.5803	6.1718	MARCH (6)	62.960
	ARCH LM (4)	0.1430	1.4630		
Japan	Ljung–Box Q(4)	3.1363	0.6912	Q (4)	$\chi^2(16)$ = 5.9311
	McLeod–Li (4)	6.3886	13.623***	MARCH (8)	67.180
	ARCH LM (4)	1.6690	3.2210**		
South Korea	Ljung–Box Q(4)	2.4629	1.1038	MQ (4)	$\chi^2(16)$ = 7.2805
	McLeod–Li (4)	1.1888	6.0506	MARCH (8)	66.770
	ARCH LM (4)	0.2700	0.9320		
Panel B: Specification tests					
China	Diagonal VAR	$H_0: \gamma_{12}^{(i)} = \gamma_{21}^{(i)} = 0,\ i = 1,2$			$\chi^2(4)$ = 40.903***
	Diagonal GARCH	$H_0: a_{ij} = b_{ij} = d_{ij} = 0, \text{if } i \neq j;\ i,j = 1,2$			$\chi^2(4)$ = 66.755***
	No GARCH	$H_0: a_{ij} = b_{ij} = d_{ij} = 0$ for all $i,j = 1,2$			$\chi^2(12)$ = 123.1***
	No GARCH–M	$H_0: \psi_{ij} = 0$ for all $i,j = 1,2$			$\chi^2(4)$ = 53.618***
	No Asymmetry	$H_0: d_{ij} = 0$ for all $i,j = 1,2$			$\chi^2(4)$ = 5.9365
Japan	Diagonal VAR	$H_0: \gamma_{12}^{(i)} = \gamma_{21}^{(i)} = 0,\ i = 1,2$			$\chi^2(10)$ = 343.7***
	Diagonal GARCH	$H_0: a_{ij} = b_{ij} = d_{ij} = 0, \text{if } i \neq j;\ i,j = 1,2$			$\chi^2(6)$ = 31.756***
	No GARCH	$H_0: a_{ij} = b_{ij} = d_{ij} = 0$ for all $i,j = 1,2$			$\chi^2(12)$ = 3287***
	No GARCH–M	$H_0: \psi_{ij} = 0$ for all $i,j = 1,2$			$\chi^2(4)$ = 75.716***
	No Asymmetry	$H_0: d_{ij} = 0$ for all $i,j = 1,2$			$\chi^2(4)$ = 200.55***
South Korea	Diagonal VAR	$H_0: \gamma_{12}^{(i)} = \gamma_{21}^{(i)} = 0,\ i = 1,2$			$\chi^2(10)$ = 317.4***
	Diagonal GARCH	$H_0: a_{ij} = b_{ij} = d_{ij} = 0, \text{if } i \neq j;\ i,j = 1,2$			$\chi^2(6)$ = 39.832***
	No GARCH	$H_0: a_{ij} = b_{ij} = d_{ij} = 0$ for all $i,j = 1,2$			$\chi^2(12)$ = 4988.0***
	No GARCH–M	$H_0: \psi_{ij} = 0$ for all $i,j = 1,2$			$\chi^2(4)$ = 40.265***
	No Asymmetry	$H_0: d_{ij} = 0$ for all $i,j = 1,2$			$\chi^2(4)$ = 51.651***

Note: ***, **, * indicate 1%, 5% and 10% significance level, respectively.

The entries of Table 3 present the results of the model specification test and diagnostic checks for the standardized residuals of output ($z_{y,t}$) and inflation ($z_{\pi,t}$) growths equation for the respective economies. Panel A reports the Ljung–Box Q test of Ljung and Box (1978) and the McLeod–Li test of McLeod and Li (1983), statistics which refer to serial correlation and dependence in uni–variate versions of the standardized residuals at lag four for all economies under investigation, respectively. The ARCH Langrage multiplier (hereinafter, LM) test shows the remaining heteroscedasticity up to lag four with statistically significant p–values (except for the standardized residuals of the inflation growth equation ($z_{\pi,t}$) for Japan's economy). In turn, the multivariate version of the serial correlation and the ARCH test for heteroscedasticity have been applied to the vector of series as a whole. Additionally, the table entries have an asymptotic chi–square (χ^2) distribution with a degree of freedom equal to the number of restricted parameters. The Panel A further presents a multivariate version of the serial dependence test by Hosking (1980). To test for the remaining heteroscedasticity in standardized residuals, the ARCH LM test of Engle (1982) is utilized, and a multivariate test routine is proposed by the RATS software packages. All in all, the heteroscedasticity test results suggest that there is no remaining heteroscedasticity in standard errors, and the conditional mean and variance–covariance specifications of the study are found as well–specified.

Following Grier et al. (2004) we also conducted the specification tests for the adequacy of model estimation. The tests' statistics with null hypotheses are reported in Panel B of Table 3 and they are noteworthy. First, relying on preliminary data analysis, there is significant conditional heteroscedasticity in the series of concern. Also, it can be also confirmed that the parameter matrices *A*, *B* and *D* provide the jointly statistically significant parameter estimates. As given in Panel B, all the entries of the elements of parameter matrices are jointly significant, and they show a well–specified second moment equation. Second, the jointly statistical significant off–diagonal elements of these parameter matrices express that the lagged conditional variances in real output growth volatility of the respective economies have an impact on the conditional volatility of inflation growth. Next, the joint significance of the elements of parameter matrix *D* clarifies that the specified conditional second moment equation is asymmetric. As stated, the asymmetric responses are only detected for the specified model on the linkage between real output growth volatility and inflation uncertainty of the economies of Japan and South Korea; no asymmetric responses are detected for China's economy. In addition, the significance of a_{11} and d_{11} shows evidence of variance asymmetry in output growth, and it expresses that the negative innovations in real output growth for the respective economies lead to more output volatility than positive shocks. Likewise, the significance of a_{22} and d_{22} also displays the response of own variance asymmetry in inflation growth, and it implies that negative inflation shocks raise more nominal uncertainty. Finally, the elements of the matrix Ψ (psi) state the existence of GARCH–M effects. All in all, the estimated models of the respective economies under investgiation are found with sound specification.

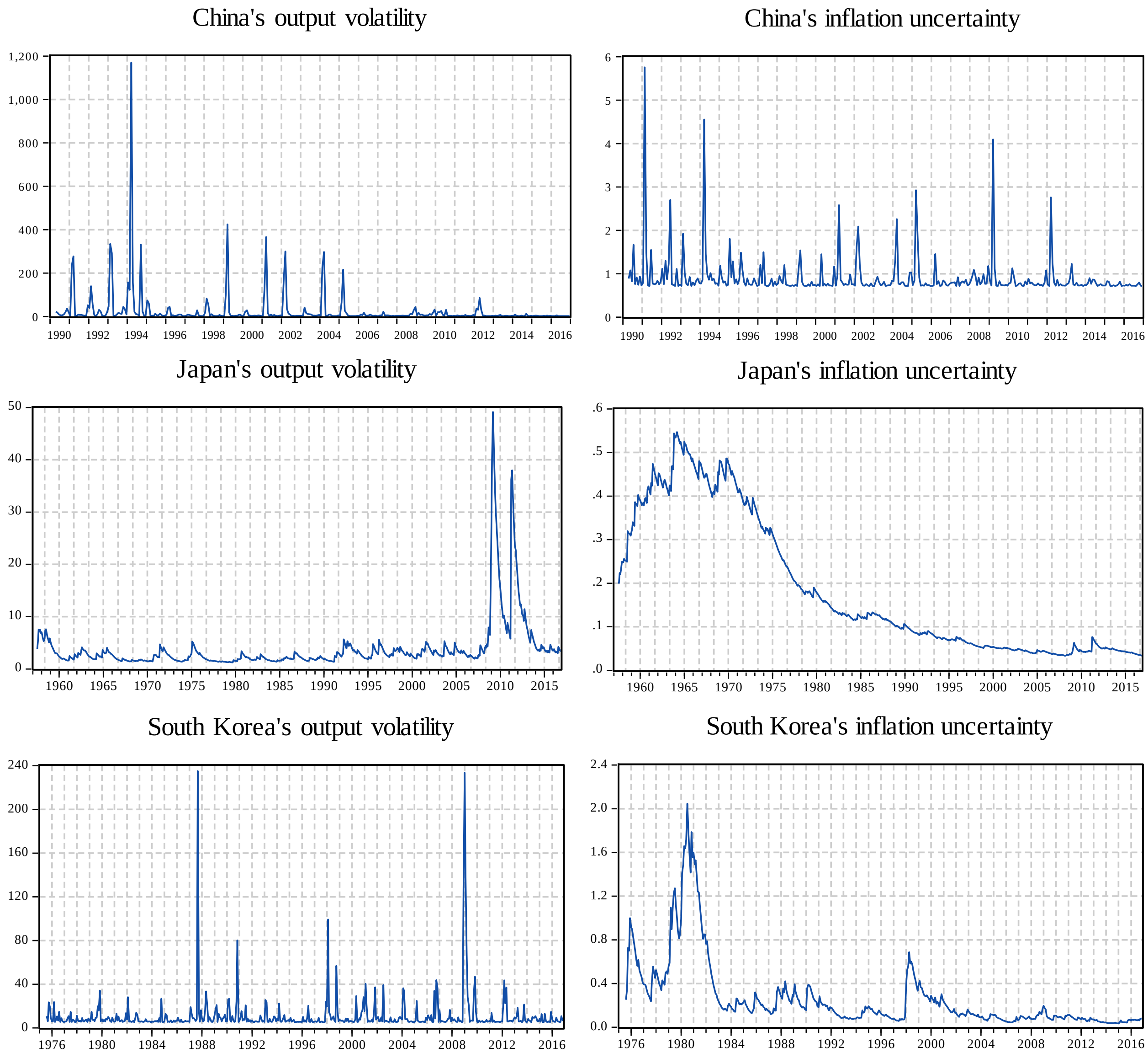

Fig. 2. Estimated conditional variances of output and inflation growths for the sample periods.

Fig. 2 illustrates visual inspections that China's economy produced a strong performance in output growth dynamics and that the conditional standard deviation was highly volatile for the first half of 1990's, reaching the highest level in 1994, then again collapsed over time from around 1200 to slightly above its original level. Likewise, Japan and South Korea's economies also exhibit highly volatile performances in output growth, but their estimated standard deviations are quite low (below 50 for Japan and below 240 for South Korea) compared to China's output growth volatility for the whole sample period, except for during the 2009 and 2012 fiscal years in Japan and during the 1988, 1998 and 2008 fiscal years in South Korea. Li and Kwok (2009) argue that the main reason for emerging volatility on Japan's output growth is due to its liquidity trap in the early 1990's, but the shocks brought by AFC had even not effected these economies. This is because until 1998, they had not yet become highly integrated with Southeast Asian economies. In a word, output growth volatility of these selected East Asian economies stayed stable over the required period, except for the stated time periods. Among them, Japan's output growth volatility was rather stable, but the negative shock that emerged following the GFC of 2008 has caused its recently high volatile output performances. In the case of South Korea's economy, it performed the most volatile output growth compared to the other economies, as its trend of estimated conditional variances reached its highest level only during the recent period of economic slowdowns. More specifically, the adverse shocks of the GFC have considerably affected its output growth, and the conditional variance reached 240 percent once more in 2008.

Fig. 2 also provides inflation uncertainty performances of the economies under investigation. The uncertainty performances of China's inflation growth series display exceedingly frequent fluctuations over the sample period; inflation growth in for China underwent its highest level in the middle of 1991, at around 5.8 percent, while the last sharpest peak was due to shocks brought by the GFC in 2008. Quite similar observations have also smarted the inflation uncertainty performances of the economies of Japan and South Korea. Initially, Japan's inflation growth uncertainty trend gradually increased and grasped it's the highest level (around 0.54 percent) in 1964; after, sudden and sharp declines were routine, and this steady decrease is enduring at around 0.04 percent until now. In the case of South Korea's economy, the nominal uncertainty performances of its inflation growth are more abnormal compared to the other two East Asian economies. Starting from 1975, three major abrupt peaks (in 1976, 1980 and 1998, among them the highest one was around 2.1 percent in the outset of 1980) and numerous incessantly sharp declines have trimmed its inflation uncertainty trend. These continuously sharp declines were ensured owing to swift responses of the recent decade's economic and financial downturns. All in all, referring to the plotted conditional standard deviations of the series of concern, it can be justified that the features of simultaneous occurrence of high (low) uncertainty in inflation growth are tracked with high (low) volatility series of output growth of the respective economies under study. Besides, all of the returns are characterized by volatility clustering, i.e., large (small) fluctuations of output growth are followed by small (large) fluctuations of inflation growth. As the clusters usually tend to occur simultaneously between the output growth volatility and inflation uncertainty, the volatility occurrences have been modeled systematically.

4.3 Variance non–causality analysis

The causal relations among the macroeconomic time series variables are also one of the recent key issues. So far, based on the Granger causality test of Granger (1969), numerous empirical works have contributed to the literature by analyzing the linear systems of time series (Hafner and Herwartz, 2008). Overall, for testing the hypothesis of variance non–causality, the literature distinguishes two approaches: mean (first moment) and variance–covariance (second moment) causality (see, e.g., Cheung and Ng, 1996; Comte and Lieberman, 2000; inter alia). Regarding the second specific objective of the study, we conducted variance non–causality analysis of Hafner and Herwartz (2008), which is achieved by estimating parameters for BEKK variance–covariance specification of Engle and Kroner (1995). This estimation method allows for establishment of a one–to–one relationship between the variables in–causality and particular testable zero restrictions imposed on the parametric model. As proposed by Hafner and Herwartz (2008), we demonstrate two conditionally heteroscedastic and stationary growth series such as $Y_{1,t}$ and $Y_{2,t}$ for $Y = y$ and π, where we denote the returns of output and inflation growths, respectively. Here, we consider that $Y_{2,t}$ does not Granger cause $Y_{1,t}$ in variance, designated by $Y_{2,t} \nrightarrow Y_{1,t}$ if,

$$Var\left(y_{1,t}\middle|\mathcal{F}_{t-1}\right) = Var\left(y_{1,t}\middle|\mathcal{F}_{t-1}\right) \forall t \in Z \tag{11}$$

Eq. (11) outlines any causality relationships; if $Y_{1,t}$ does Granger cause $Y_{2,t}$ in variance, the conditional variance of $Y_{2,t}$ can be prophesied more precisely by fitting the information set of $Y_{1,t}$. Here, the null hypothesis of Granger causality from the real output growth volatility on inflation uncertainty in the second moment equation of the economies under concern is stated as follows

$$H_0: a_{12} = b_{12} = 0 \tag{12}$$

By the same token, the null hypothesis of the Granger causality from the nominal inflation uncertainty to real output growth volatility of the respective economies is also specified as follows

$$H_0: a_{21} = b_{21} = 0 \tag{13}$$

To test these hypotheses, following to the proposed model by Hafner and Herwartz (2008), the standard Wald test statistics are constructed as follows

$$W_t = T\left(Q\hat{\vartheta}\right)'\left(Q\hat{\Sigma}_{\vartheta}Q'\right)^{-1}\left(Q\hat{\vartheta}\right)' \tag{14}$$

where, $Q = \left[0, \tilde{Q}, \tilde{Q}\right]$, $\hat{\vartheta} = \left(vech\left(\hat{C}\right)', vech\left(\hat{A}\right)', vech\left(\hat{B}\right)'\right)'$, $\hat{\Sigma}_{\vartheta} = E[u_t u_t'] < \infty$ and the asymptotic chi–squared (χ^2) distribution has a degree of freedom equal to the number of restricted parameters of the statistic, as given by

$$W_t \xrightarrow{\mathcal{L}} \chi^2_{k(K-k)}. \tag{15}$$

In the analysis, we apply Wald test statistics proposed by Hafner and Herwartz (2008) to carry out variance non–causality analysis on estimated model of the study.

Table 4
Results for variance non–causality analysis.

	China		Japan		South Korea	
	$y_t \downarrow$	$\pi_t \downarrow$	$y_t \downarrow$	$\pi_t \downarrow$	$y_t \downarrow$	$\pi_t \downarrow$
$y_t \rightarrow$	59.163***	7.2572**	1819.5***	12.253***	36.257***	9.2198***
$\pi_t \rightarrow$	61.572***	10.572***	0.0612	13322***	1.2070	26456.***

Notes: ***, **, * indicate 1%, 5% and 10% significance level, respectively. The signs "→" and "↓" represent the causative in the direction of the arrows.

Table 4 presents variance non–causality test results in conditional variances, and they follow the asymptotic chi–squared (χ^2) distribution with a degree of freedom that is unrestricted in parameter estimation. As reported in the table, there are two–directional causality in–variance between China's real output volatility and its inflation uncertainty, and uni–directional causality response is found for real output growth volatility to nominal inflation uncertainty of Japan and South Korea's economies.

4.4 Generalized Impulse Response Function Analysis

Hitherto, with the crucial linkages of the variables under study adequately discussed, an analytical framework of the dynamic impulse response of output volatility on one unit of inflation uncertainty shocks under a vector autoregression process, and vice versa, will be inspected. To scrutinize the time profile of the impact of output growth volatility shocks on future behavior of the inflation (nominal) uncertainty, we employed the generalized impulse response function (hereinafter, GIRF) proposed by Koop et al. (1996). It provides insight into how significantly an innovation in output growth volatility (often measured by one standard deviation) shock may affect uncertainty of inflation growth of the economies of concern through the dynamic interaction for the sample period. We plotted an analytical framework of impulse responses of output growth volatility to one unit of inflation uncertainty under the vector autoregression process. Following to Grier et al. (2004), the shock effects of output growth volatility to inflation uncertainty are defined through the conditional mean and with a lag through the second moment equation. Grier et al. (2004) further stress allowing for composition dependence in multivariate models of GIRF as their first advantage compared to other customary impulse response functions, and here the effects of a shock to output growth volatility are not isolated from having a contemporaneous impact on nominal uncertainty, and vice versa (see, e.g., Lee and Pesaran, 1993; Pesaran and Shin, 1998; inter alia). As given in Grier et al. (2004) the GIRF of the study is detailed as follows

$$GIRF_K(n, \varrho_t, \omega_{t-1}) = E[K_{t+n}|\varrho_t, \omega_{t-1}] - E[K_{t+n}|\omega_{t-1}] \tag{16}$$

where, $n = 0,1,2,3$ …, thus the GIRF is conditional on ϱ_t and ω_{t-1} and constructs the response by averaging out future shocks given in the past and present. By giving this, a natural reference point for GIRF is the conditional expectation of K_{t+n} given only the history ω_{t-1}, and in this benchmark response the current shock is also averaged out.

The analytical framework of the GIRF of output volatility to one standard deviation shocks of inflation uncertainty under the vector autoregression process and inflation uncertainty to one standard deviation shocks to output volatility of the respective economies under concern are illustrated in Fig. 3.

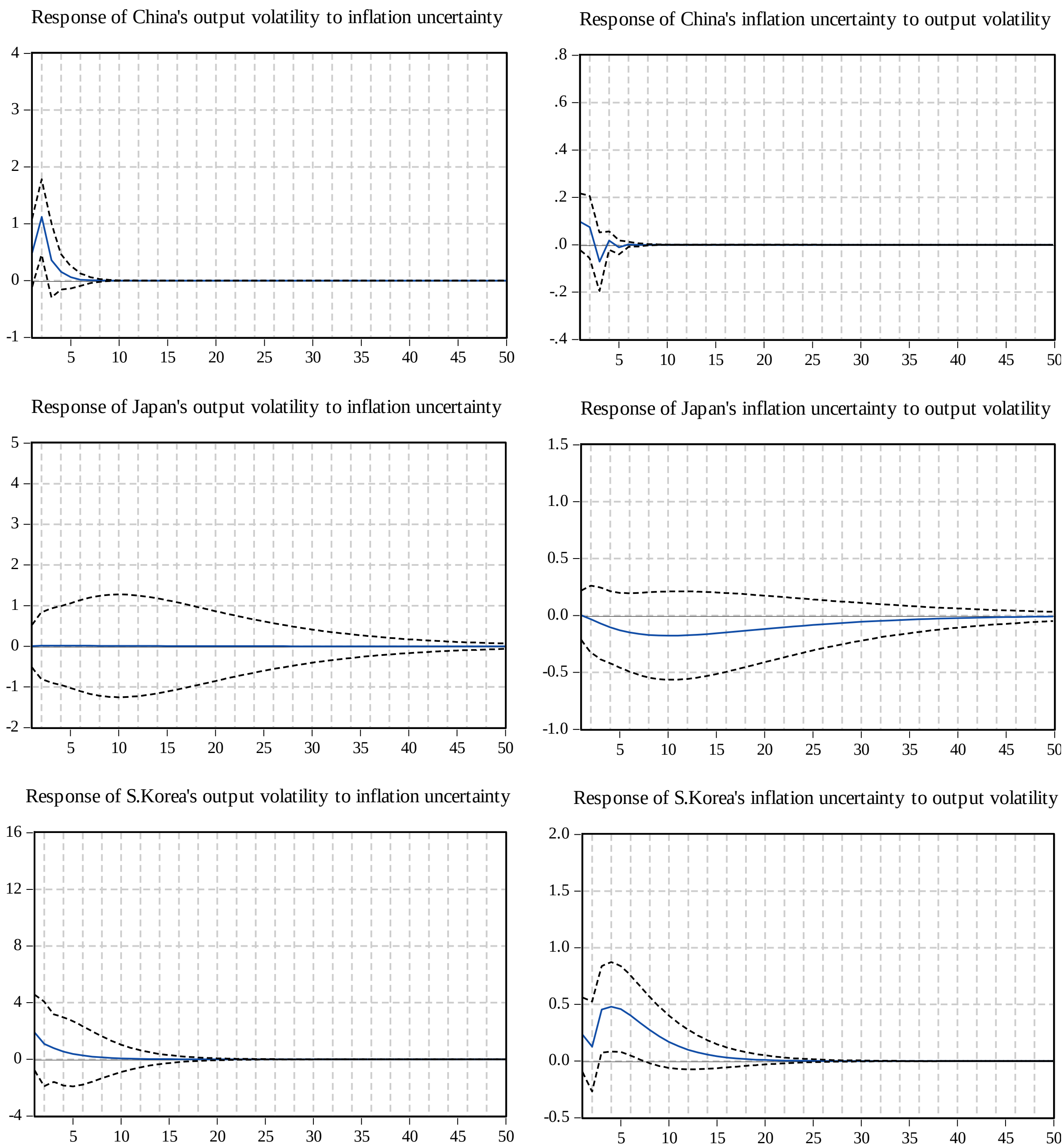

Fig. 3. Generalized impulse response function of output volatility under vector autoregression process to a unit (one standard deviation) shock of inflation uncertainty, and vice versa. Innovations $\pm$ 2.S.E.

Referring to Fig. 3, the solid blue line is the response to a unit of shock innovations, while the black dashed lines are the confidence intervals; each unit time horizon denotes a month. There is evidence to suggest that the shocks of output volatility have a positive and statistically significant impact on the inflation uncertainty of China and South Korea's economies, while ambiguously exposing Japan's economy. Prior to the effect of the shock, the real output volatilities of the economies of China, Japan, and South Korea have an immediate response of approximately 0.5%, 0% and 2, respectively, relative to inflation uncertainty shocks of one standard deviation. The GIRF grows after the shock effect in China's economy and reaches one percent point of the initial unit shock within two months; this effect takes around eight months for fully dissipate. In the case of Japan's economy, the initial unit sized growth rate is zero percent, and no changes occur after the shock; full recovery requires more than fifty months. For China's economy, the same scenario occurs as with South Korea's economy: the adversely initiated growth rate starts from two percent and continues to steadily decrease over the half year; the effect takes more than one and half years to fully disappear following the shock. In total, the real industrial output volatility of China and South Korea's economies seem to have a positive and significant impact in the response to shocks of inflation uncertainty, whereas there is an insignificant, ambiguous impact on Japan's economy for the same sample period.

5. Concluding remarks

This study mainly focuses on the dynamic linkages between the real industrial output volatility and inflation uncertainty of three East Asian economies. Moreover, the causal relationships between the variables under study and an analytical framework of impulse responses is conducted. These framework involves responses of output volatility for the respective economies after inflation uncertainty shocks and is examined using a vector autoregression process. In order to achieve the specified objectives of the study, a bivariate version of an asymmetric and unrestricted econometric approach, labeled as VAR(p)–BGARCH–M–BEKK, is employed, and the results are noteworthy. First, in the light of bi–variate GARCH–in–mean model assessment, the conditional variances of China's real output growth volatility have a significant adverse impact on inflation uncertainty of the economy, while in Japan and South Korea economies' there is a significant positive impact on inflation uncertainty. That is, whether or not volatility is accommodated for in China economy, real industrial output growth is subject to promote inflation uncertainty. Next, relying on variance non–causality analysis, we find evidence of two–directional shock transmissions between the growth series of China. There is also evidence that output growth volatility of the economy is mutually causative, with its inflation uncertainty and the nexus between them quite robust, even after controlling the uncertainty. However, only one–way causality from output growth volatility to inflation uncertainty of the economies is detected for Japan and South Korea economies, respectively. The analytical framework of impulse responses function under a vector autoregression model and express that the model incorporating uncertainty tends to show the effect of shocks lasting longer. Importantly, the results of this model are instructive for policymakers in selected East Asian economies, as output volatility provides a beneficial conduit for higher inflation uncertainty, even during the volatile periods.

In spite of high economic potential in the economies of China, Japan, and South Korea, there still remains problems with an inflexible currency regime and extensive capital controls for businesses (He and Chen, 2014). Specifically, small– and medium–sized enterprises have limitations in financing opportunities from the capital markets. As long as China and South Korea are still considered developing countries, the numerous institutional deficiencies and policy distortions in the economies will be an internal source of instability. In a similar way, Li and Kwok (2009) argue that Japan and South Korea have been suffering from banking constraints by the increasing size of the non–performing loans and internal debts. Eventually, these constraints will reduce economic growth and slow down recovery in the event of a large economic shock (Prasad, 2008).

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