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Research Article

Economic Policy Uncertainty Linkages among Asian Countries: Evidence from Threshold Cointegration Approach

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This paper employs the threshold cointegration methodology to assess the long- and short-run dynamics of asymmetric adjustment between economic policy uncertainty (EPU) of China-India, China-Japan, China-Korea, India-Japan, India-Korea, and Japan-Korea pairs using monthly EPU data ranging from January 1997 to April 2020. The relationship between the EPU pairs is examined in terms of Engle-Granger and threshold cointegrations. The findings provide evidence of long-run threshold cointegration and that the adjustments towards the long-run equilibrium position are asymmetric in the short run for the China-India and India-Japan EPU pairs in M-TAR specification with nonzero threshold values. Also, the results suggest a unidirectional causal relationship between China-India, China-Japan, and India-Korea EPU pairs in the long and short run using the spectral frequency domain causality approach. However, a bidirectional causal relationship between China-Korea, India-Japan, and Japan-Korea pairs exists in the long and short run. Therefore, the findings provide some clues to economic policymakers within the Asian subregion for possible policy uncertainty synergies and spillovers among the Asian countries.

1. Introduction

Weakening global economic growth in recent years has been attributed to heightened uncertainty in the economic policies of advanced economies. Global issues such as 1997-98 Asian financial crisis, September 11 terrorist attacks in the United States (US), Gulf War II, 2008 global financial crisis, European sovereign debt crisis, Brexit referendum, and Covid-19 pandemic are perceived to have raised economic policy uncertainty (EPU) with consequential effects on private domestic demand in many economies. Usually, the rise in uncertainty after such events may lead to a gradual widespread of "wait" and "see" attitude, resulting in postponed spending projects until anticipation for economic activity to become more obvious [1, 2]. Although EPU linkages are considered at the cross-country level, the impact of such uncertainty on economic activity and the behaviour of economic agents at the household and firm levels cannot be underestimated [3–6]. The key question that has lingered in the minds of international macroeconomists and policymakers is the extent to which EPU shocks emanating from one country affect the economic policy uncertainty as well as the business cycles in another country. Specifically, small open economies with free capital mobility, sizeable openness, and a large financial sector are greatly influenced by the international transmission of EPU shocks.

Emerging market economies have experienced large swings in business cycles, financial market returns, and macroeconomic fundamentals due to EPU shock transmission from advanced and developed economies. For example, EPU spillovers from the United States (US) and the European Union (EU) would have crucial global consequences because of their relatively large size, strong trade, and financial linkages with other economies. Aside looking at EPU impact from the global perspective, regional and subregional EPU linkage is eminent due to regional and subregional economic integration. Economies within the same subregion with considerable large financial sectors are likely to experience increased EPU codependency, especially during postmajor economic, financial, and political shocks.

The Asian economies have emerged as force in the global economic architecture in production, trade, and financial sector. For example, the Asian financial sector which is highly susceptible to shock from uncertainty represents 37% of the total world banking and insurance market capitalisation [7–9]. Again, the Asian economy over the last decade has increased its share of global Gross Domestic Product (GDP) from 24% to 31%, and with a deepened regional integration, the possibility of policy uncertainty shocks to transmit from one Asian country to another would be apparent. This regional integration that underpins the economic policy linkages of the Asian economies is evident in the revival of China's relationships with India, Japan, and South Korea, as well as the reboot of China, Japan, and the Republic of Korea trilateral summit. The possible regional integration of the Asian economy and its contagion effects prevailed during the 1997/98 Asian financial crisis that started in Thailand and spread across the subregion. Therefore, examining the comovement of EPU among the Asian countries is of great importance because of its impact on array of economic activities such as stock markets, housing price, commodity prices, and many more [10–22].

In this paper, we investigate the EPU linkages among four Asian countries, comprising China, Japan, South Korea, and India using threshold cointegration techniques to determine the long- and short-run asymmetric adjustments and comovement of EPUs between these countries. Our choice of threshold cointegration method over the traditional linear cointegration method is based on its ability to detect the presence of a long-run relationship between time series variables and to unearth asymmetries in adjustment towards fundamental values with respect to positive and negative shocks. Thus, the power of linear cointegration test is lower in an asymmetric adjustment process [23]. Moreover, because the nexus of time series variables is higher in harsh periods than in tranquil periods, it makes it important to use threshold cointegration to be able to detect the presence of long-run equilibrium relationship with asymmetric adjustments towards the fundamental values between EPUs. Enders and Siklos [24] threshold cointegration method is employed to study the asymmetric long-run relationship between EPU of the four Asian countries because the impact of economic issues such as EPU is mostly nonstationary and nonlinear.

In addition, we focus our study on Asian countries, precisely China, Japan, South Korea, and India because of their economic size and power within the Asian subregion, and the EPU shock of one of these countries can easily influence the EPU and other macroeconomic factors of the other. Moreover, the widespread 1997/98 Asian financial crises across other Asian countries give a clear indication of how contagious policy inconsistency in one of these countries could be, which is motivating enough for our study to focus on Asian countries. Focusing on Asia as an emerging economy brings about interesting dynamics to the study of EPU linkages among countries because of the central role the Asian economy plays in global production networks [25–28] and the evidence of most of the Asian emerging economies catching up financially with the matured economies [29, 30]. Although any of the Asian countries could have been selected for this study but due to limitation of data on EPU of most of the developing countries, only these four Asian countries have complete data over the whole sample period.

Studies that focus on EPU regarding the Asian economy investigate the impact of EPU spillovers of advanced economies on the Asian financial markets, most especially the stock markets [31–33]. None of the previous studies explicitly focused on investigating EPU shock transmission among the Asian countries that showed widespread contagion of the 1997/98 Asian financial crisis. The study that is close to ours is Balcilar et al.'s study [34], they investigated the impact of EPU shock transmission of US and EU on local EPU and other macroeconomic factors of the Asian economy using quantile vector autoregression (QVAR) but did not examine EPU linkages among the Asian countries. To the best of our knowledge, this is the first study to investigate the transmission of EPU shocks from one Asian country to another using a threshold cointegration approach. The findings of the study reveal a long-run relationship between the EPU pairs of the countries and the adjustment of positive deviations in the short run was more rapid in general than negative deviations, implying that EPU of one country responds quickly when another country's EPU increases. Additionally, the Granger causality tests in the frequency domain suggest both unidirectional and bidirectional causalities of the EPU pairs in the long and short run.

Therefore, using the threshold cointegration would uncover the upward and downward adjustments of the short-run deviation of one country's EPU shock transmission to other country's EPU in the long-run. Knowing the extent to which local EPUs of Asian countries link together would help policymakers of the Asian economies to be on their guard and watch economic policies of not only the advanced economies but countries within their subregion so that they can mitigate any possible adverse effects these uncertainties may bring to bear on their economies. The results from the empirical analysis showed long-run threshold cointegration with asymmetry in the short run, in particular for China-India and India-Japan. Again, a unidirectional causal relationship between China-India, China-Japan, and India-Korea EPU pairs in the long and short run using the spectral frequency domain causality approach were observed. Finally, long- and short-run bidirectional causal relationship between China-Korea, India-Japan, and Japan-Korea pairs were found. These findings present important policy implication for dealing with uncertainty spillover in the region. The rest of the study is organised as follows. Section 2 reviews the relevant literature while Section 3 outlines the methodology and description of the data. Section 4 presents the empirical results, and the conclusion of the study is provided in Section 5.

2. Literature Review

EPUs over the period have surged at the bane of the global financial crisis (GFC) and the Eurozone's serial crises as well as partisan policy disputes in the US. As suggested by the Federal Open Market Committee [35], the uncertainty about US and European fiscal, regulatory and monetary policies contributed to a steep economic decline in 2008/09 and slow recovery afterwards. According to Klößner and Sekkel [36], the uncertainty spillover that increases notably around turbulent times accounts for more than 25% of the dynamics of the policy uncertainty index. These policy uncertainties have a significant impact on financial markets and a growing interest in the literature relating to the link between EPU and international financial markets, most especially the stock markets, which have led several researchers to focus on this area. Thus, studies by Brogaard and Detzel [37], Arouri et al. [38], Bahmani-Oskooee and Saha [12], Adam [39], Asafo-Adjei et al. [40], and Chiang [41] have demonstrated that heightened uncertainty hurts stock returns. Moreover, Pastor and Veronesi [42], Liu and Zhang [43], Tsai [44], and Jurado et al. [19] with different research orientations focus on the impact of uncertainty on stock market volatility and find that the inclusion of EPU can enhance the predictability of stock returns. This assertion confirms Hansen et al.'s [45] finding, which indicates that an upward shift in stock volatility is due to heightened policy uncertainty.

Specifically, Pastor and Veronesi [46] showed that higher policy uncertainty is associated with lower stock prices, higher volatility, and higher correlations among stock returns. Using Granger causality tests, Sum [47] investigated the effect of US EPU on five ASEAN countries comprising of Indonesia, Malaysia, Philippines, Singapore, and Thailand and found that US EPU harms stock market returns of related countries. Chuliá et al. [48] examined the impact of US policy and US equity market uncertainties on domestic and other stock market returns. Their findings provide evidence that an uncertainty shock lessons stock market returns both in developed and developing countries in uncertain times. In addition, Trung [49] tests the impact of U.S. uncertainty on emerging economies and finds that an upward shift in U.S. policy uncertainty inhibits international capital inflows and investment activity, which causes stock prices to fall in emerging economies. Bhattarai et al. [50] investigated the spillover indices of US uncertainty shock on fifteen emerging market economies (EMEs) by utilizing the panel vector autoregressive (VAR) method. They found evidence that the US uncertainty has harmful effects on EME stock prices, exchange rates, country spreads, and capital inflows into them. Akadiri et al. [51] found evidence of causality between international tourism arrivals (ITAs) and EPU of three regions of America, Europe, and Asia-Pacific using annual frequency panel data that consist of 12 countries in a multivariate Granger causality model setting. Their results revealed two-way causality relationship between ITAs and EPU in France, Ireland, and United States and one-way causality relationship from ITAs to EPU in Brazil, Canada, China, and Germany, while between ITAs and EPU in Chile, Japan, South Korea, Russia, and Sweden, there were no causality relationships. To establish the nexus between EPU and carbon dioxide (CO_2) emissions, Adams et al. [52] used the World Uncertainty Index to analyse the long-run relationship of EPU, energy consumption, and CO_2 emissions for countries including Brazil, China, India, Israel, Russia, Saudi Arabia, South Africa, Turkey, Ukraine, and Venezuela over the period 1996 to 2017. Their results based on the panel pooled mean group-autoregressive distributed lag model showed a significant association between EPU and CO_2 emissions in the long run. The causality analysis conducted also revealed bidirectional relationship between EPU and CO_2 emissions.

Beside numerous studies that focus on EPU, stock price movements, and other macroeconomic variables, other studies focus on cross-country effects of uncertainty. For example, Klößner and Sekkel [36] used the policy uncertainty index to examine cross-country EPU effects of six developed countries and found evidence of a significant spillover effect of policy uncertainty from the US and the United Kingdom (UK) to other countries which are the recipients of policy uncertainty shock during and after the crises period. Luk et al. [53] studied EPU spillovers of US, Europe, Mainland China, and Japan in small open economies, using Hong Kong as a case study. They constructed EPU for Hong Kong from 1998 to 2016 and found large spillovers of uncertainty from major economies to Hong Kong. Cekin et al. [54] investigated the dependence structure of EPU in four Latin American economies (Brazil, Chile, Colombia, and Mexico) and by employing vine copula modelling with various forms of tail dependence, they found significant dependencies in economic uncertainty among the Latin American economies. By adopting QVAR model approach, Balcilar et al. [34] extended their examination of external EPU spillovers of US and EU to the local EPU of five Asian economies (China, Hong Kong, Japan, South Korea, and India). They found that global economic policy uncertainties makeall Asian countries' domestic EPU fragile, except China and Hong Kong. Bai et al. [55] investigated the economic risk contagion among major economies including the US, UK, Germany, France, Japan, and China using an innovative spillover analysis method in time and frequency domains. The empirical results showed that in time-domain framework, the economic uncertainty of the six largest economies are strongly connectedwith the US happens as both major risk spillover contributors and receiver in the frequency domain, especially, at the short-term frequency. Their results also revealed that the static net EPU spillover effects indicate on average that the US is the key transmitter, while the UK and China are the major spillover receivers.

Even though there are vast number of studies on EPU shock spillover linkages between developed and developing economies and its impact on the financial markets, none of the studies enumerated examines EPU linkages among the Asian countries, having in mind the widespread contagion of the 1997/98 Asian financial crisis across the Asian subregion. Also, our adoption of the threshold cointegration method of Enders and Siklos [24] differentiates our study from the existing studies in terms of methodology as none of the studies to the best of our knowledge has used the approach employed in our study to investigate EPU shock transmission across developed and emerging economies in general and among the Asian economies in particular.

3. Methodology and Data Description

3.1. Threshold Cointegration and Error Correction Model. To investigate the dynamic adjustment properties from EPU of one country to the other, the threshold cointegration test technique introduced by Enders and Siklos [24] is followed to identify the existence of an asymmetric long-run relationship between the EPUs of four Asian countries. To start with, Engle and Granger [56] long-run cointegration test is used to establish the stability, linearity, and long-run relationship between the EPU pairs of countries. The test is performed under the assumption that the linearity in the adjustment to the long-run equilibrium, as well as an increase or decrease in the deviation from the long-run equilibrium relationship, is symmetric. The long-run relationship between the EPU pairs of the four countries is estimated as follows:

$$\mathrm{EPU}_{j,t} = \alpha_0 + \alpha_1 \mathrm{EPU}_{i,t} + \mu_t, \tag{1}$$

where $\mathrm{EPU}_{j,t}$ and $\mathrm{EPU}_{i,t}$, respectively, represent EPU of country j and i at time t and μ_t is the normally distributed residual or error term with zero expected mean and constant variance.

To cater for the presence of nonlinearity in the variables and the adjustment process, the linear cointegration technique cannot detect as such. Therefore, for this reason, we apply Enders and Siklos [24] threshold cointegration where the long-run cointegration is linear but the adjustment to long-run equilibrium level is nonlinear. Therefore, we employ the threshold autoregressive (TAR) and momentum threshold autoregressive (M-TAR) models of Enders and Siklos [24] threshold cointegration to estimate the long-run cointegration and nonlinear adjustments to the long-run equilibrium level. The TAR model is specified as follows:

$$\Delta\mu_t = I_t \rho_1 \mu_{t-1} + (1 - I_t)\rho_2 \mu_{t-1} + \sum_i^k \gamma_i \Delta\mu_{t-i} + \varepsilon_t, \tag{2}$$

where μ_t is the residual in equation (1) substituted into equation (2) and ε_t is a zero-mean, constant variance, independent identically distributed (iid) random variable. I_t denotes the Heaviside indicator function specified as

$$\mathrm{TAR}: I_t = \begin{cases} 1, & \text{if } \mu_{t-1} \geq \tau, \\ 0, & \text{if } \mu_{t-1} < \tau, \end{cases} \tag{3}$$

where τ is the threshold value that is endogenously suggested by Chan [57]. The M-TAR model is also specified by replacing the indicator variable I_t and the level of previous period's residual μ_{t-1} in equation (2), respectively, by M_t and the change in the level of previous period's residual $\Delta\mu_{t-1}$ with the Heaviside indicator function stated as follows:

$$\mathrm{M-TAR}: M_t = \begin{cases} 1, & \text{if } \Delta\mu_{t-1} \geq \tau, \\ 0, & \text{if } \Delta\mu_{t-1} < \tau. \end{cases} \tag{4}$$

If μ_{t-1} and $\Delta\mu_{t-1}$ are above the threshold value τ, then the adjustment coefficient is $\rho_1\mu_{t-1}$, while on the other hand, the adjustment coefficient becomes $\rho_2\mu_{t-1}$, if μ_{t-1} and $\Delta\mu_{t-1}$ are below the threshold value τ. The threshold procedure involves three stages. The first stage is to estimate the TAR and M-TAR models for the cointegration procedure. At this stage, the null hypothesis of no cointegration (H_0: $\rho_1 = \rho_2 = 0$) is tested by comparing the critical values of the F-statistics with their corresponding actual values Φ in accordance with Enders and Siklos [24]. If the null hypothesis of no cointegration is rejected, the long-run cointegration between the EPU of country j and the EPU of country i exists and this takes us to the second stage of the threshold cointegration procedure. In the second stage, the symmetry of the null hypothesis (H_0: $\rho_1 = \rho_2$) is estimated. If we reject the null hypothesis of symmetry, thus $|\rho_1| \neq |\rho_2|$, there exists nonlinear threshold cointegration between the EPU of country j and the EPU of country i. We proceed to stage three where we estimate the threshold vector error correction model (TVECM) to adjust the short-run deviation from the long-run equilibrium. TAR specification of TVECM expression for EPU of country j and country i is stated as follows:

$$\Delta\mathrm{EPU}_{j,t} = \alpha_{j,0} + I_t \rho_{1,j} \mu_{t-1} + (1 - I_t)\rho_{2,j} \mu_{t-1} + \sum_k^n \beta_{j,k} \Delta\mathrm{EPU}_{i,t-k} + \sum_k^n + \varepsilon_{j,t}, \tag{5}$$

$$\Delta\mathrm{EPU}_{i,t} = \alpha_{i,0} + I_t \rho_{1,i} \mu_{t-1} + (1 - I_t)\rho_{2,i} \mu_{t-1} + \sum_k^n \alpha_{i,k} \Delta\mathrm{EPU}_{i,t-k} + \sum_k^n \beta_{i,k} \Delta\mathrm{EPU}_{j,t-k} + \varepsilon_{i,t}, \tag{6}$$

where ρ_1 and ρ_2, respectively, denote the speed of adjustment parameters for positive (above) and negative (below) deviations for one country's EPU from its long-run equilibrium, $\alpha_{j,0}$ and $\alpha_{i,0}$ are the constant terms, and $\Delta\mathrm{EPU}_{j,t-k}$ and $\Delta\mathrm{EPU}_{i,t-k}$ are the adjustments of EPU for country j and country i. respectively. $\alpha_{j,t}$, $\alpha_{i,t}$, $\beta_{j,t}$, and $\beta_{i,t}$ are the coefficients that quantify the short-term relationship among the EPU of country j, its lag, and the EPU of country i while $\varepsilon_{j,t}$

and $\varepsilon_{i,j}$ represent white noise disturbance terms. For M-TAR specification of TVECM expression, we replace I_t in equations (5) and (6) by M_t as defined by equation (4).

Because the parameters of the vector autoregression (VAR) model comprise of complex nonlinear functions, it complicates the statistical inference for the feedback measures over time [58, 59]. For this reason, we follow a Granger causality test in the frequency domain introduced by Breitung and Candelon [60] which is more useful if the causal links between variables change according to frequency such as the short and long run. We, therefore, adopt the spectral frequency domain approach to investigate the causal relationship between two time series variables based on bivariate spectral density matrix of VAR among different frequencies. According to Breitung and Candelon [60], the null hypothesis $(H_0)M_{y\longrightarrow x}(\omega) = 0$ corresponds to H_0: $R(\omega)\beta = 0$, where β is the vector of the coefficients on a given EPU index and

$$R(w) = \begin{pmatrix} \cos(\omega) & \cos(2\omega) & \ldots & \cos(p\omega) \\ \sin(\omega) & \sin(2\omega) & \ldots & \sin(p\omega) \end{pmatrix}. \tag{7}$$

The F-statistics in equation (7) are distributed as $F(2, T-2p)$ for $\omega \in (0, \pi)$, where T is the number of observations that measure the VAR model of order p. Performing the frequency domain analysis would allow us to observe nonlinear and causality cycles for high or low frequencies, and by presenting the relationship between the EPUs of the countries in a VAR system, a bidirectional relationship between the EPU_j of country j and EPU_i of country i in the short- and long-run is expressed as follows:

$$\text{EPU}_j = \lambda_1 \text{EPU}_{j,t-1} + \cdots + \lambda_p \text{EPU}_{j,t-p} + \theta_1 \text{EPU}_{i,t-1} + \cdots + \theta_p \text{EPU}_{i,t-p} + \varepsilon_{j,t}, \tag{8}$$

$$\text{EPU}_i = \lambda_1 \text{EPU}_{i,t-1} + \cdots + \lambda_p \text{EPU}_{i,t-p} + \theta_1 \text{EPU}_{j,t-1} + \cdots + \theta_p \text{EPU}_{j,t-p} + \varepsilon_{i,t}. \tag{9}$$

3.2. Data Description. The monthly EPU index data complied on four major Asian countries including China, India, Japan, and South Korea by Baker et al. [61] is used for the study. The index is based on the news coverage frequency of policy-related economic issues which serves as a proxy for policy-related economic uncertainty. There are many uncertainty measures for developed economies but less is said about emerging and developing economies as available EPU indices for developing countries are scanty in time scope. The EPU index that provides a scaled measure of the appearance of uncertainty in news surrounding economic issues is sourced from http://www.policyuncertainty.com. The data range from January 1997 to April 2020 during which the world experienced different categories of regional and global financial crises, such as the 1997-1998 Asian financial crisis, 2007–2009 global financial crises, 2010 European debt crisis, and 2015 stock market crash in China.

Figure 1 presents the time series plots of EPU of China, India, Japan, and South Korea. As shown in Figure 1, major regional and global events such as the 1997-98 Asian financial crises, 9/11 terrorist attacks in 2001, 2007-08 global financial crises, 2010 European debt crises, and Chinese stock market crash in 2015 broadly reflect spikes in the comovement of EPUs among China, India, Japan, and South Korea. We observe that the comovement of countries' EPUs during periods of such crises intensifies, confirming the fact that economic policymakers and the public including both local and foreign investors, are usually uncertain about the consequences of policy directions of a country in the periods of economic crisis.

Table 1 reports the main descriptive statistics of the variables over the period January 1997 to April 2020. On average, China has the highest EPU index among the EPU of the other countries, with India recording the lowest EPU index. The EPU of China exhibits higher variability than the other EPUs as shown by its minimum, maximum, and the standard deviation statistics, while Japan has lowest fluctuations in its EPU index as shown by its low standard deviation over the entire sample period. Overall, the EPU index of all countries are not normally distributed according to their skewness, kurtosis, Jarque–Bera test, and Shapiro test which indicate the presence of fat tails and confirms the stylized fact about the distribution of time series data being asymmetric.

Preliminary investigation of the comovement between the EPU pairs is carried out by assessing the unconditional correlation between the pairs and the results are presented in Table 2. The results show that the correlation coefficients of all the pairs are positive, indicating that the EPUs move in the same directional in pairs. The results also reveal a strong correlation between the Korea-China pair and Japan-India pair while a weak correlation can be observed between India-China pair.

4. Results and Discussion

4.1. Unit Root Test. We start our analysis by performing unit root tests to check whether the series are stationary or not using an autoregressive model. The augmented Dickey–Fuller (ADF) test, Phillips–Perron unit root test and Kwiatkowski, Phillips, Schmidt, and Shin (KPSS) test [62–64] are applied to test for the stationarity of the time series data used in our study as per the following equation:

$$\Delta x_t = (\varnothing - 1)x_{t-1} + \sum_{i=1}^{k-1} \lambda_i \Delta x_{t-i} + \mu_t + \nu_t, \tag{10}$$

where x_t is the series at time t, $\mu_t = \mu_0 + \mu_1 t$ is the deterministic term (μ_0 is the constant term and $\mu_1 t$ is the deterministic trend), and ν_t is a white noise process.

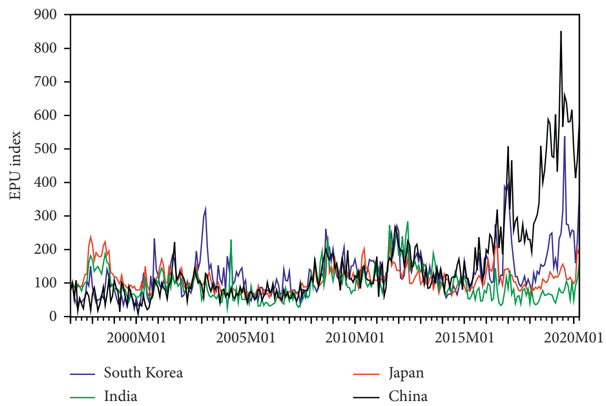

Figure 1: The relationship between the EPU indices

Table 1: Descriptive statistics.

	EPU_China	EPU_India	EPU_Japan	EPU_Korea
Mean	154.9016	94.7735	110.1837	128.7583
Min.	8.3459	24.9398	48.8858	22.4275
Max.	852.0525	283.6891	239.0284	538.1768
Std. Dev.	138.6951	46.954	36.4347	70.4422
Skewness	2.1668	1.2326	1.1455	1.7319
Kurtosis	4.7689	1.8143	1.4204	5.0573
Jarque–Bera	493.03*	111.54*	86.495*	446.74*
Shapiro	0.7389*	0.9122*	0.9210*	0.8779*

Note: *denotes the rejection of null hypothesis at 5% significance level.

Table 2: Unconditional linear correlation.

	EPU_C	EPU_I	EPU_J	EPU_K
EPU_C	1			
EPU_I	0.0004	1		
EPU_J	0.2171	0.6212	1	
EPU_K	0.6594	0.2587	0.3509	1

The results of the unit root tests of all series are shown in Table 3. In Table 3, the null hypothesis of all series at the level having unit roots cannot be rejected based on McKinnon [65] critical values at 5% level of significance using the ADF test, but after taking the first difference of the series, the ADF unit root test shows that all series are stationary at 5% level of significance, indicating that the series are integrated of order 1, $I(1)$. In addition to the ADF test, the Perron unit root test rejects the existence of unit root for all series at level, except Chinese EPU (EPU_C), which cannot be rejected at level, but after taking the first difference, the null hypothesis is rejected at 5% significance level, implying stationarity of EPU_C at first difference.

Moreover, the results of KPSS test show the rejection of the null hypothesis of stationarity of all series at level at 5% significant level. However, at first difference of all series, the stationarity hypotheses cannot be rejected at 5% level of significance. This indicates the presence of unit roots in the series at level but are however stationary after taking the first difference. It is key to note that all the variables have unit root problems in the presence of structural breaks. The structural break occurs around in 1997 mostly for EPU of India, Japan and South Korea, which brings to light the commencement of the Asian financial crisis. Moreover, a substantial break in China's EPU occurs around February 2018 where the Sino-US trade conflict intensified. In all, the ADF test and KPSS test confirm the presence of unit root in all series at level but the series are stationary after taking the first difference while the Phillip–Perron unit root test results indicate stationarity of series at level, except EPU_C that is stationary after taking first difference. Because the Phillip–Perron unit root test suffers from serious size distortions in the pure autoregressive case even in moderately large samples [66], we conclude based on the results of the ADF.

4.2. Testing for Nonlinear Characteristics of the Variables. To be able to proceed with the nonlinear cointegration analysis, we employ the nonlinear unit root test proposed by Kapetanios et al. [67], which has been considered as a nonlinear version of the ADF test. The purpose of Kapetanios, Snell, and Shin (KSS) test is to outline a testing procedure to specify the presence of nonstationary against a nonlinear exponential smoothing transition autoregressive (ESTAR) process which is globally stationary. The KSS is given by the following ESTAR specification:

$$\Delta y_t = \varphi y_{t-1}\left[1 - e^{-\theta\left(y_{t-1}-c\right)^2}\right] + \varepsilon_t, \quad (11)$$

where y_t is the time series of interest, φ is the unknown parameter, and ε_t is an iid error with mean zero and constant variance. The exponential transition function $[1 - e^{-\theta(y_{t-1}-c)^2}]$ is adopted in the test to present the nonlinear adjustment. When $c = 0$ is assumed, then equation (11) becomes

$$\Delta y_t = \varphi y_{t-1}\left[1 - e^{-\theta\left(y_{t-1}\right)^2}\right] + \varepsilon_t. \quad (12)$$

The null hypothesis of unit root, H_0: $\theta = 0$, is tested against the nonlinear ESTAR process, $H_1 > 0$, in equation (12). Because according to Kapetanios et al. [67] the null hypothesis cannot be directly tested, a reparameterization of equation (12) is suggested by computing a first-order Taylor series approximation to obtain auxiliary regression equation given by

$$\Delta y_t = \gamma y_{t-1}^3 + \varepsilon_t. \quad (13)$$

The case where the errors in equation (13) are serially correlated, it is extended with p augmentations to correct for serially correlated errors to become

$$\Delta y_t = \gamma y_{t-1}^3 + \sum_{j=1}^{p} p_j \Delta y_{t-j} + \varepsilon_t. \quad (14)$$

The null hypothesis of nonstationarity to be tested with either equation (13) or (14) is H_0: $\gamma = 0$ against the alternative H_1: $\gamma < 0$. The t-test statistics is given by $t = (\hat{\gamma}/se(\hat{\gamma}))$, where $\hat{\gamma}$ is the ordinary least square (OLS) estimate of γ and $se(\hat{\gamma})$ is the standard error of $\hat{\gamma}$. The critical values of the t statistics of the KSS unit root test are given for

Table 3: Unit root test results for all series.

ADF unit root test								Phillip–Perron unit root test						KPSS unit root test					
Level	Lag	Break date	Statistics	First difference	Lag	Break date	Statistics	Level	Lag	Statistics	First difference	Lag	Statistics	Level	Lag	Statistics	First difference	Lag	Statistics
EPU_C	4	201802	−4.0197	EPU_C	4	2019M06	−19.8666*	EPU_C	3	−3.7357	EPU_C	3	−27.7487*	EPU_C	1	1.5498*	EPU_C	1	0.0159
EPU_I	3	199704	−4.9129	EPU_I	2	2004M05	−19.5823*	EPU_I	5	−6.4978*	EPU_I			EPU_I	2	0.549*	EPU_I	3	0.022
EPU_J	5	199706	−3.2941	EPU_J	2	1997M12	−14.4354*	EPU_J	4	−5.5395*	EPU_J			EPU_J	1	0.5333*	EPU_J	1	0.0131
EPU_K	8	201904	−3.3005	EPU_K	5	1997M05	−20.2156*	EPU_K	2	−5.7414*	EPU_K			EPU_K	1	0.1909*	EPU_K	2	0.0146

Note: *denotes the rejection of the null hypothesis at 5% significance level. Lag lengths are chosen according to AIC. Critical values are from McKinnon [65] and KPSS unit root tests that all the series under consideration are stationary after taking first difference indicating that all are $I(1)$ (integrated of order one).

three cases referred to the model with the raw data, the demeaned data, and the detrended data at 1%, 5%, and 10% levels. In the case of our study, we present the KSS test results using critical values for the model with the raw data.

Table 4 presents the results of KSS unit root test. The null hypothesis of linear stationary cannot be rejected for China's EPU at 5% significant level indicating that China's EPU does not exhibit nonlinear characteristics. However, the null hypothesis of linear stationary is rejected for India's EPU, Japan's EPU, and Korea's EPU, which imply they are nonlinear stationary. Because three out of the four EPU indices exhibit nonlinear behaviour, it is therefore worth employing nonlinear models to investigate the nonlinear relationships between the EPU pairs.

4.3. Engle-Granger Cointegration. We apply the Engle-Granger cointegration test procedure as the first step to our cointegration analysis based on the estimation of equation (1) to ascertain the presence of long-run relationship between the EPU pairs of countries. Table 5 presents the models' residuals for all EPU pairs and the test results show that the null hypothesis of no cointegration is rejected at 5% level. This means that each EPU is cointegrated with one another, confirming the long-run relationship between all the EPU pairs of the countries. According to the long-run regression results in Table 5, a change in one country's EPU would influence the movement in other country's EPU in the same direction.

4.4. Enders-Siklos Cointegration Test Results. We employ Enders and Siklos [24] test to investigate the nonlinear threshold cointegration and the results are displayed in Tables 6 and 7. Both Tables 6 and 7 show the threshold effects and focus on convergence, threshold cointegration, and adjustment in the long-run equilibrium following a deviation in EPUs in the model expressed as the linear combinations of the pair of EPU variables. In both tables, the first column shows the various cointegration model specifications and the second and the third columns show the values of the adjustment parameters ρ_1 and ρ_2, while the fourth and the fifth columns, respectively, show the F statistic for the null hypothesis of no cointegration and the test results for the symmetric adjustment. Specifically, Table 6 exhibits the TAR parameter estimates by assuming a threshold value for each model to be zero which is deterministic in nature. The point estimates in the TAR model show the convergence of long-run equilibrium and that the speed of convergence for positive divergence is almost the same as the speed of convergence of negative divergence from the long-run equilibrium of all the paired EPU models, although the larger of the t statistics is the positive adjustment parameter ρ_1 which is greater than the 5% critical value, except the EPUs of India-Japan model where the larger of the t statistics is the negative adjustment parameter ρ_2. For all the models, the F-joint statistics (thus hypothesis that $\rho_1 = \rho_2 = 0$) are greater than the 5% critical value, implying that the null hypothesis of no cointegration is rejected at 5% significance level. This suggests long-run relationship between the EPU pairs of countries. On the other hand, F-equal statistics (thus hypothesis that $\rho_1 = \rho_2$) that test the null hypothesis of symmetric adjustment is lower than the 5% critical value for all the models, indicating that the null hypothesis of symmetric adjustment cannot be rejected. This implies the speed of adjustment from positive deviation is not significantly different from the speed of adjustment from negative deviation, signifying the rate at which one country's EPU responds to rise or fall in another country's EPU which is almost the same according to the TAR model.

Table 4: KSS unit root test results.

	EPU_C	EPU_I	EPU_J	EPU_K
Test statistics	−2.3455	−4.509*	−4.2881*	−5.6913*

Note: *represents significance at 5% level corresponding to −2.94 critical value.

Because the threshold value is not always zero, we follow the approach of Chan [57] to search for approximate threshold values to estimate consistent M-TAR models of all the EPU pairs. The threshold values with a minimum value of Akaike Information Criteria (AIC) obtained are 34.814 for China-India EPU pair, −31.058 for China-Japan EPU pair, 57.357 for China-Korea EPU pair, 7.932 for India-Japan EPU pair, 25.746 for India-Korea EPU pair, and 2.575 for Japan-Korea EPU pair. Table 7 shows a similar analysis as in Table 6 using the M-TAR specification to check for asymmetric movement in one country's EPU in relation to changes in another country's EPU. Similarly, the larger of the t statistics is the positive adjustment parameter ρ_1 which is greater than 5% critical value, implying the test statistics are significant at 5% level except for the EPUs of China-Korea model where the larger of the t statistics is the negative adjustment parameter ρ_2. The M-TAR model estimates suggest convergence in the long-run equilibrium and the speed of convergence for positive deviation is faster than the speed of convergence for negative deviation for China-India and India-Japan models, indicating an asymmetric adjustment in EPU pairs between these countries. The null hypothesis of no cointegration is rejected in all models as the value of F-joint statistics is greater than 5% critical value. This indicates that all models show a long-run equilibrium relationship between the EPU pairs of countries. To detect the possibility of asymmetric adjustment, the null hypothesis of symmetric adjustment cannot be rejected for most of the models, except for China-India and India-Japan combinations, where the F-equal statistics are greater than 5% critical value, indicating that the speed of adjustment of positive and negative deviations from long-run equilibrium is different. Thus, China's EPU reverts quickly to the equilibrium path whenever the EPU of India rises more than a fall and vice versa. Likewise, a rise in Japan's EPU leads to India's EPU reverting quickly to the equilibrium path more than a fall in Japan's EPU and it is also true for the converse.

4.5. The Error Correction Model Estimation Results in M-TAR Specification. To finally analyse the asymmetric cointegration adjustment, we estimate the M-TAR error correction

Table 5: Engle-Granger cointegration results.

	EPU_C/EPU_I	EPU_C/EPU_J	EPU_C/EPU_K	EPU_I/EPU_J	EPU_I/EPU_K	EPU_J/EPU_K
Test statistics	−9.0694*	−8.7004*	−16.3803*	−18.9545*	−16.6225*	−17.7117*

Note: *represents 5% significance level with a corresponding critical value equal to −1.95 level. Each column represents EPU pair combination of the models' residuals.

Table 6: Enders-Siklos cointegration test results according to the TAR model.

Model	ρ_1	ρ_2	$\rho_1 = \rho_2 = 0$	$\rho_1 = \rho_2$	Conclusion
EPU_C/EPU_I	−0.228* (−4.267)	0.202* (−2.782)	12.21* (0.000)	0.091 (0.764)	Cointegration exist/symmetric adjustment
EPU_C/EPU_J	−0.249* (−4.954)	−0.153* (−2.209)	14.223* (0.000)	1.315 (0.253)	Cointegration exist/symmetric adjustment
EPU_C/EPU_K	−0.432* (−6.481)	0.350* (−4.138)	27.252* (0.000)	0.651 (0.420)	Cointegration exist/symmetric adjustment
EPU_I/EPU_J	0.229* (−3.870)	−0.304* (−4.278)	15.224* (0.000)	0.723 (0.396)	Cointegration exist/symmetric adjustment
EPU_I/EPU_K	−0.214* (−4.091)	−0.206* (−2.751)	11.546* (0.000)	0.007 (0.933)	Cointegration exist/symmetric adjustment
EPU_J/EPU_K	−0.231* (−4.511)	−0.172* (−2.383)	12.419* (0.000)	0.461 (0.498)	Cointegration exist/symmetric adjustment

Note: *represents significance at 5% level. Numbers in parenthesis and square brackets are t-values and p values, respectively.

Table 7: Enders-Siklos cointegration test results according to the M-TAR model.

Model	ρ_1	ρ_2	$\rho_1 = \rho_2 = 0$	$\rho_1 = \rho_2$	Conclusion
EPU_C/EPU_I	−0.365* (−4.601)	−0.163* (−3.184)	14.812* (0.000)	4.871* (0.028)	Cointegration exist/asymmetric adjustment
EPU_C/EPU_J	−0.223* (−4.081)	−0.508* (−3.279)	13.520* (0.000)	0.034 (0.853)	Cointegration exist/symmetric adjustment
EPU_C/EPU_K	−0.368* (−3.371)	−0.410* (−6.868)	26.940* (0.000)	0.13 (0.719)	Cointegration exist/symmetric adjustment
EPU_I/EPU_J	−0.375* (−5.645)	−0.160* (−2.384)	17.766* (0.000)	5.686* (0.018)	Cointegration exist/asymmetric adjustment
EPU_I/EPU_K	−0.304* (−4.116)	−0.165* (−3.141)	12.699* (0.000)	2.515 (0.114)	Cointegration exist/symmetric adjustment
EPU_J/EPU_K	−0.265* (−4.604)	−0.151* (−2.447)	13.207* (0.000)	1.91 (0.168)	Cointegration exist/symmetric adjustment

Note: *represents significance at 5% level. Numbers in parentheses and square brackets are t-values and p values, respectively.

model (M-TVECM) specified in the modified equations (5) and (6) to establish the short-run relationships between the EPUs of the countries. Though only China-India and India-Japan EPU models produced asymmetry in the adjustment mechanism as shown in Table 7, the adjustment parameters ρ_1 and ρ_2 which represent the coefficients of the long-run relationship between the EPU pairs for the remaining models showed the significance of both positive and negative adjustments at 5% level, and since $|\rho_1| \neq |\rho_2|$, we estimate the M-TVECM for all the EPU pairs and the results are presented in Table 8. In Table 8, we have 12 M-TVECM estimated and each model comprises of each country's EPU as a dependent variable yielding three models each with the corresponding independent variables. The results suggest that the speed of adjustment of positive deviation is quicker than the negative one for most models except for three models where Korea is the dependent variable with the adjustment of negative deviations being more rapid than the positive ones. Because the coefficients of the error correction term which represent the coefficients of the long-run relationship are significant, we conclude that long-run relationship exists between the EPU pairs. Specifically, for the models where China's EPU is the dependent variable, the adjustment of the positive deviation of India's EPU is significant at 10% level, showing a positive relationship between India's EPU and China's EPU in the short run while the adjustment of the positive deviation of Korea's EPU is significant at 1% level and the joint coefficient of Korea's EPU positively impacts China's EPU in the short run. The adjustment of both positive and negative deviations of Japan's EPU is not significant at 5% level, indicating the failure of China's EPU to respond to the deviation of Japan's EPU in the short run. These results imply that upward movements in the EPU of India and Korea will cause upward movement in China's EPU as well.

In addition, models where India's EPU acts as the dependent variable, the adjustment of the positive deviation of China's EPU is although significant at 5% level and India's EPU rarely reacts to the deviation of China's EPU in the short run but instead converges to the equilibrium value in the long run. The adjustment of the positive deviation of Japan's EPU is significant at 1% level and Japan's EPU influences India's EPU positively in the short run. In the same breath, India's EPU responds to a positive deviation of Korea's EPU in the short run, as the adjustment of the positive deviation of Korea's EPU is significant at 1% level, indicating a positive relationship between Korea's EPU and India's EPU in the short run. These results imply that the increase in Japan's EPU and Korea's EPU caused an increase in India's EPU. Furthermore, having Japan's EPU as the dependent variable in the model, the adjustment of positive deviation of China's EPU is significant at 5% level but Japan's EPU rarely responses to short-run movements in China's EPU, instead it returns to the equilibrium path in the long run. Similarly, the fluctuations in India's EPU neither influences Japan's EPU in the short run nor the adjustment in either direction of the deviation of India's EPU is significant at 5% level, implying that the movement in India's

Table 8: The M-TVECM coefficient estimates.

Dep var		EPU_C			EPU_I			EPU_J			EPU_K	
Ind var	EPU_I	EPU_J	EPU_K	EPU_C	EPU_J	EPU_K	EPU_C	EPU_I	EPU_K	EPU_C	EPU_I	EPU_J
ΔEPU_C1	−0.702***	−0.559***	−0.676***							−0.210*		
	(−5.190)	(−4.543)	(0.5.636)							(−1.710)		
ΔEPU_C2	−0.630***	−0.548***	−0.596***									
	(−5.085)	(−4.345)	(−4.612)									
ΔEPU_I1	0.671***			0.338***	−0.484***	−0.466***						
	(3.586)			(−2.633)	(−3.541)	(−3.439)						
ΔEPU_I2				−0.416***	−0.420***	−0.432***						
				(−3.114)	(−3.063)	(−3.359)						
ΔEPU_J1					0.257*							0.428**
					(1.670)							(−2.048)
ΔEPU_J3							−0.317**	−0.383***	−0.0.320**			−0.490**
							(−2.382)	(−2.619)	(−2.455)			(−2.212)
ΔEPU_J4		−0.349*										
		(−1.755)										
ΔEPU_K1			0.304***			0.142**					−0.300**	
			(2.768)			(2.219)					(−2.082)	
ΔEPU_K2			0.421***									−0.203*
			(3.440)									(−1.750)
ΔEPU_K3									−0.091*		−0.318***	
									(−1.766)		(−2.670)	
ρ_1	0.097*	0.006	−0.176***	−0.233**	−0.260***	−0.161*	−0.150**	−0.155	−0.256***	−0.201	−0.023	−0.164**
	(1.837)	(0.163)	(−2.668)	(−2.239)	(−2.735)	(−1.661)	(−2.434)	(−1.628)	(−3.294)	(−1.110)	(−0.220)	(−1.996)
ρ_2	−0.050	−0.044	0.105***	−0.087	−0.038	−0.094	0.198	−0.117	0.012	−0.402***	−0.211***	−0.235***
	(−0.1.452)	(−0.935)	(2.606)	(−1.551)	(−0.461)	(−1.610)	(1.047)	(−1.355)	(0.161)	(−4.661)	(−2.810)	(−2.651)

Note: ***, **, and *represent 1%, 5%, and 10% levels, respectively. Numbers in parentheses represent t-values. "Dep var" denotes dependent variables and "Ind var" denotes independent variables. Each column under a dependent variable represents a model comprising of the dependent variable and an independent variable.

EPU does not impact the movement in Japan's EPU in the short run. In addition, the adjustment of the positive deviation of Korea's EPU is significant at 1% level, showing a negative relationship between Korea's EPU and Japan's EPU in the short run, implying that a rise in Korea's EPU leads to a fall in Japan's EPU.

Moreover, having Korea's EPU as the dependent variable in the model, we observe that the adjustment of the negative deviations of China's EPU and India's EPU and that of Japan's are significant at 1% level and these variables have a negative relationship with Korea's EPU in the short run except India's EPU, implying that upward movements in EPUs of China and Japan result in downward movement in Korea's EPU. Therefore, the abovementioned findings suggest both unidirectional and bidirectional movements between the EPUs of the countries in the short run either in the same or opposite direction to the bidirectional movements. In summary, movements in India's EPU affect China's EPU in the short run but not in the reverse case; while there is no significant short-run relationship between China's and Japan's EPUs, Japan influences China's EPU in the long run and this is consistent with Bai et al. [53] who found China to be the principal receiver of EPU spillover. For the case of China-Korea and India-Korea EPU pairs, there is a bidirectional movement between the pairs but in the opposite direction. Thus, while upward movement in Korea's EPU results in the upward movements in China's and India's EPUs, upward movements in both China's and India's EPUs result in downward movement in Korea's EPU and this relationship suggests a diversification potential for investors. Also, an upward movement in Japan's EPU causes India's EPU to move in the same direction in the short run but the movement in India's EPU in either direction does not impact movement in Japan's EPU in the short run. We can again infer from the M-TVECM that movements in EPUs of Japan and Korea impact the movement in the other in the opposite direction, indicating a significant negative relationship between these two variables in the short run. Thus, a rise in Japan's EPU results in a decline in Korea's EPU while an increase in Korea's EPU also results in a decline in Japan's EPU which again provides diversification opportunities for investors.

4.6. Estimating Causality among the EPUs in the Frequency Domain. As a final step in our analysis, we explore the existence of spectral causality among the EPU indices of the countries over the short and long run by estimating equations (7), (8), and (9). The test statistics for all frequencies in the interval $(0, \pi)$ are computed at 5% significance level and the frequencies correspond to a wavelength of $(2\pi/\omega) \sim 2$ years. The 5% critical value for the F-statistics with 2 and $(T - 2p)$ degrees of freedom corresponding to 2 and 272 degrees of freedom where the value of T is 280 observations and p is 4 (VAR order) is computed. Figure 2 shows the Granger causality between the EPUs of China and India in the frequency domain and at the 5% significance level. The null hypothesis that China's EPU does not Granger cause India's EPU cannot be rejected, implying that China's EPU does not significantly influence India's EPU in short and long run. On the other hand, India's EPU does cause China's EPU at frequencies corresponding to 5 to 6 months in the long run and 2 to 3 months in the short run. This shows a unidirectional causality between China's and India's EPU in the short and long run which indicates that the movement in India's EPU affects China's EPU through the short and long run.

Figure 3 shows the causality of China's EPU and Japan's EPU in the short and long run. The figure reveals China's EPU does not Granger cause Japan's EPU either in the short or long run at 5% significance level. This implies that the movement in China's EPU does not influence Japan's EPU in the short and long run. Japan's EPU Granger causes China's EPU at frequencies corresponding to 3 to 4 months in the long run, indicating unidirectional causality. This finding indicates that movement in Japan's EPU affects the movement in China's EPU in the long run. Figure 4 shows the frequency domain causality of China's EPU and Korea's EPU in the short and long run. Korea's EPU Granger causes China's EPU at 5% significant level at frequencies corresponding to 3 to 4 months in the long run, while China's EPU Granger causes Korea's EPU at frequencies corresponding 2 to 6 months in the short through to the long run. This is an indication of bicausality which implies that EPUs of both China and Korea influence each other in the long run.

Figure 5 displays the Granger causality of Japan's EPU and India's EPU in the frequency domain. The figure reveals the rejection of the null hypothesis that EPU of Japan does not Granger cause the EPU of India at all frequencies significant at 5% level which is rejected, indicating the movement in EPU of Japan affects movement in India's EPU in the long and short run. In the reverse case, India's EPU Granger causes Japan's EPU at significant frequencies corresponding to 2 to 3 months in the short run. The result shows bidirectional causality between the EPUs of Japan and India implying that movement in one EPU influences the other in the short and long run. Figure 6 depicts the frequency domain Granger causality of Korea's EPU and India's EPU at frequencies significant at 5% level. The null hypothesis that India's EPU does not Granger cause Korea's EPU cannot be rejected at all frequencies at 5% significance level, indicating that movement in India's EPU does not significantly influence the movement in Korea's EPU while Korea's EPU Granger causes India's EPU at frequencies corresponding to 2 to 3 months in the long run, implying the influence of Korea's EPU on the movement in India's EPU.

Finally, Figure 7 shows the Granger causality of Korea's EPU and Japan's EPU in the frequency domain. The EPU of Korea Granger causes the EPU of Japan at significant

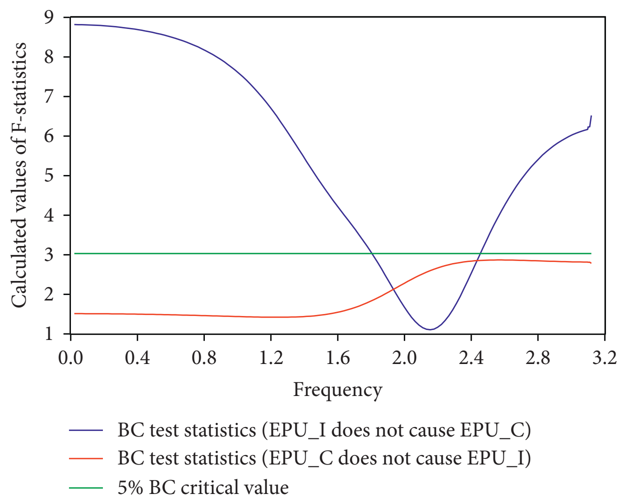

Figure 2: The frequency domain causality between China's and India's EPU indices. The part of the lines mentioned above the critical value-line indicates rejection of the null hypothesis of no Granger causality.

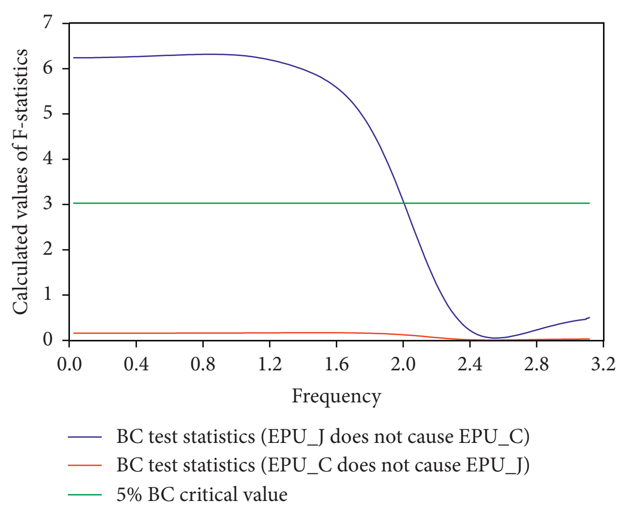

Figure 3: The frequency domain causality between China's and Japan's EPU indices. The part of the lines above the critical value-line indicates rejection of the null hypothesis of no Granger causality.

frequencies at 5% level corresponding to 2 to 3 months in the short run, implying that movement in Korea's EPU influences Japan's EPU in the short run. For the converse, Japan's EPU Granger causes Korea's EPU at frequencies corresponding to 3 to 5 months in the long run and 2 to 3 months in the short run. This implies a bidirectional causality between Korea's EPU and Japan's EPU in the long and short run. Thus, movements in both EPUs impact the other. The causal relationship between the EPU pairs of the countries is evidenced by the integration of the Asian economy through the formation of greater trade and investment linkages underpinned by East Asia's supply chain and production fragmentation and served as an engine of global trade and economic growth [66].

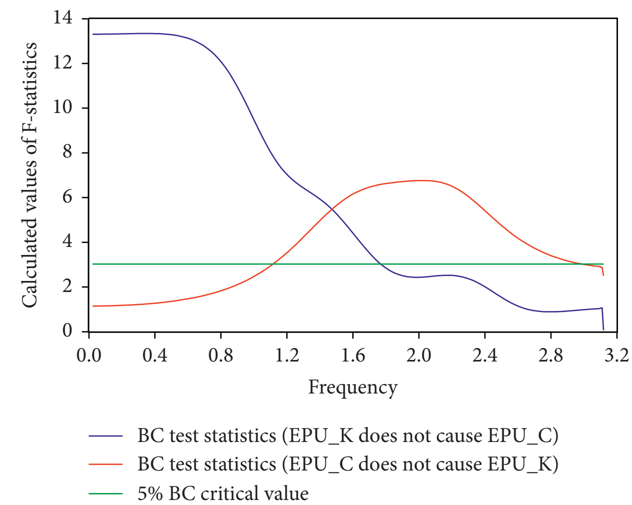

Figure 4: The frequency domain causality between China's and Korea's EPU indices. The part of the lines above the critical value-line indicates rejection of the null hypothesis of no Granger causality.

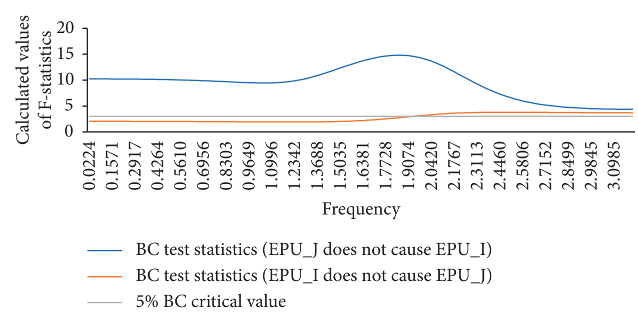

Figure 5: The frequency domain causality between Japan's and India's EPU indices. The part of the lines above the critical value-line indicates rejection of the null hypothesis of no Granger causality.

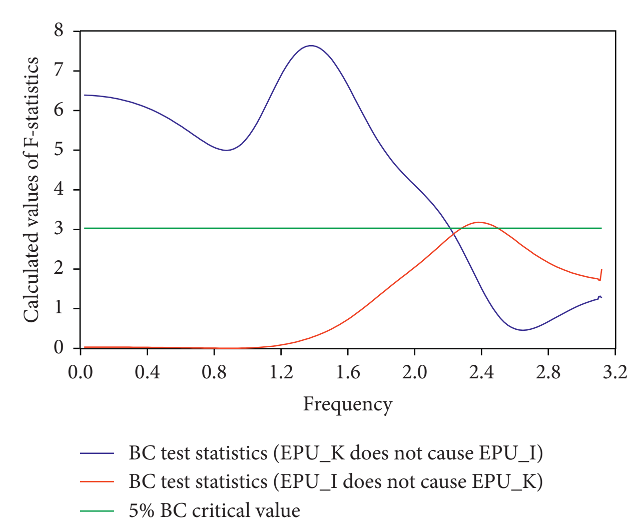

Figure 6: The frequency domain causality between Korea's and India's EPU indices. The part of the lines above the critical value-line indicates rejection of the null hypothesis of no Granger causality.

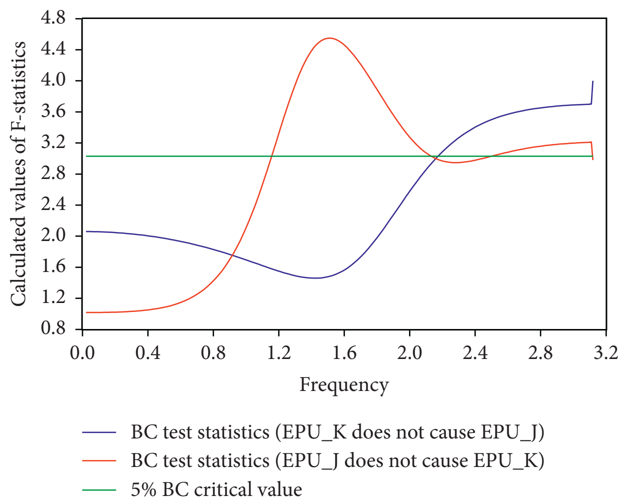

Figure 7: The frequency domain causality between Korea's and Japan's EPU indices. The part of the lines above the critical value-line indicates rejection of the null hypothesis of no Granger causality.

5. Conclusion

We have investigated the linkages between EPU pairs of four Asian countries from January 1997 to April 2020 by examining the cointegration, asymmetric cointegration, and causal relationship in the frequency domain between the EPU pairs of China, India, Japan, and South Korea, allowing for asymmetric adjustments towards long-run equilibrium. The Engle-Granger cointegration test reveals the existence of long-run relationships between the EPU pairs. Because the Engle-Granger cointegration lacks a threshold adjustment in the long-run, we employed the TAR and M-TAR models, following Enders and Siklos [24], to determine the asymmetric response of each EPU in the combination of China-India, China-Japan, China-Korea, India-Japan, India-Korea, and Japan-Korea EPU models. The TAR and M-TAR models support the threshold adjustment between the EPU pairs, which further discloses asymmetries in the EPU model adjustment process. Though the null hypothesis of no cointegration was rejected for both TAR and M-TAR models, the null hypothesis of symmetric adjustment was not rejected for all models with TAR specification with zero threshold value. However, for the M-TAR model with nonzero threshold values, the symmetric adjustment null hypothesis was rejected for China-India and India-Japan EPU pairs, indicating asymmetry in the adjustment of positive and negative divergence from the long-run equilibrium. We, therefore, estimated the M-TVECM using the M-TAR specification. The findings show that the EPUs influence each other in the short run and the threshold error term shows the speed of adjustment for positive deviation which is faster than the negative deviations for all models, except the case where Korea's EPU is the dependent variable where the speed of adjustment for negative deviations is more rapid than the adjustment of positive deviations. In all, apart from the Korea's EPU that responds quickly to the decline in other EPUs in the short run, the remaining EPUs respond quickly to a rise in value of the other EPU indices.

In addition, the bivariate analysis to establish long- and short-run relationship between the EPU pairs of the countries in the frequency domain reveals both unidirectional and bidirectional Granger causality between the EPU pairs. The findings suggest a unidirectional causality between China-India, China-Japan, and India-Korea EPU pairs where India and Japan Granger cause China's EPU in the long and short run while Korea's EPU Granger causes India's EPU in the long run, indicating that both India and Japan influence movement in China's EPU and Korea's EPU which, on one hand, influence the movement in India's EPU. However, the bidirectional causality between China-Korea pair in the long and short run exists. The findings also reveal bidirectional causality between India-Japan and Japan-Korea EPU pairs in the long and short run showing that each EPU influences movement in the other EPU in the pair in either long run or short run or both.

The linkages and comovements between the EPU pairs of countries established in our study provide policy implications to the policymakers and local and international investors of these countries, as well as the countries within the Asian subregion. Heightened economic policy inconsistency spawns fear in investors, leading to “wait” and “see” attitudes which can “impede business prospects and households' consumption and this can threaten all facets of the economy including weakened stock market performance, increased unemployment rate, volatile financial market, rising inflation, etc. Therefore, economic policymakers should be aware of the potential EPU linkages among countries so that prudent measures could be put in place to instill confidence of a growing economy in the investment community.

Data Availability

The economic policy uncertainty data were supplied by https://www.policyuncertainty.com/about.html and economic policy uncertainty is under license and so cannot be made freely available.

Conflicts of Interest

The authors declare that they have no conflicts of interest.

References

[1] N. Bloom, “The impact of uncertainty shocks,” *Econometrica*, vol. 77, no. 3, pp. 623–685, 2009.

[2] G. Caggiano, E. Castelnuovo, and G. Nodari, *Uncertainty and Monetary Policy in Good and Bad Times*, Melbourne Institute, Melbourne, Australia, 2017.

[3] B. S. Bernanke, “Irreversibility, uncertainty, and cyclical investment,” *The Quarterly Journal of Economics*, vol. 98, no. 1, pp. 85–106, 1983.

[4] C. D. Carroll, “Buffer-stock saving and the life cycle/permanent income Hypothesis,” *The Quarterly Journal of Economics*, vol. 112, no. 1, pp. 1–55, 1997.

[5] N. Bloom, S. Bond, and J. V. Reenen, *The Dynamics of Investment under Uncertainty*, IFS Working Papers from Institute for Fiscal Studies, London, UK, 2001.

[6] R. Bansal and A. Yaron, "Risks for the long run: a potential resolution of asset pricing Puzzles," *The Journal of Finance*, vol. 59, no. 4, pp. 1481–1509, 2004.

[7] S. Gilchrist, J. W. Sim, and E. Zakrajšek, *Uncertainty, Financial Frictions, and Investment Dynamics*, National Bureau of Economic Research, Cambridge, MA, USA, 2014.

[8] D. Caldara, C. Fuentes-Albero, S. Gilchrist, and E. Zakrajšek, "The macroeconomic impact of financial and uncertainty shocks," *European Economic Review*, vol. 88, pp. 185–207, 2016.

[9] A. Sheng, C. S. Ng, and C. Edelmann, *ASIA FINANCE 2020 Framing a New Asian Financial Architecture*, Oliver Wyman, Fung Global Institute, Hong Kong, China, 2020.

[10] C. Christou, J. Cunado, R. Gupta, and C. Hassapis, "Economic policy uncertainty and stock market returns in Pacific-Rim countries: evidence based on a Bayesian panel VAR model," *Journal of Multinational Financial Management*, vol. 40, pp. 92–102, 2017.

[11] F. Balli, G. S. Uddin, H. Mudassar, and S.-M. Yoon, "Cross-country determinants of economic policy uncertainty spillovers," *Economics Letters*, vol. 156, pp. 179–183, 2017.

[12] M. Bahmani-Oskooee and S. Saha, "On the effects of policy uncertainty on stock prices: anasymmetric analysis," *Quantitative Finance and Economics*, vol. 3, no. 2, pp. 412–424, 2019a.

[13] T. C. Chiang, "Economic policy uncertainty, risk and stock returns: evidence from G7 stock markets," *Finance Research Letters*, vol. 29, pp. 41–49, 2019.

[14] E.-C. Chung and D. R. Haurin, "Housing choices and uncertainty: the impact of stochastic events," *Journal of Urban Economics*, vol. 52, no. 2, pp. 193–216, 2002.

[15] H. Yu, "Government policies and housing price instability," *Public Policy Review*, vol. 22, no. 2, pp. 74–115, 2008.

[16] D. Su, X. Li, O.-R. Lobonţ, and Y. Zhao, "Economic policy uncertainty and housing returns in Germany: evidence from a bootstrap rolling window," *Zbornik Radova Ekonomskog Fakulteta U Rijeci: Časopis Za Ekonomsku Teoriju I Praksu/Proceedings of Rijeka Faculty of Economics: Journal of Economics and Business*, vol. 34, no. 1, pp. 43–61, 2016.

[17] J.-H. Jeon, "The impact of asian economic policy uncertainty: evidence from Korean housing market," *The Journal of Asian Finance, Economics and Business*, vol. 5, no. 2, pp. 43–51, 2018.

[18] L. Karnizova and J. Li, "Economic policy uncertainty, financial markets and probability of US recessions," *Economics Letters*, vol. 125, no. 2, pp. 261–265, 2014.

[19] K. Jurado, S. C. Ludvigson, and S. Ng, "Measuring uncertainty," *American Economic Review*, vol. 105, no. 3, pp. 1177–1216, 2015.

[20] H. Mumtaz and K. Theodoridis, "Common and country specific economic uncertainty," *Journal of International Economics*, vol. 105, pp. 205–216, 2017.

[21] I. O. Olanipekun, H. Güngör, and G. Olasehinde-Williams, "Unraveling the causal relationship between economic policy uncertainty and exchange market pressure in BRIC countries: evidence from bootstrap panel Granger causality," *SAGE Open*, vol. 9, no. 2, 2 pages, 2019.

[22] P. Alessandri and H. Mumtaz, "Financial regimes and uncertainty shocks," *Journal of Monetary Economics*, vol. 101, pp. 31–46, 2019.

[23] N. S. Balke and T. B. Fomby, "Threshold cointegration," *International Economic Review*, vol. 38, pp. 627–645, 1997.

[24] W. Enders and P. L. Siklos, "Cointegration and threshold adjustment," *Journal of Business & Economic Statistics*, vol. 19, no. 2, pp. 166–176, 2001.

[25] T. Ito and P.-L. Vézina, "Production fragmentation, upstreamness, and value added: evidence from Factory Asia 1990-2005," *Journal of the Japanese and International Economies*, vol. 42, pp. 1–9, 2016.

[26] M. Helble and B.-L. Ngiang, "From global factory to global mall? East Asia's changing trade composition and orientation," *Japan and the World Economy*, vol. 39, pp. 37–47, 2016.

[27] J. Aizenman and S.-I. Fukuda, "The pacific rim and the global economy: future financial and macro challenges," *Journal of International Money and Finance*, vol. 74, pp. 229–231, 2017.

[28] B. Shepherd, "Mega-regional trade agreements and Asia: an application of structural gravity to goods, services, and value chains," *Journal of the Japanese and International Economies*, vol. 51, pp. 32–42, 2018.

[29] S.-I. Fukuda, "Finance in asia rising: growth and resilience in an uncertain global economy," 2013.

[30] H. Ito and M. Kawai, "Trade invoicing in major currencies in the 1970s-1990s: lessons for renminbi internationalization," *Journal of the Japanese and International Economies*, vol. 42, pp. 123–145, 2016.

[31] Y. Wanhai, Y. Guo, H. Zhu, and Y. Tang, "Oil price shocks, economic policy uncertainty and industry stock returns in China: asymmetric effects with quantile regression," *Energy Economics*, vol. 68, pp. 1–18, 2017.

[32] R. Li, S. Li, D. Yuan, and K. Yu, "Does economic policy uncertainty in the U.S. influence stock markets in China and India? Time-frequency evidence," *Applied Economics*, vol. 52, no. 39, p. 4300, 2020.

[33] T. C. Chiang, "Economic policy uncertainty and stock returns: evidence from the Japanese market," *Quantitative Finance and Economics*, vol. 4, no. 3, pp. 430–458, 2020.

[34] M. Balcilar, Z. A. Ozdemir, H. Ozdemir, and M. Wohar, "Transmission of US and EU economic policy uncertainty shock to Asian economies in bad and good times," *Discussion Paper Series, IZA Instittute of Labor Economics*, vol. 18, 2020.

[35] http://www.federalreserve.gov/monetarypolicy/fomcminutes20091216.htmFederal Open Market Committee, "Minutes of the December 2009 Meeting," 2009, HYPERLINK http://www.federalreserve.gov/monetarypolicy/fomcminutes20091216.htm.

[36] S. Klößner and R. Sekkel, "International spillovers of policy uncertainty," *Economics Letters*, vol. 124, no. 3, pp. 508–512, 2014.

[37] J. Brogaard and A. Detzel, "The asset-pricing implications of Government economic policy uncertainty," *Management Science*, vol. 61, no. 1, pp. 3–18, 2015.

[38] M. Arouri, C. Estay, C. Rault, and D. Roubaud, "Economic policy uncertainty and stock markets:long run evidence from the us," *Finance Research Letters*, vol. 18, pp. 136–141, 2016.

[39] A. M. Adam, "Susceptibility of stock market returns to international economic policy: evidence from effective transfer entropy of Africa with the implication for open innovation," *Journal of Open Innovation: Technology, Market, and Complexity*, vol. 6, no. 3, p. 71, 2020.

[40] E. Asafo-Adjei, D. Agyapong, S. K. Agyei, S. Frimpong, R. Djimatey, and A. M. Adam, "Economic policy uncertainty and stock returns of Africa: a wavelet coherence analysis," *Discrete Dynamics in Nature and Society*, vol. 2020, Article ID 8846507, 2020.

[41] T. C. Chiang, “Financial risk, uncertainty and expected returns: evidence from Chinese equity markets,” *China Finance Review International*, vol. 9, no. 4, pp. 425–454, 2019.
[42] Ľ. Pástor and P. Veronesi, “Political uncertainty and risk premia,” *Journal of Financial Economics*, vol. 110, no. 3, pp. 520–545, 2013.
[43] L. Liu and T. Zhang, “Economic policy uncertainty and stock market volatility,” *Finance Research Letters*, vol. 15, pp. 99–105, 2015.
[44] I.-C. Tsai, “The source of global stock market risk: a viewpoint of economic policy Uncertainty,” *Economic Modelling*, vol. 60, pp. 122–131, 2017.
[45] L. P. Hansen, T. J. Sargent, and T. D. Tallarini, “Robust permanent income and pricing,” *Review of Economic Studies*, vol. 66, no. 4, pp. 873–907, 1999, https://econpapers.repec.org/article/ouprestudhttps://econpapers.repec.org/article/ouprestud/.
[46] L. Pástor and P. Veronesi, “Uncertainty about government policy and stock prices,” *The Journal of Finance*, vol. 67, no. 4, pp. 1219–1264, 2012.
[47] V. Sum, “The ASEAN stock market performance and economic policy uncertainty in the United States,” *Economic Papers: A Journal of Applied Economics and Policy*, vol. 32, no. 4, pp. 512–521, 2013.
[48] H. Chuliá, R. Gupta, J. M. Uribe, and M. E. Wohar, “Impact of US uncertainties on emerging and mature markets: evidence from a quantile-vector autoregressive approach,” *Journal of International Financial Markets, Institutions and Money*, vol. 48, pp. 178–191, 2017.
[49] N. B. Trung, “The spillover effect of the US uncertainty on emerging economies: a Panel VAR Approach,” *Applied Economics Letters*, vol. 26, no. 3, pp. 210–216, 2019.
[50] S. Bhattarai, A. Chatterjee, and W. Y. Park, “Global spillover indices of US uncertainty,” *Journal of Monetary Economics*, vol. 19, 2019.
[51] S. S. Akadiri, A. A. Alola, and G. Uzuner, “Economic policy uncertainty and tourism: evidence from the heterogeneous panel,” *Current Issues in Tourism*, vol. 23, no. 20, p. 2507, 2019.
[52] S. Adams, F. Adedoyin, E. Olaniran, and F. V. Bekun, “Energy consumption, economic policy uncertainty and carbon emissions; causality evidence from resource rich economies,” *Economic Analysis and Policy*, vol. 68, pp. 179–190, 2020.
[53] P. Luk, M. Cheng, P. Ng, and K. Wong, “Economic policy uncertainty spillovers in small open economies: the case of Hong Kong,” *Pacific Economic Review*, vol. 25, no. 1, 2018.
[54] S. E. Cekin, A. K. Pradhan, A. K. Tiwari, and R. Gupta, “Measuring co-dependencies of economic policy uncertainty in Latin American countries using vine copulas,” *The Quarterly Review of Economics and Finance*, vol. 14, 2019.
[55] L. Bai, X. Zhang, Y. Liu, and Q. Wang, “Economic risk contagion among major economies: new evidence from EPU spillover analysis in time and frequency domains,” *Physica A: Statistical Mechanics and its Applications*, vol. 535, 2019 https://ideas.repec.org/s/eee/phsmap.html.
[56] R. F. Engle and C. W. J. Granger, “Co-integration and error correction: representation, estimation, and testing,” *Econometrica*, vol. 55, no. 2, pp. 251–276, 1987.
[57] K. S. Chan, “Consistency and limiting distribution of the least squares estimator of a threshold autoregressive model,” *The Annals of Statistics*, vol. 21, no. 1, pp. 520–533, 1993.
[58] J. Geweke, “Measurement of linear dependence and feedback between multiple time Series,” *Journal of the American Statistical Association*, vol. 77, no. 378, pp. 304–313, 1982.
[59] Y. Hosoya, “The decomposition and measurement of the interdependency between second-order stationary processes,” *Probability Theory and Related Fields*, vol. 88, no. 4, pp. 429–444, 1991.
[60] J. Breitung and B. Candelon, “Testing for short- and long-run causality: a frequency-domain approach,” *Journal of Econometrics*, vol. 132, no. 2, pp. 363–378, 2006.
[61] S. R. Baker, N. Bloom, and S. J. Davis, “Measuring economic policy uncertainty∗,” *The Quarterly Journal of Economics*, vol. 131, no. 4, pp. 1593–1636, 2016.
[62] D. A. Dickey and W. A. Fuller, “Likelihood ratio statistics for autoregressive time series with a unit root,” *Econometrica*, vol. 49, no. 4, pp. 1057–1072, 1981.
[63] P. C. B. Phillips and P. Perron, “Testing for a unit root in time series regression,” *Biometrika*, vol. 75, no. 2, pp. 335–346, 1988.
[64] D. Kwiatowski, P. C. B. Peter, P. Schmidt, and Y. Shin, “Testing the null hypothesis of stationarity against the alternative of a unit root: how sure are we that economic time series have a unit root?,” *Journal of Econometrics*, vol. 54, pp. 159–178, 1992.
[65] J. G. McKinnon, “Numerical distribution functions for unit root and cointegration tests,” *Journal of Applied Econometrics*, vol. 11, no. 6, pp. 601–618, 1996.
[66] G. W. Schwert, “Tests for unit roots: a Monte Carlo investigation,” *Journal of Business and Economic Statistics*, vol. 7, pp. 147–159, 1989.
[67] G. Kapetanios, A. Snell, and Y. Shin, “Testing for unit root in the nonlinear STAR framework,” *Journal of Econometrics*, vol. 112, pp. 359–379, 2003.