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Nonlinear Cointegration and Asymmetric Adjustement between Economic policy uncertainty and Gold price: Evidence from the United States

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Abstract

The main objective of this paper is to estimate the interaction between gold price and US- EPU by using monthly data during the period span from January 1997 to September 2020. The threshold cointegration approach focus on TAR, C-TAR, M-TAR and consistent-MTAR is employed. Results indicate the evidence of asymmetry in the adjustment process to equilibrium. Unidirectional causality is reported between variables. In addition, gold price became cointegrated with economic uncertainty. The adjustment mechanism is asymmetric and the speed of adjustment to the equilibrium is different when the last equilibrium error has different signs.

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1. Introduction

As part of portfolio diversification, investors use gold as a hedging strategy against risk, especially in times of financial and economic crises (Lau et al., 2017; O'Connor et al., 2015). Gold is characterized by monetary and financial functions. In February 1973 and specifically since the collapse of the Bretton Woods system, the function of gold has weakened. Wang et al (2016) indicate that the financial asset function is very important. Moreover, gold is crucial for governments and central banks as it can be used as a reserve to defend the value of their currencies (Baur and McDermott, 2010).

The dynamic relationships between gold and other financial markets have attracted a special attention from both practitioners and academics. A strong relationship between them would have important implications for international capital budgeting decisions and economic policies because negative shocks affecting one market may be quickly transmitted to another through contagious effects. This issue has become more critical with the occurrence of recent black swan events such as the US 2007 subprime crisis. Future uncertainties and the "black swan theory" have gained momentum.

Theoretically, Bredin et al. (2015), Joy (2011) and Reboredo (2013) suggest that gold protects against the risk of financial instruments and also protects against the risk of inflation (Lucey et al., 2017). In the context of economic uncertainty, (Bialkowski et al. (2015)) suggest that gold is considered as a tool for hedging against uncertainty in the global financial system.

In times of political and economic uncertainty, there is a great attractiveness of gold to investors and this leads various theorists such as Balcilar et al (2016) as well as Li and Lucey (2017) to study the relationship between measures of uncertainty and gold. Jones and Sackley (2016) incorporate an index of the United States and European economic policy uncertainty into a gold-pricing model and find that gold prices are positively related to EPU.

Various empirical studies have been interested in the interactions between the price of gold and other macroeconomic and financial variables. By considering the effects of level and volatility. For example, Bekiros et al. (2017) use a GARCH-based on copula methodology and a wavelet approach to study the hedging and diversification roles of gold for BRICS stock markets. Results indicate the evidence of two-way causality linkages, suggesting that gold acts as an agent of equity. On the other hand, Sarwar (2017) considers the links between the implied volatilities of US stock markets and the gold; the author disregards the relationships in large emerging markets.

Recently, Sarwar and Khan (2017) use GARCH models and the Granger causality test, and they indicate that intensified US market uncertainty reduces emerging market returns in Latin America and increases the variance of returns. Bouri et al. (2017a) find the presence of nonlinearity and significant co-integration in the interaction between the implied volatilities of gold, crude oil and Indian equities. Bouri et al. (2017b) examine the short- and long-term causality between the implied volatility of gold and that of India and China. The authors show evidence of a feedback effect. Yang and Zhou (2017) identify the significant impact of quantitative easing on the implied volatility spillovers across US Treasury bonds commodities and developed stock market indices.

More recently, Rehman et al. (2018) use the structural vector autoregression (SVAR) model proposed by Kilian and Park (2009) and examine the effect of oil price shocks on precious metal returns. They detect the variability in the effects through a rolling window impulse response function by extending the dynamic connectedness approach of Diebold and Yılmaz (2014) using the structural forecast error variance decomposition. They report a time-varying effect of the disintegrated structural oil shocks on precious metal returns with a significant increase during the global financial crisis period of 2008–2009.

Direct and indirect relationships between gold prices and economic policy uncertainty highlight the importance of conducting research that examines the interaction between economic uncertainty and gold prices. In this context, the non-parametric causality-in-quantiles approach is employed by Balcilar et al. (2016a) to study economic policy uncertainty and gold prices, return and volatility. Many uncertainty measures are used, which include those of Baker et al. (2016), Rossi and Sekhposyan (2015), and Jurado et al. (2015), and mixed results are reported. Their empirical results of monthly and daily data indicated that uncertainty measures affect gold prices, return and volatility. In contrast, the results of quarterly data showed weak causality and were significant for gold volatility only.

On the other hand, Raza et al. (2018) use the standard linear Granger causality test and the nonparametric causality-in-quantiles approach to examine the association between economic policy uncertainty and gold prices. The period span from January, 1995 to March, 2017 with monthly frequency data. Results indicate that the standard linear Granger causality test shows that no causal association exists between economic policy uncertainty and gold prices. The nonparametric causality-in-quartiles test shows the rejection of null hypothesis, which implies that economic policy uncertainty, causes gold prices in all the examined countries.

The main objective of this paper is to study the interaction between gold price and economic policy uncertainty by considering the United States economy. We employed the nonlinear cointegration such as the threshold effect focus on TAR model, consistent TAR, momentum TAR and consistent-MTAR. We examine the long-term relationship between gold and EPU. We use the Enders and Siklos (2001) asymmetric cointegration model to analyze the long-run asymmetric equilibrium relationship between uncertainty and gold price. To be specific, the adjustment coefficient of the error correction term is different when the equilibrium error is positive from when it is negative.

This paper is organized as follows: section 2 discusses the data and empirical methodology, section 3 presents the preliminary analysis, section 4 presents the empirical results and section 5 concludes the paper.

2. Data and empirical methodology

In this paper, we use two variables, namely economic policy uncertainty (EPU) and gold prices from United States at monthly frequency during the period span from January, 1997 to September, 2020. The main objective is to study the nonlinear cointegration and asymmetric adjustment between variables. The gold price is extracted from the World Bank Commodity Price Data and the EPU is sourced from policyuncertainty.com.

In recent years, threshold cointegration has been increasingly used in price transmission studies. Cointegration has been frequently employed to study the interaction between price variables. The two methods of cointegration are Johansen and Engle-Granger two-step approaches. Both of them assume symmetric relationship between variables. Balke and Fomby (1997) used a two-step approach to examine threshold cointegration on the basis of the approach developed by Engle and Granger (1987). Enders and Granger (1998) and Enders and Siklos (2001) further generalize the standard Dickey-Fuller test by allowing for the possibility of asymmetric movements in time-series data. This makes it possible to test for cointegration without maintaining the hypothesis of a symmetric adjustment to a long-term equilibrium. Thereafter, the method has been widely applied to analyze asymmetric price transmission. The conventional tests of cointegration such as Engle and Granger (1987) are residual-based test that analyze the validity of long-run relationships between gold price and EPU by estimating the following model:

$$Y_t = \beta_0 + \beta_1 X_t + \varepsilon_t \tag{1}$$

Where Y_t is the gold price of United states at time t and X_t is the EPU. ε_t is the residual in equation (1) and β_0 and β_1 are coefficients.

The implicit assumption of linear and symmetric adjustment (Engle and Granger, 1987) is problematic. Enders and Siklos (2001) proposed a two-regime threshold cointegration approach to entail asymmetric adjustment in cointegration analysis. They argued that the Engle-Granger cointegration test is likely to lead to misspecification errors when the adjustment of the error correction term is asymmetric. They remedy this error by expanding the Engle-Granger two-step cointegration test to incorporate an asymmetric error correction term. In the next step, we determine whether or not the disturbance term ε_t is stationary by considering an asymmetric test methodology in the form of Threshold Autoregressive (TAR) cointegration model as proposed by Enders and Granger (1998) and Enders and Siklos (2001).

The equation of a TAR process is:

$$\Delta\varepsilon t = I_t\, \rho_1(\varepsilon_{t-1}-\tau) + (1-I_t)\, \rho_2\, (\varepsilon_{t-1}-\tau) + \mu_t \tag{2}$$

Where ρ_1, ρ_2 are coefficients, τ is the value of the threshold, μ_t is a white-noise disturbance and I_t is the Heaviside indicator such that:

$$I_t = \begin{cases} 1 \; if\ \varepsilon t - 1 \geq \tau \\ 0 \; if\ \varepsilon t - 1 < \tau \end{cases} \tag{3}$$

Under the null hypothesis of no cointegration between the variables, the -statistic for the null hypothesis ρ_1=ρ_2=0 has a nonstandard distribution. Rejecting this assumption means that Eq. (2) is an attractor such that the equilibrium value of the $\{\varepsilon_t\}$ is τ. The adjustment process is ($\rho_1\varepsilon_{t-1}-\tau$) if the lagged value of ε_t is above its long-run equilibrium value, while if the lagged value of ε_t is below its long-run equilibrium value, the adjustment is $\rho_2(\varepsilon_{t-1}-\tau)$. If $-1<|\rho_1|<|\rho_2|<0$, negative discrepancies will be more persistent than positive discrepancies. Moreover, Tong (1983) showed that the OLS estimates of ρ_1 and ρ_2 have an asymptotic multivariate normal distribution if the sequence $\{\varepsilon_t\}$ is stationary. Therefore, if the null assumption $\rho_1=\rho_2=0$ is rejected, it is possible to test for symmetric adjustment (i.e., $\rho_1=\rho_2$) using a standard F-test. Rejecting both the null assumptions $\rho_1=\rho_2=0$ and $\rho_1=\rho_2$ indicates the existence of threshold cointegration and asymmetric adjustment.

Since the exact nature of the nonlinearity may not be known, Enders and Siklos (2001) consider another kind of asymmetric cointegration test methodology that allows the adjustment to be contingent on the change in ε_{t-1} (i.e., $\Delta\varepsilon_{t-1}$) instead of the level of ε_{t-1}. In this case, the Heaviside indicator of Eq. (3) becomes.

$$I_t = \begin{cases} 1 \; if \; \Delta\varepsilon t-1 \geq \tau \\ 0 \; if \; \Delta\varepsilon t-1 < \tau \end{cases} \tag{4}$$

This specification is especially relevant when the adjustment is such that the series exhibits more "momentum" in one direction than in the other (Thompson, 2006; Kuo and Enders, 2004; Enders and Siklos, 2001; Enders and Granger, 1998). That is, the speed of adjustment depends on whether ε_t is increasing (i.e., widening) or decreasing (i.e., narrowing). According to Thompson (2006), among others, if $|\rho_1|<|\rho_2|$, then increases in ε_t tend to persist whereas decreases revert back to the threshold quickly. The resulting model is called momentum-threshold autoregressive (M-TAR) cointegration model. The TAR model captures asymmetrically deep movements if, for instance, positive deviations are more prolonged than negative deviations. The M-TAR model allows the autoregressive decay to depend on $\Delta\varepsilon_{t-1}$.

As such, the M-TAR specification can capture asymmetrically "sharp" movements in $\{\varepsilon_t\}$ sequence (Caner and Hansen, 2001).

In both the TAR and M-TAR cointegration processes, the null assumption of $\rho_1=\rho_2=0$ could be tested, while the null hypothesis of symmetric adjustment may be tested by the restriction $\rho_1=\rho_2$. Generally, there is no presumption to whether to use TAR or M-TAR specifications. Thus, it is recommended to select the adjustment mechanism by a model selection criterion such as AIC or BIC. Furthermore, if the errors in Eq. (2) are serially correlated, it is possible to use the augmented form of the test:

$$\Delta\varepsilon_t = \mathrm{I}_t\, \rho_1(\varepsilon_{t-1}-\tau) + (1-\mathrm{I}_t)\, \rho_2\, (\varepsilon_{t-1}-\tau) + \textstyle\sum_{i=1}^{P} \varphi_i\, \Delta\varepsilon_{t-i} + v_t \tag{5}$$

To use the tests, we first regress ε_t on a constant and call the residuals, $\{\widehat{\varepsilon_t}\}$ which are the estimates of $(\varepsilon_{t-1}-\tau)$. In a second step, we set the indicator according to Eq. (3) or Eq. (4) and estimate the following regression:

$$\Delta\widehat{\varepsilon_t} = \mathrm{I}_t\, \rho_1(\widehat{\varepsilon_{t-1}}-\tau) + (1-\mathrm{I}_t)\, \rho_2\, (\widehat{\varepsilon_{t-1}}-\tau) + \textstyle\sum_{i=1}^{P} \varphi_i\, \Delta\widehat{\varepsilon_{t-1}} + v_t \tag{6}$$

The number of lags p is specified to account for serially correlated residuals and it can be selected using AIC, BIC, or Ljung-Box Q test. In several applications, there is no reason to expect the threshold to correspond with the attractor (i.e., $\tau=0$). In such circumstances, it is necessary to estimate the value of along with the values of ρ_1 and ρ_2. A consistent estimate of the threshold t can be obtained by adopting the methodology of Chan (1993). A super consistent estimate of the threshold value can be attained with several steps. First, the process involves sorting in ascending order the threshold variable, i.e., $\widehat{\varepsilon_{t-1}}$ for the TAR model or the

$\Delta\widehat{\varepsilon_{t-1}}$ for the M-TAR model. Second, the potential threshold values are determined. If the threshold value is to be meaningful, the threshold variable must actually cross the threshold value (Enders, 2004). Thus, the threshold value τ should lie between the maximum and minimum values of the threshold variable.

The equilibrium correction specification (ECM) of Engle and Granger (1987) assumes that the adjustment process due to disequilibrium among the variables is symmetric. In order to incorporate asymmetries, two extensions on the ECM model have been made. Error correction terms and first differences on the variables are decomposed into positive and negative values, as proposed by Granger and Lee (1989). The second extension adds the threshold cointegration mechanism to the Granger and Lee (1989) approach. The resulting asymmetric error correction model with threshold cointegration has the form:

$$\Delta X_t = \theta_x + \delta_x{}^{+} Z_{t-1}{}^{+} + \delta_x{}^{-} Z_{t-1}{}^{-} + \sum_{j=1}^{J} \alpha_{xj}{}^{+} \Delta X_{t-j}{}^{+} + \sum_{j=1}^{J} \alpha_{xj}{}^{-} \Delta X_{t-j}{}^{-} + \sum_{j=1}^{J} \beta_{xj}{}^{+} \Delta Y_{t-j}{}^{+} + \sum_{j=1}^{J} \beta_{xj}{}^{-} \Delta Y_{t-j}{}^{-} + v_{x,t} \quad (7)$$

And

$$\Delta Y_t = \theta_y + \delta_y{}^{+} Z_{t-1}{}^{+} + \delta_y{}^{-} Z_{t-1}{}^{-} + \sum_{j=1}^{J} \alpha_{yj}{}^{+} \Delta X_{t-j}{}^{+} + \sum_{j=1}^{J} \alpha_{yj}{}^{-} \Delta X_{t-j}{}^{-} + \sum_{j=1}^{J} \beta_{yj}{}^{+} \Delta Y_{t-j}{}^{+} + \sum_{j=1}^{J} \beta_{yj}{}^{-} \Delta Y_{t-j}{}^{-} + v_{y,t} \quad (8)$$

Where k = {1, 2}, $Z_{t-1}{}^{+} = \mathrm{I}_t\, \widehat{\varepsilon_{t-1}}$ *and* $Z_{t-1}{}^{-} = (1 - \mathrm{I}_{t})\widehat{\varepsilon_{t-1}}$

3. Preliminary analysis

Table 1 reports summary statistics of gold price and economic policy uncertainty (EPU). The highest mean and standard deviation are observed for gold during the period. Asymmetry is measured by the values of skewness and kurtosis is a measurement for flatted distribution. We see that the two variables have a positive skewness. The Jarque-Bera test statistic rejects the null hypothesis of normality.

Table I. Descriptive statistics of gold and EPU

variables	GOLD	EPU
Mean	879.9235	127.5997
Maximum	1968.630	503.963
Minimum	256.0800	44.7827
SD	500.722	62.7238
Skewness	0.1514	2.2600
Kurtosis	1.5894	11.0948
Jarque-Bera	24.7182***	1020.750***
	0.0000	0.0000
Observations	285	285

Note: *, ** and *** denote the significance at 10%, 5% and 1% levels.

Figure1 display the time series plots for the gold price and economic policy uncertainty. We see that EPU and gold have an evident co-movement in general, which reveals a high possibility of cointegration between these series. In addition, the pair of series display divergent movement indicating possible nonlinear cointegration. Table 2 shows the results of the stationarity tests, namely the ADF and PP. The observation of the results indicates that all the series are stationary in first difference. We conclude that gold and EPU are integrated processes of order one I (1), or unit root processes.

4. Empirical results

Results of the Engle and Granger (1987) cointegration tests are reported in table 3. Results provide evidence for the alternative hypothesis of linear cointegration. Indeed, the parameters β_0 and β_1 are statistically significant. To study the possibility of asymmetric transmission mechanism between gold price and EPU, we conduct a nonlinear cointegration analysis by using the threshold auto-regression models. Four models are used in this paper such as the TAR with $\tau=0$, consistent TAR with τ estimated, Momentum -TAR with $\tau=0$ and consistent M-TAR with τ estimated.

To study the existence of a serial correlation in the residual series, we choose an optimal lag for each model. For the empirical diagnostic analysis, we focus on three information criteria namely AIC, SBIC and L-Jung Box statistics at different orders 4, 8 and 12. The value of the threshold τ is unknown and has to be estimated along the values of ρ_1 and ρ_2. We follow the Chan's (1993) method to estimate the threshold values for consistent TAR and M-TAR models.

Table 4 reports the empirical results of the threshold cointegration tests for the TAR, consistent-TAR, momentum TAR and consistent-MTAR. Through the four nonlinear models, the results indicate the rejection of the null hypothesis of threshold cointegration ($\rho1=\rho2=0$) for the gold-EPU pair by considering the consistent-MTAR model. This result confirms the evidence of a cointegrating relationship between the gold price and EPU. In this case we can examine whether their adjustment coefficients are different across positive and negative errors. This procedure serves to verify the evidence of an asymmetric cointegration through the hypothesis H0: $\rho1=\rho2$. If the two previous tests reject the null assumption, so asymmetry test makes sense. Based on information criteria AIC and SBIC and L-Jung Box statistics, we observe that the C-MTAR is the most applicable model for variables' adjustment to long-run equilibrium for the pair Gold-EPU.

Figure 2 illustrates the variations of the SSE for the C-MTAR model considering a lag of 8. By observing the gold-EPU pair, we see that the lowest SSE for the consistent-MTAR model is 7700.000 at the threshold value of -132.205. The consistent-MTAR model is the best model characterized by the lowest AIC statistic of 3586.885and BIC statistic of 3626.709.

As shown in Table 4, we found limited evidence of asymmetric transmission between gold price and EPU. Therefore, gold price became cointegrated with the uncertainty, the adjustment mechanism is asymmetric and the speed of adjustment to the equilibrium is different when the last equilibrium error has different signs. This means that the change in the equilibrium error has a different impact on the adjustment speed to the new equilibrium.

Considering the Gold-EPU pair, we observe for the C-MTAR model that the F test relating to the null hypothesis of absence of cointegration admits a statistic of 3.771 which is significant at a level of 5%. This result indicates that gold and EPU are cointegrated with an adjustment threshold. In addition, the F statistic for the null hypothesis of symmetric price transmission has a value of 6.168 and it is significant at the 5% level. Therefore, the adjustment process is asymmetric when gold and EPU adjust to achieve the long-term equilibrium.

In order to investigate the movement of the gold price and EPU series in a long-run equilibrium relationship, we analyze the asymmetric error correction model. The results of the

C-MTAR model are reported in Table 5. Diagnostic analyses on the residuals with AIC, BIC and Ljung-Box Q statistics select a lag of four for the model. The consistent-MTAR model is the best from the threshold cointegration analyses and the error correction terms are constructed using Equation (4) and Equation (6). The result indicates that gold is cointegrated with EPU and it also exhibits asymmetric adjustments. In addition, the short-term equilibrium adjustment process mainly occurs with EPU since $\delta+ = \delta-$.

For regimes with positive shocks (EPU is higher than gold), the adjustment coefficient for gold is -0.0035 and 0.0149 for EPU, which means that, in the next period, gold price will go up and the price deviation will increase. Considering regimes with negative shocks (gold price is lower than EPU), the adjustment coefficient for gold is 0.0178 and 0.0587 for EPU, which means that, in the next period, gold price will go down. The adjusted R-squared value is 0.080 for the gold and 0.188 for EPU. Moreover, the AIC and BIC statistics for EPU are both larger than those for the gold price. This means that the model specification is better fitted on the EPU. The Granger causality between this pair is analyzed by the F-tests. The F-statistic of 2.002 reveals that gold does Granger cause EPU. The F-statistic of 0.976 also indicates that EPU does not Granger cause gold.

Table II. Unit root tests results

	ADF-Level		ADF-difference		I(d)	PP-Level		PP-difference		I(d)
Variables	t-statistic	p-value	t-statistic	p-value		t-statistic	p-value	t-statistic	p-value	
Gold Price	-1.6781	0.7586	-13.7897***	0.0000	I(1)	-1.7079	0.7455	-13.7731***	0.0000	I(1)
EPU	-0.6979	0.4139	-13.8344***	0.0000	I(1)	-1.1996	0.2108	-36.3690***	0.0000	I(1)

Note: *, ** and *** denote the significance at 10%, 5% and 1% levels.

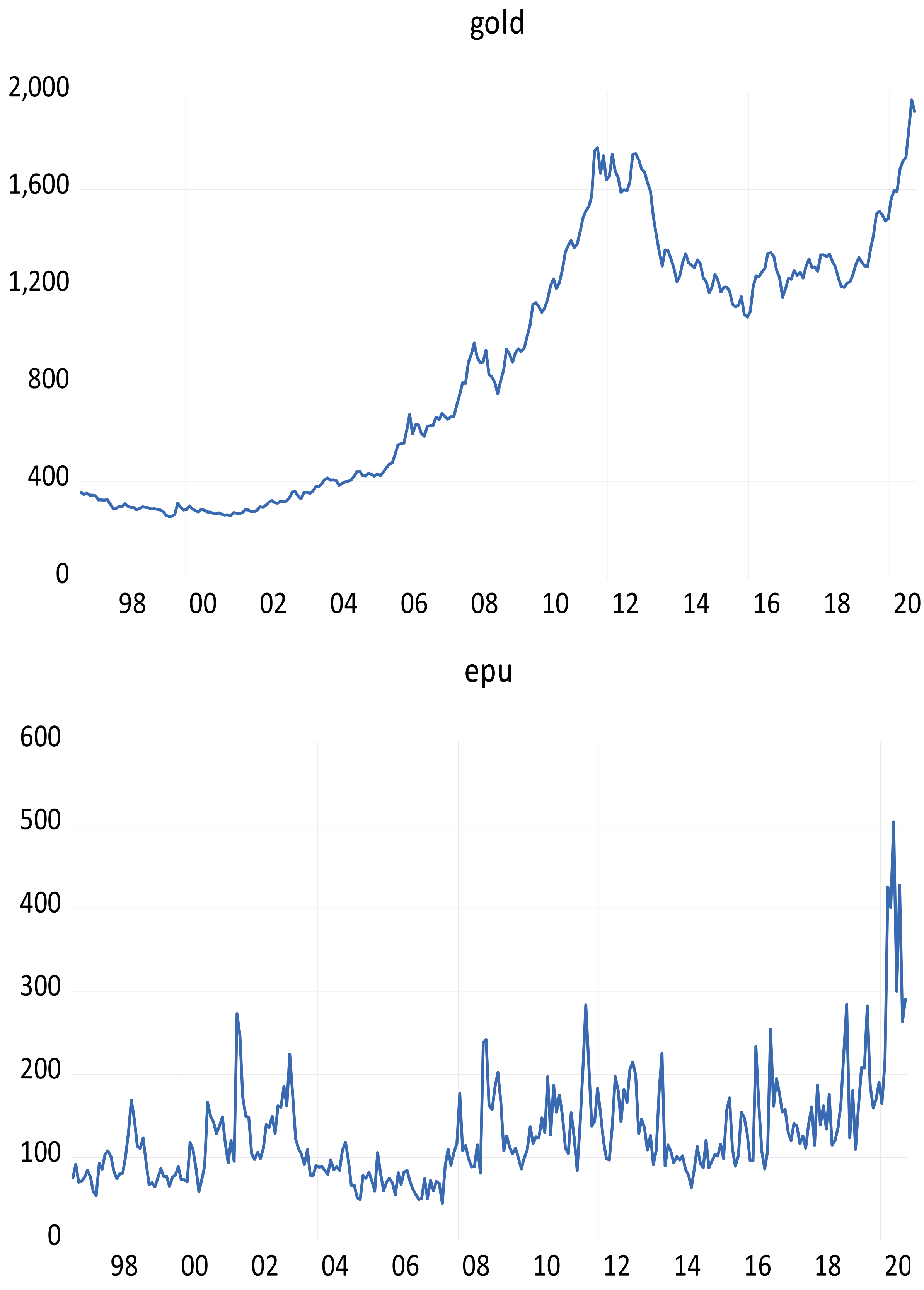

Figure1. Dynamics of monthly Gold price and Economic policy uncertainty

Table III. Linear cointegration results tests (Engle and Granger (1987))

	β_0		β_1	
Pairs of variables	coefficient	p-value	coefficient	p-value
GOLD- EPU	376.176***	0.0000	3.948***	0.0000

Note: *, ** and *** denote the significance at 10%, 5% and 1% levels.

Table IV. Results of the nolinear cointegration tests (threshold)

Pairs of Variables	GOLD-EPU			
	TAR	C-TAR	M-TAR	**C-MTAR**
lags(p)	8	8	8	8
threshold	0	-235.705	0	-132.205
rho1	-0.031	-0.044	-0.005	-0.007
	(-0.959)	(-1.38)	(-0.163)	(-0.273)
rho2	-0.023	-0.01	0.058*	-0.158***
	(-0.73)	(-0.31)	(-1.657)	(-2.745)
total obs	285	285	285	285
coint obs	276	276	276	276
AIC	3593.178	3592.576	3591.762	**3586.885**
BIC	3633.002	3632.4	3631.586	**3626.709**
LB(4)	1	0.999	1	0.999
LB(8)	0.998	0.998	0.999	1
LB(12)	1	1	1	1
No CI: ∅	0.688	0.979	1.375	3.771**
H0: $\rho_1 = \rho_2 = 0$	0.5036	0.3768	0.2545	0.0242
No APT	0.032	0.613	1.401	6.168**
H0: $\rho_1 = \rho_2$	0.857	0.434	0.238	0.014

Note: Number in parentheses are the t-value. *, ** and *** denote the significance at 10%, 5% and 1% levels.

Table V. Results of asymmetric ECM with threshold cointegration

Variables	C-MTAR (lag=4)			
	GOLD		EPU	
	Coefficients	t-statistic	Coefficients	t-statistic
θ	-0.6609	-0.131	13.7851*	2.481
${\alpha_1}^+$	0.2401**	2.705	-0.1730	-1.775
${\alpha_2}^+$	0.0408	0.432	-0.1760	-1.696
${\alpha_3}^+$	0.0835	0.829	-0.4117***	-3.721
${\alpha_4}^+$	0.0905	0.875	-0.0079	-0.070
${\alpha_1}^-$	-0.0446	-0.385	-0.4139**	-3.253
${\alpha_2}^-$	0.0099	0.086	0.0099	0.078
${\alpha_3}^-$	0.0172	0.163	0.0085	0.073
${\alpha_4}^-$	-0.2417*	-2.227	-0.1069	-0.897
${\beta_1}^+$	0.1185	1.159	-0.1347	-1.200
${\beta_2}^+$	-0.3006**	-2.863	-0.0426	-0.370
${\beta_3}^+$	0.1274	1.188	-0.1695	-1.440
${\beta_4}^+$	0.0756	0.715	0.0340	0.293
${\beta_1}^-$	0.2854*	2.163	0.0707	0.488
${\beta_2}^-$	0.3352*	2.521	-0.0551	-0.377
${\beta_3}^-$	-0.0541	-0.407	0.3270*	2.241
${\beta_4}^-$	-0.1697	-1.277	0.0511	0.351
δ^+	-0.0035	-0.458	0.0149	1.750
δ^-	0.0178	1.203	0.0587***	3.609
Diagnostic	**F-stat**	**P-value**	**F-stat**	**P-value**
R-Squared	0.140	-	0.240	-
Adjusted R-Squared	0.080	-	0.188	-
F-Stat	2.356***	0.0000	4.589***	0.0000
AIC	2824.613	-	2877.055	-
BIC	2897.309	-	2949.750	-
Q(4)	0.976	-	0.611	-
Q(8)	0.447	-	0.272	-
Q(12)	0.055	-	0.163	-
Granger Causality Test				
$H_{01}: {\alpha_{yj}}^+ = {\alpha_{yj}}^-$	2.002**	0.047	4.370***	0.000
$H_{02}: {\beta_{xj}}^+ = {\beta_{xj}}^-$	2.807***	0.005	0.976	0.455

Notes: Q_{LB}(P) denote the significance level for the Ljung-Box Q statistic, The P-Value are in parentheses, *, ** and *** denote the significance at 10%, 5% and 1% levels.

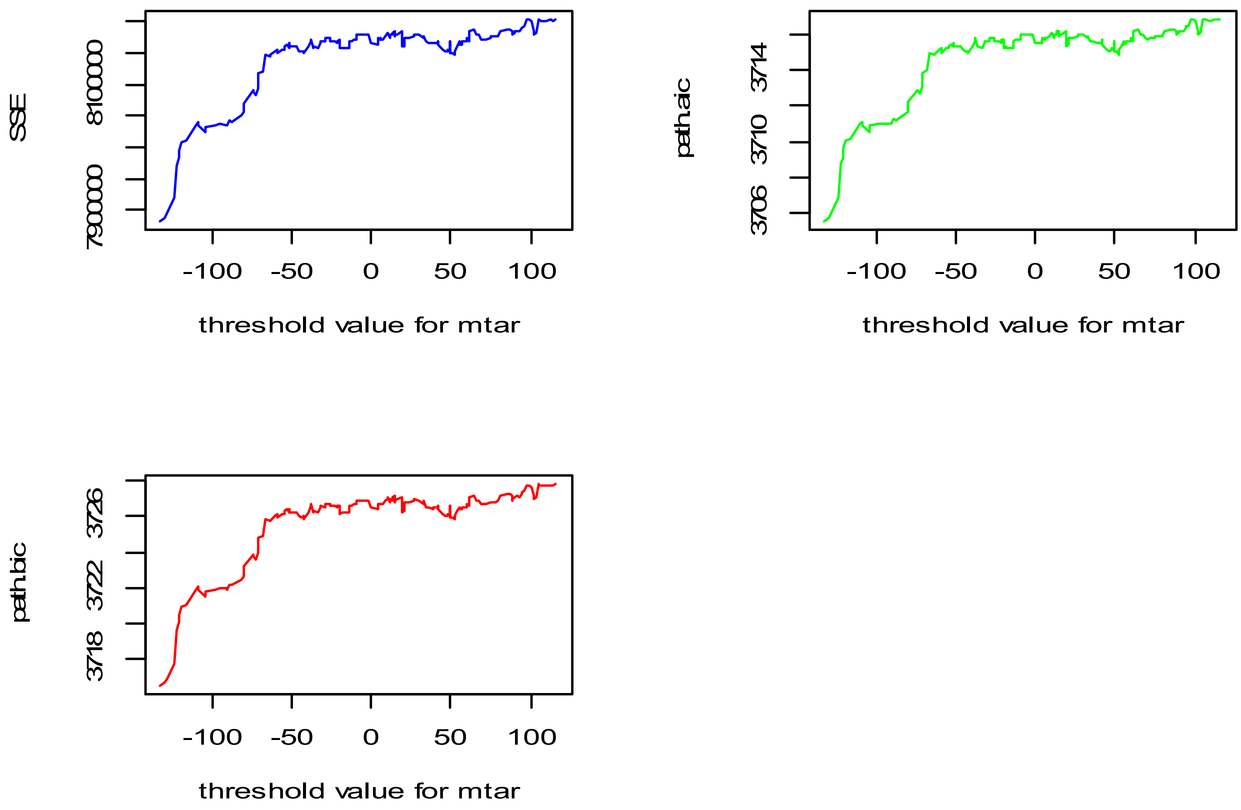

Figure2. Threshold value for M-TAR

5. Conclusion

In this paper, we study the dynamic interaction between gold price and economic policy uncertainty by considering the United States economy. Specifically, we focused on the linkages between variables in both the short-run and long-run horizons under both the linear and nonlinear threshold cointegration framework. We employ the methodology developed by Enders and Siklos (2001), focused on a nonlinear (threshold) cointegration model allowing for nonlinear adjustment to long-run equilibrium. From the linear cointegration approaches, we can reject the null hypothesis of no cointegration for the consistent-MTAR. In addition, using the consistent-MTAR specification, we found evidence of asymmetry in the adjustment process to equilibrium.

Our findings indicate the presence of asymmetric effect between gold price and economic policy uncertainty. In addition, we observe the evidence of a co-movement among the price of gold and economic uncertainty. These results offer great insights into practices of the portfolio diversification and risk management. Indeed, when the level of economic policy uncertainty is high, gold is used as a safe haven. If the inflation rate rises, people no longer trust the local currency and gold becomes of paramount importance. Sometimes, tight monetary measures are taken to control economic uncertainty or reduce its impact. Such decisions may be successful and they may fail. So, investing in gold can be the ideal solution for investors to protect their wealth and avoid the negative effects of economic uncertainty.

In the future research, we can consider other measures of economic uncertainty such as the VIX, the partisan conflict and the global EPU to study the impact of uncertainty measures on the price of gold.

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