Foreign Exchange Rate Movements of Fragile Five Economies: Do They Follow the U.S. Dollar Index?

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Abstract: In this paper, we examine the long-run relationship between Dollar Index and foreign exchange rates of 'Fragile Five' economies, respectively. We analyze foreign exchange rates of Turkey, Indonesia, Brazil, South Africa, India, and weighted average of the foreign exchange value of the US dollar against the currencies of both the broad group of major U.S. trading partners and group of the major currencies. We employ nonlinear cointegration framework and Granger causality tests on the weekly data covering January 2002 – June 2018. The empirical results that the foreign exchange rates do not have significant long-run relationship with the trade weighted US Dollar index. However, the dollar index does have significant impact on the foreign exchange rates of Fragile Five, respectively, in the short-run.
Keywords: Fragile Five, Dollar Index, Nonlinear Cointegration, Nonlinear Causality

Kırılgan Beşli Ekonomilerinin Döviz Kuru Hareketleri: Dolar Endeksini Takip Ediyorlar mı?
Öz: *Bu çalışmada, dolar endeksi ile 'Kırılgan Beşli' ülkeleri döviz kurları arasındaki uzun dönemli ilişki incelenmektedir. Türkiye, Endonezya, Brezilya, Güney Afrika ve Hindistan'ın döviz kurları ile birlikte ana ticaret ortakları ve başlıca para birimleri ağırlıklı dolar endeksleri analiz edilmiştir. Ocak 2002 – Haziran 2018 dönemini kapsayan haftalık verilere doğrusal olmayan eşbütünleşme sistemi ve Granger nedensellik testleri uygulanmıştır. Sonuçlar, döviz kurlarının, ana ticaret ortakları ağırlıklı dolar endeksiyle uzun vadeli ilişkiye sahip olmadığını göstermektedir. Bununla birlikte, dolar endeksinin, kısa vadede Kırılgan Beşli döviz kurları üzerinde anlamlı etkisinin olduğu görülmektedir.*
Anahtar Kelimeler: *Kırılgan Beşli, Dolar Endeksi, Doğrusal Olmayan Eşbütünleşme, Doğrusal Olmayan Nedensellik.*

I. Introduction

Foreign exchange market is the biggest and the most popular financial market in the world with daily transactions amounting up to 5.1 trillion a day as of April 2016 (BIS, 2016). The size of the market is both due to increasing trade triggered by globalization as well as due to the speculators and hedgers trying to generate profits or hedge their open positions. The exports market competitiveness and thus the current account balances of the countries are dependent on the foreign exchange rate. Since the level of exchange rate is

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significant also from the government perspective, central banks are correspondingly actors that try to intervene to the foreign exchange market either directly or indirectly using policy tools. Considering the amount of parties in the market and the size of the market, estimating the foreign exchange rate and to note its determinants have always been a noteworthy topic in the literature by academics.

The pace of globalization resulted in higher world trade in addition to the increased capital flows. The capital accumulated in developed economies ran to emerging economies especially after 2000s, because of the higher return potential in those countries. Among the emerging markets BRICS countries take the lead in terms of the capital flow and trade flow (Sui and Sun, 2016), given that they are the major trading partners with developed countries like US, Japan, etc. Among them China is separated from the rest since its economy paced excessively and placed itself as the second biggest country in terms of GDP as of 2017 and it has budget surpluses since the last two decades. Following the global turmoil Morgan Stanley coined a new term "fragile five" countries, namely Turkey, Brazil, India, South Africa and Indonesia, which have become too dependent on hot money to finance their growth. After the tapering of the Federal Reserve, these countries started to be judged as most risky countries. But yet later in 2017 current account deficits have declined markedly in these countries, invalidating the doubts. Still, these countries keep the doubt alive with their increasing indebtedness.

Considering these developments and the interest in fragile five countries, the aim of this paper is to evaluate the long-run relationships of the foreign exchange rates of these countries with US dollar indices. The paper adopts two versions of dollar index, that is the dollar value against its major trading partners and the other one is the value of dollar against the major currency values. The data is for the period from January 2002 to June 2018, covering the phase when the amount of the capital flows to the emerging markets were at the highest as well as the aftermath of the global financial crisis. Our study contributes to the existing literature by examining the long-run relationship between dollar indices and foreign exchange rates of Fragile Five economies, respectively. Given that the economic and financial variables exhibit generalized autoregressive conditional heteroskedasticity (GARCH) and stochastic volatility behavior (see *inter alia*, Chou, 1988; Kim et al., 1998), we use a novel econometric methodology that considers those stylized facts of the variables and that is appropriate for the nonlinear structure of the foreign exchange rates data (see *inter alia*, Granger, 1989; Granger and Teräsvirta, 1993; Taylor and Peel, 2000). To the best of our knowledge, this is the first study employing Maki's (2015a, 2015b) nonlinear cointegration framework on the foreign exchange rates data. The nonlinear cointegration framework includes wild bootstrap unit root test in exponential smooth transition autoregressive (ESTAR) models and wild bootstrap cointegration test in ESTAR error correction model. These tests have better size and power properties in the presence of unknown heteroskedastic variances, multivariate GARCH errors and stochastic volatility than (linear) conventional tests as their statistical significances are calculated by wild bootstrapping. Moreover, we apply the causality test developed by Diks and Panchenko (2006) to investigate the nonlinear Granger causality linkages among the exchange rates.

The remainder of the paper is as follows. Section 2 provides a brief literature, Section 3 presents the econometric methodology, Section 4 presents the data and the results, and the final section concludes.

II. Literature Review

The literature on foreign exchange linkages start with papers that find cointegrating relationships between exchange rates using OLS or MLE methods (see for example Hakkio and Rush, 1989; Copeland, 1991; Rapp and Sharma, 1999; Ferré and Hall, 2002;). These papers are afterwards criticized since the OLS and MLE methods are inadequate to find cointegration (Kang, 2008). As the econometric methodology advanced more papers analyze the various aspects of foreign exchange rates and more are yet to come.

Some papers investigate the linkages among exchange rate series from the viewpoint of volatility spillovers since the seminal paper by Engle, Ito and Lin (1990). They found that exchange rate uncertainty arises due to the shocks in individual markets as well as due to shocks transmitted across markets. Inagaki (2007) applied a cross-correlation function to analyze the volatility spillovers between euro and the pound and reported a unidirectional causality-in-variance from euro to pound. Kitamura (2010) analyzed the intraday interdependence and volatility spillover among the euro, the pound and the Swiss franc using varying-correlation model of multivariate GARCH. His results show that return volatility in the euro spills to the latter two.

Nikkinen, Sahlström, and Vähämaa (2006) concluded that the implied volatility of euro effects the pound and the Swiss franc by using VAR and Granger causality models. Antonakakis (2012) examine return co-movements and volatility spillovers of euro, British pound, Japanese yen and Swiss franc against the US dollar, for the period before and after the introduction of the euro. He reports co-movements and volatility spillovers between the series but notes that their magnitude is lower in the post euro period.

Another strand of the literature focus on the interdependence of exchange rates by analyzing time-varying correlations. Among them Pérez-Rodríguez (2006) used Engle (2002) Dynamic Conditional Correlation GARCH model with country specific effects to analyze the effects of conditional volatilities in returns of the euro and other major currencies against US dollar rate and reported contemporaneous and lagged volatility spillovers in the yen, dollar and euro series. Patton (2006) used copula models to test the asymmetric dependence between Deutsche Mark and yen and noticed different degree of correlations during joint appreciation against US dollar versus during joint depreciations. Wang and Yang (2009) also reports evidence of asymmetric volatility in the Australian dollar, pound and Yen against US dollar. According to their results a depreciation against the US dollar leads to significant greater volatility than an appreciation for the Australian dollar and the British pounds, whereas the opposite is not true for Japanese yen. Applying conditional copulas before and after the introduction of the euro Boero, Silvapulle and Tursunalieva (2010) analyzed whether the launch of the new currency had an impact on the dependence between exchange rates. They report varying degrees of co-movements for the euro, the pound and yen against US dollar. Applying multivariate asymmetric conditional correlation GARCH model Tamakoshi and Hamori (2014) use model and report a higher dependency between dollar, euro, pound and Swiss franc during periods of joint appreciation.

Some of the literature focus on the co-movements during the times of the crises. Baig and Goldfajn (1999) searched for contagion between financial markets for Thailand, Malaysia, Indonesia, Korea and Philippines by using cross-market correlation coefficient method. They conclude that the correlations in currency and sovereign spreads surge during the crisis period. Khalid and Kawai (2003) analyzed the impulse response functions for the Asian crisis. They reported that introducing a shock to the Thai foreign exchange market only effects Indonesian market, whereas the other currencies in the region are marginally affected. AuYong et al. (2004) analyzed the cointegration level and directions of causality of the foreign exchange rates during 1994 Mexican, 1997 Asian, 1998 Russian and 1999 Brazilian crisis. According to the results of Granger causality tests and impulse response analysis most of the pre-Mexican causality disappears and significant numbers of new causality emerge in the 1994 Mexican crisis while the 1997 Asian crisis generates significant spillover effects into the later part of the 1998 Russian and 1999 Brazilian crises. Chung (2006) also applied Engle's (2002) methodology and concluded that dollar-won co-movement decreased since 1997 currency crisis and the effect of yen increased over the years.

The literature note a number of papers applying ESTAR models in their empirical analysis. Taylor and Peel (2000) modeled the relationship between the deviations of dollar-sterling and dollar-mark exchange rates and simple monetary fundamentals. Rothman et al. (2001) stressed that ESTAR model is without doubt a better approach to capture money and output relationship. Maki (2006) applied the ESTAR model to analyze the term structure of interest rates. Yoon (2010) adopted ESTAR models to test the validity of the Purchasing Power Parity and concluded that TAR and ESTAR models should be considered to analyze the dynamics of exchange rates.

Our paper differs from the previous literature as we employ nonlinear cointegration and Granger causality framework, capturing the nonlinearities in the foreign exchange rates data. We test the long-run relationship between dollar indices and each foreign exchange rates of Fragile Five economies, respectively, over the period of January 2002 – June 2018 which covers global financial crisis, European debt crisis and the other socio-economic developments.

III. Methodology

A. Unit Root Test

Maki (2015a) proposes the following regression model for testing the unit root in ESTAR models:

$$\Delta y_t = \rho y_{t-1} F(.) + \sum_{j=1}^{p} \psi \Delta y_{t-j} + e_t, \tag{1}$$

where e_t is a zero mean error and $F(.)$ is a smooth transition function of y_{t-1}. $F(.)$ can be defined as follows:

$$F(y_{t-1};\kappa) = 1 - \exp\left\{-\kappa^2 y_{t-1}^2\right\}, \tag{2}$$

where κ is a parameter which determine the smoothness of the above function. The value of the $F(.)$ is bounded between 0 and 1 under the assumption of $\kappa > 0$. In the above ESTAR specifications, y_t is a near-unit root process when y_{t-1} is near zero 2014:477). Maki (2015a) introduces the wild bootstrap of the following unit root test statistic based on (1) and (2):

$$t = \inf_{\kappa \in [\kappa_{\min}, \kappa_{\max}]} \frac{\hat{\rho}}{\text{s.e.}(\hat{\rho})} \tag{3}$$

where $\hat{\rho}$ is the OLS estimate of ρ and s.e.($\hat{\rho}$) is the standard error of $\hat{\rho}$. The test statistic is using infimum-type statistics and we set $\kappa_{\min} = 10^{-1} V_T$ and $\kappa_{\max} = 10^{3} V_T$ where $V_T = \sqrt{\left(\sum_{t=1}^{T} y_t^2 / T\right)}$. The null and the alternative hypothesis of the test can be written as follows:

$$H_0 : \rho = 0, \quad H_1 : \rho < 0. \tag{4}$$

where the null hypothesis is basically unit root against the alternative hypothesis of a ESTAR process with (1) and (2). Maki (2015a) implements the wild bootstrap procedure in order to obtain the *p*-value of the test statistic (3). First, equation (1) with (2) is estimated for obtaining the residuals, and then the bootstrap sample is generated by

$$y_t^* = y_{t-1}^* + u_t^* \tag{5}$$

where $u_t^* = \varepsilon_t \hat{e}_t$ and $\varepsilon_t \sim \text{i.i.d.} N(0,1)$. Maki (2015a) states that $\hat{e}_t$ are the residuals which minimize the unit root *t*-statistics (3) of ρ across each possible

κ in the regression equation (1) with (2) (Maki, 2015a:479). Accordingly, the bootstrap test we implement in our paper is based on the following regression:

$$\Delta y_t^* = \rho^b y_{t-1}^* \left(1 - \exp\left\{-\kappa^{b2} y_{t-1}^{*2}\right\}\right) + \sum_{j=1}^{p} \psi_j^b \Delta y_{t-j}^* + \tilde{e}_j, \quad (6)$$

$$t^b = \inf_{\kappa^b \in \left[\kappa_{\min}^b, \kappa_{\max}^b\right]} \frac{\hat{\rho}^b}{\text{s.e.}\left(\hat{\rho}^b\right)} \quad (7)$$

For (6), $\tilde{e}_t$ is an error term. For (7), $\kappa_{\min}^b$ and $\kappa_{\max}^b$ are set to $\kappa_{\min} = 10^{-1} V_T^*$ and $\kappa_{\max} = 10^3 V_T^*$, respectively where $V_T = \sqrt{\left(\sum_{t=1}^{T} y_t^2 / T\right)}$. The bootstrap *p*-values associated with the unit root test statistics (7) are calculated as follows:

$$P_b(t) = \frac{1}{B} \sum_{n=1}^{B} I\left(t^b < t\right) \quad (8)$$

where *B* is the number of bootstrap repetitions and $I(\cdot)$ is an indicator function which takes value of 1 if $(\cdot)$ is true and 0 otherwise.

B. Cointegration Test

Maki's (2015b) cointegration test statistic is calculated based on the following error correction model (ECM) and the marginal vector autoregressive (VAR) model (Maki, 2015b:293):

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$$\begin{aligned} \Delta y_t &= \rho u_{t-1} \left(1 - \exp\left\{-\gamma u_{t-1}^2\right\}\right) + \omega' \Delta \mathbf{x}_t + \sum_{i=1}^{p} \psi_i' \Delta \mathbf{z}_{t-i} + e_t, \\ \Delta \mathbf{x}_t &= \sum_{i=1}^{p} \mathbf{\Gamma}_{xi} \Delta \mathbf{z}_{t-i} + \eta_t, \end{aligned} \quad (9)$$

Define

$$\mathbf{z}_t = \left(y_t, \mathbf{x}_t'\right)', \mathbf{x}_t = \left(x_{1t}, \ldots, \mathrm{x}_{mt}\right)' \quad (10)$$

where $\mathbf{z}_t$ is the $n \times 1$ vector of observable *I*(1) variables; y_t is a scalar and $\mathbf{x}_t$ is an $m \times 1$ vector. For (9), e_t and η_t are zero-mean errors, ω , ψ_i are $(m \times 1)$ and $(n \times 1)$ vectors, respectively, $\mathbf{\Gamma}_{xi}$ is an $m \times n$ matrix, and $u_t = y_t - \beta' \mathbf{x}_t$ with β'

is the $(m \times 1)$ cointegrating vector. The wild bootstrapped version of the following cointegration test statistic (Maki, 2015b: 293):

$$t_C = \inf_{\gamma \in [\gamma_{\min}, \gamma_{\max}]} \frac{\hat{\rho}}{\text{s.e.}(\hat{\rho})} \tag{11}$$

where $\hat{\rho}$ is the OLS estimate of ρ and s.e.($\hat{\rho}$) is the standard error of $\hat{\rho}$. The test statistic is using infimum-type statistics and we set $\gamma_{\text{b}\min} = 10^{-1} V_T^*$ and $\gamma_{\text{b}\max} = 10^3 V_T^*$ where $V_T^* = \sqrt{\left(\sum_{t=1}^{T} \hat{u}_t^{*2} / T\right)}$. The null and the alternative hypothesis of can be defined as:

$$H_0 : \rho_b = 0 \, , \; H_0 : \rho_b < 0 \tag{12}$$

Based on the equation (9), Maki (2015b) proposes the cointegration test using the wild bootstrap procedure. The null hypothesis of no cointegration can be tested using the following process (Maki, 2015b:293):

$$\Delta y_t^* = \omega' \Delta \mathbf{x}_t + \sum_{i=1}^{p} \psi_i' \Delta \mathbf{z}_{t-i} + e_t^*, \tag{13}$$

where $e_t^* = \varepsilon_t \hat{e}_t$; $E(\varepsilon_t) = 0$ and $E(\varepsilon_t^2) = 1$. Using the process (11), the bootstrap cointegration test is can be written as follows:

$$\Delta y_t^* = \rho \hat{u}_{t-1}^* \left(1 - \exp\left\{-\gamma_b u_{t-1}^{*2}\right\}\right) + \omega_b' \Delta \mathbf{x}_t + \sum_{i=1}^{p} \psi_{bi}' \Delta \mathbf{z}_{t-i} + v_t, \tag{14}$$

$$t_{bC} = \inf_{\gamma_b \in [\gamma_{b\min}, \gamma_{b\max}]} \frac{\hat{\rho}_b}{\text{s.e.}(\hat{\rho}_b)} \tag{15}$$

For (14), v_t is an error term; $\hat{u}_t^*$ is the error correction term based on the bootstrap sample and is given by $\hat{u}_t^* = y_t^* - \hat{\beta}_b' \mathbf{x}_t$, where $\hat{\beta}_b'$ is the estimate of the cointegration vector in the bootstrap sample. For (15), $\hat{\rho}_b$ is the OLS estimate of ρ_b and s.e.($\hat{\rho}_b$) is the standard error of $\hat{\rho}_b$. The test statistic is using infimum-type statistics and we set $\gamma_{\text{b}\min} = 10^{-1} V_T^*$ and $\gamma_{\text{b}\max} = 10^3 V_T^*$ where

$V_T^* = \sqrt{\left(\sum_{t=1}^{T} \hat{u}_t^{*2} / T\right)}$. The bootstrap *p*-value associated with the cointegration *t*-test statistic is calculated as follows:

$$P_b(t_C) = \frac{1}{B}\sum_{j=1}^{B} I\left(t_{bC} < t_C\right) \tag{16}$$

where *B* is the number of bootstrap repetitions and $I(\cdot)$ is an indicator function which takes value of 1 if $(\cdot)$ is true and 0 otherwise.

C. Granger Causality

Toda and Yamamoto (1995) suggest estimating the following VAR(*p*+*d*) model where d is the maximum integration degree of the variables:

$$y_t = c + B_1 y_{t-1} + \cdots + B_p y_{t-p} + \cdots + B_{p+d} y_{t-(p+d)} + \varepsilon_t. \tag{17}$$

where y_t is a vector of *k* variables, *c* is a vector of intercepts, ε_t is a vector of error terms, and *B* is the matrix of parameters. By imposing zero restriction on the first *p* parameters in (6), we obtain Wald statistics following χ^2 distribution, with *p* degrees of freedom, under the null hypothesis of Granger (1969) non-causality against the alternative hypothesis of Granger causality.

We test the Granger non-causality from one strictly stationary time-series (X_t) to another (Y_t). In a nonparametric setting with finite lags (i.e. l_X and l_Y), the null hypothesis of Granger non-causality test can be stated as Y_{t+1} is conditionally independent of X_t, X_{t-1}, …, X_{t-l_X} , given Y_t, Y_{t-1}, …, Y_{t-l_X} , which can be formulated as (Diks and Panchenko, 2006:1649):

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$$H_0 : Y_{t+1} \mid \left(X_t^{l_X} ; Y_t^{l_Y} \right) \sim Y_{t+1} \mid Y_t^{l_Y}, \tag{18}$$

where $X_t^{l_X} = \left(X_{t-l_X+1},\ldots,X_t\right)$ and $Y_t^{l_Y} = \left(Y_{t-l_Y+1},\ldots,Y_t\right)$, and l_Y and l_X respectively denote the lag lengths of *X* and *Y*. When we assume $l_Y = l_X = 1$ and $Z_t = Y_{t+1}$, and drop the time index in (18), we can specify a continuous random variable as $W = (X, Y, Z)$ indicating a three-variate random variable, distributed as $W_t = \left(X_t, Y_t, Y_{t+1}\right)$. Under the null hypothesis (18), the conditional distribution of *Z* given $(X, Y) = (x, y)$ is the same as that of *Z* given $Y = y$ only, and the joint probability density function $f_{X,Y,Z}(x, y, z)$ and its marginals must satisfy (Diks and Panchenko, 2006:1650):

$$\frac{f_{X,Y,Z}(x,y,z)}{f_{X,Y}(x,y)}=\frac{f_{Y,Z}(y,z)}{f_Y(y)} \tag{19}$$

for each vector (x,y,z) in the support of (X,Y,Z). Then, Diks and Panchenko (2006) state that the null hypothesis of nonlinear no causality implies:

$$q\equiv E\left[f_{X,Y,Z}(X,Y,Z)\,f_Y(Y)-f_{X,Y}(X,Y)\,f_{Y,Z}(Y,Z)\right]=0 \tag{20}$$

where $\hat{f}_W(W_i)$ is a local density estimator of a d_W-variate random vector W at W_i, defined by $\hat{f}w(W_i)=(2\varepsilon_n)^{-d_W}(n-1)^{-1}\sum_{jj\neq i}\mathrm{I}_{ij}^{w}$ that $\mathrm{I}_{ij}^{w}=I\left(\left\|W_i-W_j\right\|<\varepsilon_n\right)$ with the indicator function $I(\cdot)$, and the bandwidth ε_n, which depends on the sample size of n. Given the estimator q specified in (20), test statistic can be written as:

$$T_n(\varepsilon_n)=\frac{(n-1)}{n(n-2)}\sum_i\left(\hat{f}_{X,Y,Z}(X_i,Y_i,Z_i)\,\hat{f}_Y(Y_i)-\hat{f}_{X,Y}(X_i,Y_i)\,\hat{f}_{Y,Z}(Y_i,Z_i)\right) \tag{21}$$

If the bandwidth depends on the sample size as $\varepsilon_n=Cn^{-\beta}$ where $C>0$, and $\beta\in(1/4,1/3)$, then the test statistic in (21) satisfies:

$$\sqrt{n}\frac{T_n(\varepsilon_n)-q}{S_n}\xrightarrow{D}N(0,1) \tag{22}$$

where S_n is an estimator of the asymptotic variance of $T_n(\cdot)$, and $\xrightarrow{D}$ represents convergence in distribution. Diks and Panchenko (2006) also put forward that this nonlinear Granger non-causality test statistic is asymptotically distributed as standard normal.

D. Data and Empirical Results

Data

We analyze foreign exchange (FX) rates against the US Dollar of Fragile-Five economies (F-5), namely Turkey (TRY), Indonesia (IDR), Brazil (BRL), South Africa (ZAR), India (INR), and weighted average of the FX value of the U.S. dollar against the currencies of both the broad group of major U.S. trading partners (TWEXB) and group of the major currencies (TWEXM). We obtain weekly data from the FactSet. The data cover the period between January 2002 and June 2018. Index base value for all series are set to 100 and expressed in natural logarithms. We plot the movements of the times series in Figure 1 and report the estimated correlation coefficients in Table 1.

Table 1: *Correlation Coefficients*

	TWEXB	TWEXM	TRY	IDR	BRL	ZAR	INR
TWEXB	1.000						
TWEXM	0.970	1.000					
TRY	0.458	0.317	1.000				
IDR	0.443	0.327	0.885	1.000			
BRL	0.886	0.810	0.616	0.582	1.000		
ZAR	0.529	0.469	0.856	0.874	0.601	1.000	
INR	0.409	0.310	0.928	0.883	0.588	0.897	1.000

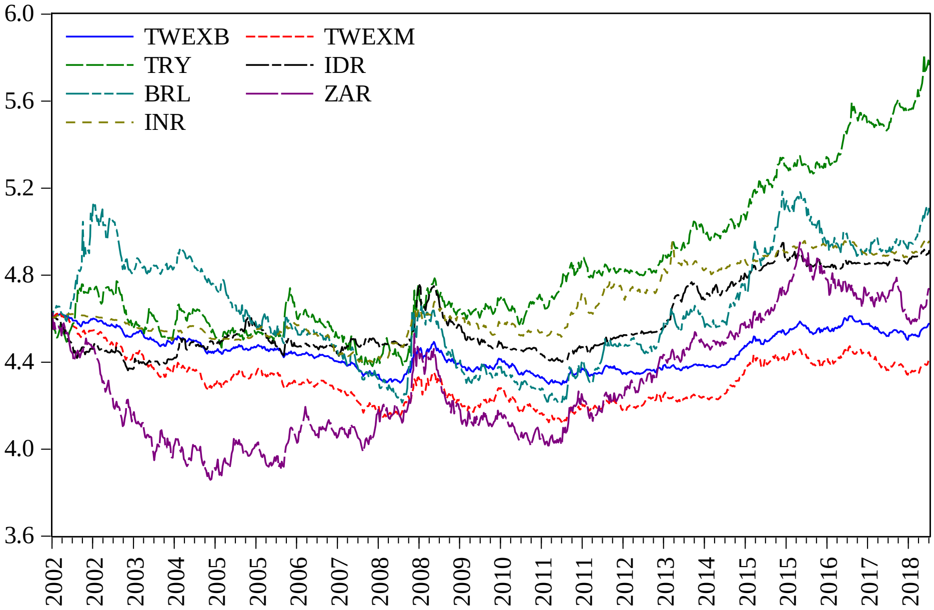

Figure 1: Foreign Exchange Rates of F-5 and Dollar Indices (Jan 2002 - June 2018)

Empirical Results

We present the results of Maki's (2015a) wild bootstrap tests for unit root in ESTAR models for the currencies and dollar indices in Table 2. Using the level of the time series, the estimated bootstrap *p*-values in Panel A of Table 2 are calculated higher than 10% significance level for all series except TWEXM, suggesting that all time-series except TWEXM have unit root. When we take the first differences of the series that have unit root, they become stationary since we can reject the null hypothesis of unit root at 1% significance level based on the estimated bootstrap *p*-values in Panel B of Table 2. These results indicate that the FX rates of F-5 and TWEXB are integrated of order one, *I*(1), while TWEXM is found to be stationary, *I*(0), at 10% significance level. We drop TWEXM from further analyses since it has a different order of integration from the others.

We employ Maki's (2015b) test to check whether the $I(1)$ time series are cointegrated. We estimate (9) with x_t as TWEXB, and each y_t being one of the FX rates of F-5 that is found to be $I(1)$ and the results are reported in Table 3. We evidence that we cannot reject the null hypothesis of no cointegration for all cases, indicating that there is no long-run relationship between TWEXB and, FX rates of F-5, respectively.

Table 2: *Maki (2015a) Wild Bootstrap Tests for Unit Root in ESTAR Models*

Series	t	$P_b(t)$	**AIC**	**lag**
Panel A. Level				
TWEXB	-2.286	0.292	-9.900	6
TWEXM	-3.069***	0.077	-9.183	12
TRY	1.028	0.996	-7.919	9
IDR	-0.530	0.960	-9.007	12
BRL	-2.062	0.608	-7.495	9
ZAR	-1.281	0.789	-7.525	10
INR	-0.556	0.891	-9.329	1
Panel B. First Difference				
TWEXB	-27.910*	0.000	-9.881	0
TWEXM	-28.480*	0.000	-9.167	0
TRY	-29.810*	0.000	-7.917	3
IDR	-26.570*	0.000	-8.993	12
BRL	-32.600*	0.000	-7.488	0
ZAR	-29.960*	0.000	-7.515	2
INR	-26.710*	0.000	-9.336	0

Note: t and $P_b(t)$ stand for unit root test statistic and the estimated bootstrap p-value, respectively. AIC is Akaike Information Criterion. *, ** and *** denote significance at 10%, 5% and 1% levels respectively. 0.000 indicates less than 0.0005.

Table 3: *Maki (2015b) Wild Bootstrap Testing for Cointegration in an ESTAR Error Correction Model Results*

	t_C	$P_b(t_C)$	**AIC**	**lag**
TRY	-0.354	0.953	-8.253	3
IDR	-1.393	0.833	-9.160	7
BRL	-3.022	0.439	-7.816	2
ZAR	-1.484	0.829	-8.008	4
INR	-1.968	0.488	-9.658	2

Note: t_C and $P_b(t_C)$ stand for cointegration test statistic and the estimated bootstrap p-value, respectively. AIC is Akaike Information Criterion.

Table 4: *Diks and Panchenko (2006) Nonlinear Granger Causality Test Results*

Lag	**Raw Returns**	**VAR Residuals**	**Raw Returns**	**VAR Residuals**
	***T*(*p*)**	***T*(*p*)**	***T*(*p*)**	***T*(*p*)**
	TWEXB $\nRightarrow$ TRY		**TRY $\nRightarrow$ TWEXB**	
1	0.248 (0.402)	-0.201 (0.580)	0.929 (0.176)	0.888 (0.187)
2	0.708 (0.239)	0.459 (0.323)	1.523 (0.064)	1.521 (0.064)
3	1.545 (0.061)	1.129 (0.129)	0.492 (0.311)	0.251 (0.401)
4	1.486 (0.069)	1.252 (0.105)	0.413 (0.340)	0.280 (0.390)
	TWEXB $\nRightarrow$ IDR		**IDR $\nRightarrow$ TWEXB**	
1	2.006 (0.022)	1.103 (0.135)	-0.983 (0.837)	-0.475 (0.683)
2	1.825 (0.034)	1.120 (0.131)	-1.010 (0.844)	0.073 (0.471)
3	1.706 (0.044)	1.474 (0.070)	-1.555 (0.940)	-0.166 (0.566)
4	1.584 (0.057)	1.676 (0.047)	-1.075 (0.859)	0.134 (0.447)
	TWEXB $\nRightarrow$ BRL		**BRL $\nRightarrow$ TWEXB**	
1	2.326 (0.010)	2.720 (0.003)	1.386 (0.083)	1.349 (0.089)
2	2.461 (0.007)	2.617 (0.004)	0.435 (0.332)	-0.722 (0.765)
3	2.386 (0.009)	2.422 (0.008)	0.274 (0.392)	0.208 (0.418)
4	2.510 (0.006)	2.857 (0.002)	0.589 (0.278)	0.508 (0.306)
	TWEXB $\nRightarrow$ ZAR		**ZAR $\nRightarrow$ TWEXB**	
1	2.319 (0.010)	2.065 (0.019)	-0.002 (0.501)	-0.026 (0.510)
2	1.350 (0.089)	1.037 (0.150)	1.021 (0.154)	1.091 (0.138)
3	0.703 (0.241)	0.702 (0.241)	1.053 (0.146)	0.955 (0.170)
4	0.469 (0.320)	1.203 (0.114)	0.984 (0.163)	0.675 (0.250)
	TWEXB $\nRightarrow$ INR		**INR $\nRightarrow$ TWEXB**	
1	0.646 (0.259)	0.290 (0.386)	0.222 (0.412)	0.465 (0.321)
2	0.995 (0.160)	0.850 (0.198)	0.998 (0.159)	0.929 (0.177)
3	0.574 (0.283)	0.716 (0.237)	1.006 (0.157)	0.871 (0.192)
4	1.108 (0.134)	1.041 (0.149)	0.765 (0.222)	0.588 (0.278)

Note: T is the nonlinear Granger causality test statistic. Numbers in parentheses are p-values. $\nRightarrow$ denotes causality direction. VAR is Vector Autoregression and l stands for lag length. Raw data indicate the series are in first differences. VAR residuals are the residuals of the VAR(p+d) models where p is the optimal lag length determined by Akaike Information Criterion, and d is the maximum order of integration of the series, which is equal to 1 in our cases.

Table 4 shows the nonlinear Granger causality test results up to 4 lags, using the first differenced data (raw returns) and residuals obtained from VAR(p+d) system. For the causality between TWEXB and Turkish Lira (TRY), the results for raw returns indicate significant (10% level) nonlinear causality from TWEXB to TRY at both third and fourth lags. However, this causality linkage is not strictly nonlinear as we do not evidence nonlinear causality running from TWEXB to TRY using VAR residuals. The results for VAR residuals also suggest nonlinear causality at 10% level from TRY to TWEXB at only second lag, indicating limited bi-directional causality among TWEXB and TRY in the short-run.

With respect to nonlinear causality among TWEXB and Indonesia Rupiah (IDR), the results for both the raw returns and VAR residuals indicate unidirectional strict nonlinear causality running from TWEXB to IDR.

For the nonlinear causality between TWEXB and Brazilian Real (BRL), the results from both raw returns and VAR residuals suggest strictly nonlinear causality from TWEXB to BRL at 1% significance level. Moreover, the nonlinear causality from BRL to TWEXB exists only at first lag and does not persist over the long-term.

For the price transmission among TWEXB and South Africa Rand (ZAR), we evidence unidirectional nonlinear causality from TWEXB to ZAR at only first lag using both raw returns and residuals obtained from VAR(p+d) system. This implies that strict nonlinear causality from TWEXB to ZAR does not persist over the long-term.

Finally, the results for the nonlinear causality among TWEXB and Indian Rupiah (INR) suggest no nonlinear causality linkages between them.

IV. Conclusion

The economies of the Fragile Five include Turkey, Brazil, India, South Africa and Indonesia. The members of this group have become too dependent on hot money to finance their growth projects as the capital accumulated in developed economies ran to emerging economies especially after 2000s. As capital flows out of emerging economies after the tapering of the Federal Reserve, the currencies of these countries have experienced considerable weaknesses, leading to difficulties to finance their current account deficits, and growth projects. Accordingly, these economies are in the group of Fragile Five, exposed to higher interest rates.

The purpose of this paper is to examine the long-run relationship between dollar indices and foreign exchange rates of Fragile Five, respectively. In this respect, we analyze foreign exchange rates of Fragile Five economies and the two versions of dollar index, the dollar value against its major trading partners, and the value of dollar against the major currency values, over the period January 2002 – June 2018. Different from the previous studies, we employ nonlinear cointegration framework of Maki (2015a, 2015b) on the foreign exchange data to better capture the nonlinearities stemming from structural breaks which eventually cause heteroskedastic variance, multivariate GARCH errors and stochastic volatility.

There are several important findings. First, dollar index measuring the value of dollar against the major currency values (TWEXM) is found to be stationary in the level, indicating that this dollar index does not have unit root and tend to revert its mean over time. Put another way, the dollar value of major currencies is somehow predictable and stable over the sample period.

Second, we do not find significant evidence of cointegration between the Trade Weighted U.S. Dollar Index (TWEXB) and the foreign exchange rates of Fragile Five economies, respectively. We also evidence statistically significant unidirectional nonlinear Granger causality from TWEXB to the foreign exchange rates of Fragile Five economies except Indian Rupiah.

These results imply that TWEXB does not have relationship with each of the foreign exchange rates of Fragile Five in the long-run, however TWEXB does have significant impact on the foreign exchange rates of Fragile Five, respectively, in the short-run. Based on the empirical evidence that the foreign exchange rate measures do not follow the dollar index in the long-run, it is worth investigating additional systematic risk factors driving the foreign exchange rates of these vulnerable Fragile Five economies; we leave this task for a future study.

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