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Dynamics of the Relationship between Inflation and Interest Rates: Testing For the Fisher Hypothesis with Structural Break(S) and Parameter Stability

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ABSTRACT: In the present study, the Fisher hypothesis that considers a one-to-one and unidirectional relationship between inflation and interest rates was tested using a quarterly frequency dataset for E-7 countries. Due to the fact that parameter constancy was not established in the study, Arai & Kurozumu and Kejriwal test techniques were used and it was found that Fisher effect was valid in weak and strong forms for country groups excluding Russia. At the same time, due to the fact that coefficient indicator modulus observed in slope parameters after structural changes had negative values, contractionary monetary policies in these economies where inflation targeting strategies were implemented resulted in weakening of the relationship between inflation and interest rate variables.

1. INTRODUCTION

The efforts to research the interaction between inflation and interest rates, which is frequently discussed and widely studied in economics literature, is significant especially in determination of the effectiveness of monetary policies mandated by central banks on the economy (Carneiro et al., 2002). In this context, Fisher hypothesis proposed by the neo-classical monetary theory reflects the situation where a positive and direct relationship between expected inflation rate and nominal interest rates occurs without affecting real interest rates (Barsky, 1987). Thus, under conditions where Fisher hypothesis is true,[1] nominal interest rates react to permanent shocks occurring for the inflation variable in the same extend and prevent real interest rates from the impact of long term monetary shocks. The assumption that shocks are transient for the real interest rate variable reflects that the related variable follows a covariance stable process, inflation neutrality does not occur between economic agents (Fried and Howitt, 1983), and the change in nominal interest rates reflects exactly the change in inflation rates (Carmichael and Stebbing, 1983). Thus, in case Fisher hypothesis is valid, it is commonly considered that the change in nominal interest rates could be utilized as a good performance indicator for future inflation rate predictions (Mishkin, 1992) and as an active monetary policy tool (Payne and Ewing, 1997). Furthermore, it was reported that public's rate of time preference and technological limitations that define real investment returns cause the constant level of real interest rates, remaining unaffected from monetary change (Woodward, 1992).

Fisher hypothesis could simply be written as shown in Equation 1:

$$i_t = r_t + \pi_t^e \tag{1}$$

In Equation 1, i_t depicts nominal interest rates, π_t^e expected inflation rate and r_t depicts the real interest rate. Since a change in π_t^e does not have a permanent effect on r_t, a change in expected inflation is completely reflected on the nominal interest rates. This relationship is also known as long-term Fisher effect. π_t^e and series i_t degree of integration I(1) could be tested as shown in long term Fisher effect co-integration equation (Equation 2) below:

[1] Obtaining valid data for Fisher hypothesis is inconsistent with efficient market hypothesis. Because, economic agents fully utilize the price changes in the market. For detailed information, see Fama, 1975.

$$i_t = \alpha + \beta_1 \pi_t^e + \varepsilon_t \qquad (2)$$

In addition to reflecting powerful validity of Fisher hypothesis and the condition that degree of co-integration is $CI[1,1]$ and ε_t is I(=), testing $\beta_1 = 1$ parameter indicates the existence of the possible substitutability among securities and capital tools (Choi, 2000). This condition demonstrates that expected inflation rate was not effective on real interest rates and the money is super neutral (Atkins and Coe, 2002). $\beta_1 \langle 1$ demonstrates weak form of Fisher effect is valid and real interest rates correlate with expected inflation in the negative direction (Lee, 2007). Thus, inability to adapt interest rates fully to the expected inflation rate reflects the existence of a monetary illusion (Coppock and Poitras, 2000). Reduction of purchasing power and consequently increasing monetary holding costs, especially in an inflationist economy, results in an increase in capital stocks by increasing the demand for alternative investment tools (Tobin, 1965). Especially the low marginal productivity of capital under decreasing returns to scale assumption results in the decrease of interest rates and $\beta_1 \langle 1$ reflects weak form Fisher effect. However, the will of investors to reorganize their portfolios favoring real assets in case of high inflation expectations and the existence of the monetary illusion, prevents the occurrence of the full Fisher effect (Modigliani and Cohn, 1979; Summers, 1982). On the contrary, limitation of interest payments by state regulations and the existence of a high degree substitutability between money and bonds could result in a reverse Fisher form. In other words, while nominal interest rates could not be adapted for inflationary changes, real interest rates interact with inflation rate in the negative direction. Thus, in classical Fisher equation, monetary policy changes would not affect nominal interest rates and the trend would be a stable process (Choudhry, 1997). Finally, co-integration coefficient would change within the interval of the related parameter $[1.3-1.5]$ as a result of the tax effect (Crowder and Hoffman, 1996).[2]

Fisher hypothesis was continuously used in the literature to determine the effectiveness of monetary policies, their use in prediction of expected inflation, and understanding the effect of real interest rates on foreign trade and capital flow (Ito, 2009). Although findings in conducted studies demonstrated evidence supporting the validity of Fisher hypothesis, there are several studies that found the hypothesis was invalid due to factors such as the inability of frequency structure, time period, methodological differences and external shocks on inflation variable to create sufficient permanent shocks to test the hypothesis and fractional integration and long memory processes of the related variable as well.[3]

A literature review would reveal that Fisher hypothesis was tested using unit root and co-integration tests. In the basis of the use of unit root tests in testing Fisher hypothesis lies the assumption that the change based on transient shocks in real interest rates that represent the difference between nominal interest and inflation rates would be balanced by continuous shocks on nominal interest and inflation rates (Evans and Lewis, 1995; Bonham, 1991). Econometrically, this expression reflects that real interest rates have the dynamics to return to the average. Studies that reported Fisher hypothesis as valid were those by Lai(1997), Lanne (2001), Million (2003), Malliaropulos (2000), while Inder and Silvapulle (1993) found that the hypothesis was invalid. The studies that arrived to the conclusion that the hypothesis was valid using the co-integration technique were Dutt and Ghosh (1995), Granville and Mallick (2004), Gül and Açıkalın (2008), Million (2004), Atkins and Coe (2002), Westerlund (2008), Wallace and Warner (1993), Juntilla (2001), Şimşek and Kadılar (2006), Thornton (1996), Özcan and Arı (2015), while Koustas and Serletis (1999), Pelaez (1995), Hawtrey (1997), Ghazali and Ramle (2003), Christopoulos and Leon-Ledesma (2007), and Yılancı (2009) reported that Fisher hypothesis was not valid.

The present study that tests whether Fisher effect was valid for E-7 countries includes four sections. The introduction presents required theoretical background and current literature for Fisher hypothesis. Second and third sections contain econometric methodology and analysis results, respectively. In the final section, findings of the study are discussed.

2. ECONOMETRIC METHODOLOGY

The efforts to research long term relationships between economic variables started with the seminal studies of Granger (1983), Engle & Granger (1987) and quite powerful and alternative tests were introduced via development of economic techniques (methods) and software. It is necessary for the obtained predictors to follow a stable course to utilize the economic relationship as a predictor in economic policies and tests that take this structure into account could create instructive results for policy makers. The existence of factors such as policy changes and shocks in time series regression analyses results in parameter instability problem and variation in estimated parameters within the time dimension (One potential problem with time series

[2] For detailed information see: Darby (1975).
[3] For detailed information see: Jensen (2009).

regression models is that the estimated parameters may change over time.) (Hansen, 1992). Kejriwal (2008) stated that factors such as structural changes and policy variations prevent the parameter constancy condition and could result in spurious regression problem. To remove this problem, Kejriwal and Perron (2008; 2010) proposed a test procedure that allows multiple breaks in co-integrated regression models. Kejrival and Perron (2010) test procedure, in addition to being a useful method for both identification of parameter instabilities and to obtain consistent estimators along with multiple breaks (Esteve et al., 2013), is an extension of the Bai & Perron (1998) study and could be used for models where different degrees of constant and co-integrated variables coexist (Adachi & Liu, 2009).

2.1. Structural Break Test

In the stage after the existence of unit root is verified, it is necessary to test whether the relationship is consistent in the time dimension. LM tests statistics used to test this relationship could cause certain problems. It was argued that structural changes in the marginal distribution of the independent variables reduced testing potency of LM-type techniques (Hansen, 2000). Since LM-type applications used in unit root tests are non-monotonic in finite samples, type-1 error possibility increases. This in turn is a problem related to the estimation of long term variance for error terms under the parameters are stable null hypothesis. Furthermore, in addition to causing an increase bandwidth, increase in the break magnitude reduces the testing power of LM-type techniques by causing an increase in long term variance (Kejriwal & Perron, 2008). To resolve abovementioned problems, Kejriwal & Perron (2010) developed *Sup-Wald* test that takes structural changes and parameter inconsistency into account. For this purpose, Kejriwal& Perron (2010) proposed three test statistics. The first is the *Sup – F* statistic that test an alternative hypothesis that expresses k breaks (*m = k*) against the null hypothesis that assumes no breaks (*m = 0*):

$$Sup-F_T^*(k)=\sup_{\lambda\in\wedge\varepsilon}\frac{SSR_0-SSR_k}{\hat{\sigma}^2}$$

SSR_0 and SSR_k statements in Sup – F statistic represent sum of squared residuals required to test null and alternative hypotheses, respectively; $\hat{\sigma}^2$, long term variance; ($\hat{\sigma}^2=T^{-1}\sum_{t=1}^{T}\tilde{u}_t^2+2T^{-1}\sum_{j=1}^{T-1}\omega(j/\hat{h})\sum_{t=j+1}^{T}\tilde{u}_t\tilde{u}_{t-j}$) $\omega(.)$, kernel function; $\hat{h}$ is the bandwidth parameter; $\lambda_i=T_i/T$ is the scalar vector that belong to the break fraction ($i=1,...,m,T_i$) for each I; $\lambda=\left\{\lambda_1,...,\lambda_m\right\}$ and reflects the break timing. There are two approaches to determine the break fraction. The first is the pseudo-Gaussian maximum likelihood estimator (MLE) proposed by Bai et al. (1998), and the other is the least squares estimator (OLS) proposed by Kurozumi & Arai (2005). As mentioned by Kurozumi & Arai (2005), Bai et al. (1998) paper does not contain realistic assumptions, the break fraction was estimated with OLS in the present study. Arai & Kurozumi (2007) argued that the suitability of asymptotically efficient estimation technique for limiting distribution in determination of break fraction with OLS is significant. Although this technique where efficient parameters are obtained asymptotically is an extension of the Carrion-i-Silvestre & Sanso (2005) paper, Saikkonen (1991) recognizes the asymptotically efficient estimation technique. As discussed in Carrion-i-Silvestre & Sanso (2005) paper, it was stated that asymptotically efficient estimation technique gave superior results in finite samples when compared to FM (fully modified) technique.

The second technique is the *UDmax* statistic that tests the alternative hypothesis that there are *M* ($1\leq m\leq M$) unknown breaks the most against the null hypothesis that there is no structural break ($m=0$):

$$UD\max F_T^*(M)=\max_{1\leq k\leq m}F_T^*(k)$$

The third test is the sequential procedure (*SEQ*) statistic that tests the alternative hypothesis that argued *k* + 1 breaks against the null hypothesis which states that there are *k* breaks.

$$SEQ_T(k+1|k)=\max_{1\leq j\leq k+1}\sup_{\tau\in\wedge_{j,\varepsilon}}T\left\{(SSR_T(\hat{T}_1,...,\hat{T}_k)-(SSR_T(\hat{T}_1,...,\hat{T}_{j-1},\tau,\hat{T}_j,...,\hat{T}_k))/\hat{\sigma}_{k+1}^2\right\}$$

In SEQ procedure, $\wedge_{j,\in}=\left\{\tau:\hat{T}_{j-1}+(\hat{T}_j-\hat{T}_{j-1})\in\leq\tau\leq\hat{T}_j-(\hat{T}_j-\hat{T}_{j-1})\in\right\}$ and $\hat{\sigma}_{k+1}^2$ statements show the consistent estimator value for the long term variance. It was stated that consideration of the *SEQ* statistic in determination of the break count based on the significance of *UDmax* test statistic is a good strategy (Kejriwal, 2008). Thus, sequential procedure, SEQ would be utilized as the information criterion in determination of multiple breaks in the model for Fisher hypothesis that takes the interest rate and inflation relationship into consideration in the present study.

2.2. Co-integration Test with Structural Changes (Multiple Breaks)

The hypotheses that co-integration with structural breaks specify "no co-integration" or "co-integration without any structural break" as the null were rejected in the paper by Gregory & Hansen (1996), Quintos (1997) and Seo (1998) Hence rejection of these null hypotheses is often understood as the existence of co-integration with structural breaks. Thus, it is necessary to apply more technical tests that consider the null hypothesis that propose so-integration with structural breaks. In the present study, Arai & Kurozumi (2007) co-integration technique where the break point is obtained by minimizing the sum of squared residuals and the null of cointegration with an unknown single break against the alternative of no cointegration was tested was utilized. Kejriwal (2008) study, which is the expanded version of the Arai & Kurozumi (2007) study, that tests whether the series were co-integrated under multiple breaks was further taken into consideration.

Let us assume that the linear model that considers *m* structural change (that *m* + 1 regime is realized) in Kejriwal & Perron (2008; 2010) study is as follows:

$$y_t = c_j + z'_{ft}\delta_f + z'_{bt}\delta_{bj} + x'_{ft}\beta_f + x'_{bt}\beta_{bj} + u_t,\ t = (T_{j-1}+1,...,T_j), \quad \textbf{(1)}$$

To demonstrate the sample size T in Model 1, for each j, $j = 1, 2, ..., m+1$ $T_0 = 0, T_{m+1} = T$. In Model 1, y_t is the first degree integrated (order of integration) I(1) scalar dependent variable, and $x_{ft}(p_f x1)$ and $x_{bt}(p_b x1)$ I(0) are the vectors for $z_{ft}(q_f x1)$ and $z_{bt}(q_b x1)$ I(1) independent variables. *b* and *f* subscripts in Model 1 express that parameters break across the regimes and remained fixed, respectively. In this form, although Model 1 expresses a structure that allows partial structural change, the change in parameters are quite limited. Thus, under the assumption that the intercept and the slope change ($p_f = p_b = q_f$), Model 1 is transformed into a pure structural change model and allows that change of I(1) independent variable parameters across regimes. This case is given in Equation 2:

$$y_t = c_j + z'_{bt}\delta_{bj} + u_t, (t = T_{j-1}+1,...,T_j) \quad \textbf{(2)}$$

The correlation between independent variables and error terms in Model 2 causes the endogeneity problem and thus bias in parameters estimated with OLS. To resolve this problem, Saikkonen (1991) and Stock & Watson (1993) proposed to use I(1) independent variables that contain endogeneity problem in Kejriwal & Perron's (2008; 2010) co-integration model by taking their first degree differences by including them as explanatory variable in the same model and to estimate the model with dynamic-OLS. The model to be estimated with dynamic-OLS model is given in Equation 3:

$$y_t = c_j + z'_{bt}\delta_{bj} + \sum_{j=-l_T}^{l_T} \Delta z'_{bt-j}\pi_{bj} + u_t^*,\ (if\ \ T_{i-1}\langle t \le T_i) \quad \textbf{(3)}$$

According to Model 3, LM test statistic under *m* breaks is calculated as shown in Equation 4:

$$\tilde{V}_m(\hat{\lambda}) = \left(\frac{\sum_{t=1}^{T} S_t(\hat{\lambda})^2}{T^2}\right) / \Omega_{11} \quad (4)$$

Ω_{11} expression in Equation 4 depicts the consistent estimator that belongs to the long term variance of u_t^* in Equation 3; and $\hat{\lambda}$ depicts [$\hat{\lambda} = (\hat{T}_1/T,...,\hat{T}_m/T),(\hat{T}_1,...,\hat{T}_m)$] break fraction. $\hat{\lambda}$ parameter is obtained via the dynamic algorithm technique proposed by Bai & Perron (2003) minimizing the global value of error sum of squares. While testing the null hypothesis that states variables in Equation 4 are cointegrated with single break, LM ($\tilde{V}_1(\hat{\lambda})$) test statistic could cause the rejection of the hypothesis that it is cointegrated with multiple breaks. Thus, Kejriwal (2008) developed Arai & Kurozumi (2007) study and reevaluated the null hypothesis that the variables were cointegrated with multiple break using the LM test statistic. In this method, LM test statistic is compared to Arai & Kurozumi (2007) table critical values for single break ($\tilde{V}_1(\hat{\lambda})$), while for multiple breaks ($\tilde{V}_m(\hat{\lambda})$) it is compared to simulation critical values obtained by Kejriwal (2008). Furthermore, Z_t -type test statistics proposed by Gregory & Hansen (1996) and express that alternative hypothesis is cointegrated with single break will be used in the present study.

3. AN APPLICATION OF THE FISHER HYPOTHESIS: THE CASE OF E-7 COUNTRIES

In the present study where Fisher hypothesis was tested for E-7 countries, quarterly datasets obtained from IMF-IFS database contained different periods for each country.[4] For inflation and nominal interest rates, of which the level values were considered, primarily time series properties for these variables should be established with unit root tests over the full sample.

M-type modified Dickey-Fuller and Philips-Perron unit root tests that resolves the abovementioned problems and proposed by Ng & Perron (2001) and Stock (1999) were utilized in the study due to the facts that the reduction of the testing power by the approach of the root of the autoregressive polynomial to but less than the unit (Dejong et al., 1992), negative autocorrelation of moving average polynomial roots of first-differentiated series due to size distortions (Schwert, 1989; Perron & Ng, 1996), and occurrence of different stationary properties in different latency lengths (Ng and Perron, 1995). Test statistics are given below:

$$MZ_{\alpha}^{GLS} = (T^{-1}y_T^2 - s_{AR}^2)(2T^{-2}\sum_{t=1}^{T} y_{t-1}^2)^{-1}, MSB^{GLS} = (T^{-2}\sum_{t=1}^{T} y_{t-1}^2 / s_{AR}^2)^{1/2}$$

$$MP_T^{GLS} = \left[\bar{c}^2 T^{-2}\sum_{t=1}^{T}\tilde{y}_{t-1}^2 - \bar{c}T^{-1}\tilde{y}_T^2\right] / s_{AR}^2, \text{ for } p=0$$

$$MP_T^{GLS} = \left[\bar{c}^2 T^{-2}\sum_{t=1}^{T}\tilde{y}_{t-1}^2 (1-\bar{c}) T^{-1}\tilde{y}_T^2\right] / s_{AR}^2, \text{ for } p=1$$

$$MZ_t^{GLS} = MZ_{\alpha}^{GLS} x MSB^{GLS}, s_{AR}^2 = \hat{\sigma}_k^2 / (1-\hat{\beta}(1))^2, \hat{\beta}(1) = \sum_{i=1}^{k}\beta_i$$

Table 1 reports the Ng & Perron (2001) unit root test results. Based on these results, interest rate variable has a unit root and null hypothesis of non-stationarity for interest rate variables could not be rejected for China, India, Mexico, Russia and Turley, while it could be rejected for Brazil and Indonesia. Similarly inflation variable has a unit root and the null hypothesis could be rejected for Turkey at 1% significance level. The results demonstrated that nominal interest rate variable order of integration for China, India, Mexico, Russia and Turkey was I(1). It is I(0) for inflation variable only in Turkey.

Table 1: Ng & Perron (2001) Unit Root Test Results

	Deterministic Component - $Case:\ p=1,\ \bar{c}=-13.5\ \ I(1)\ vs\ I(0)$							
Statistic	Variable	Brazil	China	India	Indonesia	Mexico	Russia	Turkey
MZ_{α}^{GLS}	i	-59.044*	-1.39	-4.7	-16.397***	-4.781	0.485	-3.037
	π	1.906	-1.591	-3.049	-3.21	-0.965	1.389	-70.149*
MSB^{GLS}	i	0.092	0.561*	0.326*	0.175**	0.311*	4.558*	0.393*
	π	0.987*	0.402*	0.4*	0.393*	0.407*	0.761*	0.083
$MZ_t^{GLS} = ADF^{GLS}$	i	-5.433*	-0.779	-1.533	-2.863***	-1.487	2.21	-1.196
	π	1.881	-0.639	-1.221	-1.262	-0.392	1.057	-5.881*
MP_T^{GLS}	i	0.415	16.293*	5.213*	1.494	5.264*	116.071*	29.089*
	π	80.685*	11.312*	8.011*	7.628*	12.874*	46.641*	1.478
Statistic	Deterministic Component - $Case:\ p=0,\ \bar{c}=-7.0\ \ I(2)\ vs\ I(1)$							
MZ_{α}^{GLS}	Δi	-	-44.802*	-35.855*	-	-37.544*	-5.985***	-41.200*
	$\Delta\pi$	-20.239*	-29.189*	-49.731*	-32.939*	-27.288*	-21.927*	-
MSB^{GLS}	Δi	-	0.106	0.118	-	0.115	0.287	0.110
	$\Delta\pi$	0.155	0.125	0.099	0.123	0.133	0.151	-
$MZ_t^{GLS} = ADF^{GLS}$	Δi	-	-4.733*	-4.234*	-	-4.330*	-1.715***	-4.536*
	$\Delta\pi$	-3.141*	-3.654*	-4.909*	-4.055*	-3.637*	-3.309*	-
MP_T^{GLS}	Δi	-	0.548	0.683	-	0.660	0.641	0.602
	$\Delta\pi$	1.353	0.954	0.692	0.753	1.084	1.124	-
Critical Values	$Case:\ p=0,\ \bar{c}=-7.0$			$Case:\ p=1,\ \bar{c}=-13.5$				
	%10	%5	%1	%10	%5	%1		

[4] Sample period includes for Brazil, India and Indonesia: 1980Q1-2012Q4, for China: 1990Q4-2012Q4, for Mexico: 1981Q2-2012Q4, for Russia: 1994Q3-2012Q4 and for Turkey: 1986Q2-2012Q4.

MZ_{α}^{GLS}	-5.7	-8.1	-13.8	-14.2	-17.3	-23.8
MSB^{GLS}	0.275	0.233	0.174	0.185	0.168	0.143
$MZ_{t}^{GLS} = ADF^{GLS}$	-1.62	-1.98	-2.58	-2.62	-2.91	-3.42
MP_{T}^{GLS}	4.45	3.17	1.78	6.67	5.48	4.03

Notes: Test equations for unit root process in level (in first difference) includes trend and intercept (intercept). *,** and *** denote significance at the %1, %5 and %10 levels, respectively. Δ indicates first difference of any series.

Following the determination of variables' order of integration, it is necessary to test whether Fisher hypothesis that takes the relationship between inflation and interest rates in the long term into consideration was stable. For this purpose, the testing procedure proposed by Kejriwal & Perron (2008; 2010) was considered with a maximum break value of 5 and $\varepsilon = 0.15$ trimming value and obtained results are reported in Table 2.

Table 2: Stability Test of the Relation Between Inflation and Interest Rate

Testing Multiple Structural Breaks in Cointegrated Regression Models for E-7 Countries														
Countries	Brazil		China		India		Indonesia		Mexico		Russia		Turkey	
$y_t = \{i_t\}, z_t = \{1, \pi\}, \varepsilon = 0.15, x_t = \{\phi\}, q = 2, M = 5$														
Statistic	*h* 13	*p* 2	*h* 18	*p* 2	*h* 19	*p* 3	*h* 19	*p* 5	*h* 18	*p* 3	*h* 10	*p* 2	*h* 15	*p* 2
Sup F(1)*	27.34*		8.798		3.605		7.054		12.48***		39.937*		23.252*	
Sup F(2)*	14.083*		11.181**		5.317		9.149		10.463**		14.367*		12.942*	
Sup F(3)*	9.367**		9.017**		3.599		7.373		5.359		9.544**		10.463**	
Sup F(4)*	7.313***		8.228**		2.714		5.292		5.856		7.059***		8.013**	
Sup F(5)*	5.848***		6.846**		2.174		4.535		4.893		5.642***		6.492**	
Udmax	27.34*		11.181		5.317		9.149		12.48***		39.937*		23.252*	
Number of Breaks Selected by Procedure *(SEQ)* and Criteria *(BIC, LWZ)*														
(BIC)	1		4		2		2		3		2		3	
(LWZ)	1		2		2		2		2		2		1	
(SEQ)	1		0		0		0		3		2		1	

Notes: y_t, z_t, q, x_t, p, h, and M denote the dependent variable, the regressor, number of regressors whose coefficients are allowed to change: the intercept and slope, the number of I(0) variables, number of first differenced regressors used for DOLS estimation, the minimum number of observations in each segment, and maximum number of structural changes allowed, respectively. *, **, and *** denote significance at the 1%, 5%, and 10% levels, respectively. The critical values are taken from Kejriwal & Perron (2010), Table 1, trending case, qb = 1.

Although parameter instability or structural change finding in cointegration regression for India and Indonesia were not supported by any statistic, BIC and LWZ information criteria indicate the existence of break for both countries. Furthermore, *UDmax* test statistic that tests the existence of at least one break for Brazil, Mexico, Russia and Turkey was found statistically different than zero at 1% significance level (10% for Mexico) and null hypothesis of no structural change in cointegration regression in related countries was rejected. In addition, although there was a finding based on LWZ information criterion that there were two breaks in Fisher cointegration regression for China, this finding was not supported by any *Sup – F* statistic. Due to the fact that in case the model used for Fisher hypothesis exhibited spurious regression, it had a tendency to reject the null of coefficient stability, it is necessary to test the cointegration relationship between inflation and interest rates. In other words, since stability tests also reject the null of coefficient stability when the regression is a spurious one, it is necessary to confirm the presence of cointegration between the variables. In this stage, cointegration techniques that take obtained breaks based on information criteria into consideration were applied. In the case of single break, Gregory-Hansen and Arai-Kurozumi, in case of multiple breaks, Kwjriwal-extended/Arai-Kurozumi cointegration regressions were tested. Initially, the cointegration relationships in country groups where there was no break based on information criteria were scrutinized. For instance, while *SEQ* procedure selects no break in cointegration regression for China, India and Indonesia, *LWZ* information criterion selects two breaks. Table 3 demonstrates test statistics that were obtained by considering the breaks selected based on information criteria in cointegration regression.

Table 3: Testing Long-Run Relation Between Inflation and Interest Rate

I. Gregory-Hansen and Arai-Kurozumi Cointegration Tests with a Single Break

Gregory-Hansen[a]			Arai-Kurozumi Cointegration[b]		
Statistic	Brazil	Turkey	Statistic	Brazil	Turkey
Z_t^*	-15.871*	-7.904*	$V_1(\hat{\lambda})$	0.0984	0.0634
Z_α^*	-132.560*	-83.108*	Break Frac.	0.2755	0.2941
ADF_t^*	-31.088*	-9.033*	Break Date	1998:Q3	1994:Q1

II. PP and ADF Cointegration Test Without Break[c]

Statistic	Chine	India	Indenosia
Z_t	-7.061*	-1.797	-3.305**
Z_α	-66.233*	-6.151	-20.894**
ADF_t	-3.038	-1.804	-2.824

III. Arai-Kurozumi Cointegration Test with Multiple Breaks[d]

Two Breaks						Three Breaks		
Statistic	China	India	Mexico	Indonesia	Russia	Statistic	Mexico	Turkey
$V_2(\hat{\lambda})$	0.083***	0.119**	0.058	0.091***	0.053	$V_3(\hat{\lambda})$	0.053	0.047
Break Dates	1999:Q1 2000:Q4	2002:Q1 2006:Q4	1994:Q1 2001:Q2	1994:Q1 2001:Q2	1999:Q4 2003:Q1	*Break Dates*	1994:Q1 1998:Q3 2001:Q2	1993:Q3 1999:Q4 2007:Q4

Notes: Rejection of the null hypothesis at 1%, 5% and 10% significance level are/is indicated by *,** and *** respectively.

a. Testing the null hypothesis of no cointegration against the alternative of cointegration in the presence of possible regime shift, the critical values are taken from Gregory & Hansen (1996), Table 1.

b. Testing the null hypothesis of cointegration with a structural break against the alternative hypothesis of no cointegration. The critical values are taken from Arai & Kurozumi (2007), Table 1 for $\tau=0.5$ since the limiting distributions of the test statistics are symmetric around $\tau=0.5$.

c. Testing the null hypothesis of no cointegration against the alternative of cointegration with no break. Critical values are taken from Maddala & Kim (1998), Table 6.2.

d. Testing the null hypothesis of cointegration with multiple structural break (2 or 3) against the alternative hypothesis of no cointegration. The critical values were obtained by simulations using 500 steps to approximate the Wiener process (by partial sums of i.i.d. N(0, 1) random variables) and 2000 replications. Critical values are 0.1545(1%), 0.0939(5%) and 0.0753(10%) for two breaks, 0.1531(%1), 0.0917(%5) and 0.0728(%5) for three breaks.

Accordingly, based on PP & ADF test results that considers cointegration relationship without break, although the existence of Fisher effect is determined in China and India, Table 3 shows that ADF statistics were not significant. *Sup – F* test results that indicate there was a single break based on information criteria for Brazil and Turkey necessitate testing of Fisher hypothesis for related countries with G-H and Arai-Kurozumi cointegration techniques that take single break into account. Based on the statistics shown in the first section of Table 3, structural change and cointegration relationship was determined both in Brazil and Turkey and it was concluded that Fisher effect was valid in related countries. Furthermore, PP & ADF test results for China, India and Indonesia that indicates there was no structural change based on SEQ information criteria are presented in the scond section of Table 3. Based on the statistics, it was concluded that Fisher hypothesis was valid for China and Indonesia, but not for India. At the final section of Table 3, whether Fisher hypothesis was valid in related countries based on multiple break results obtained with information criteria was investigated. It was determined that Fisher effect was valid with structural change for two breaks in Mexico and Russia, and for three breaks in Mexico and Turkey.

Table 4 presents the country-based comparison of parameters obtained from the findings of models that take structural break into account.

Table4: E-7 Countries Estimated Parameters Under the Breaks

Parameter	Brazil	Chine	India	Indonesia	Mexico(2)	Mexico (3)	Russia	Turkey (1)	Turkey (3)
a0	1942.309	38.109	9.372	16.099	52.119	51.856	262.828	38.097	38.026
	[184.252]	[5.625]	[19.418]	[1.735]	[7.217]	[8.39]	[9.266]	[6.164]	[6.107]
a1	-306.664	107.168	145.259	54.991	78.423	72.229	64.332	69.277	79.174
	[84.227]	[3.054]	[26.354]	[2.35]	[8.505]	[8.073]	[9.265]	[3.979]	[6.69]
a2	-	21.632	-86.311	-3.607	17.936	63.805	0.492	-	195.598
	-	[5.625]	[37.271]	[2.82]	[4.06]	[9.888]	[4.482]	-	[8.637]
a3	-	-	-	-	-	22.797	-	-	41.923
	-	-	-	-	-	[5.817]	-	-	[5.914]
bo	-71.145	0.212	0.03	-0.272	7.512	8.649	-14.073	132.603	134.53
	[9.475]	[0.354]	[1.228]	[0.254]	[3.498]	[4.067]	[3.556]	[27.643]	[27.388]
b1	-4.991	-0.166	-0.647	-0.677	-3.413	-2.642	-1.333	-0.529	-0.422
	[3.105]	[0.237]	[0.654]	[0.119]	[2.262]	[1.597]	[0.794]	[0.069]	[0.744]
b2		0.574	16.705	-0.021	-0.216	-0.846	-0.037	-	-1.824
		[0.367]	[22.652]	[0.197]	[0.121]	[0.787]	[0.114]	-	[0.462]
b3	-	-	-	-	-	-0.198	-	-	-0.132
	-	-	-	-	-	[0.342]	-	-	[0.238]
	1998:Q3	1999:Q1	2002:Q1	1994:Q1	1994:Q1	1994:Q1	1999:Q4	1994:Q1	1993:Q3
Break Date(s)	-	2000:Q4	2006:Q4	2001:Q2	2001:Q2	1998:Q3	2003:Q1	-	1999:Q4
	-	-	-	-	-	2001:Q2	-	-	2007:Q4

Notes: Standard error for related parameters are shown in brackets.

Obtained coefficient values per Table 4 were estimated within expectations with the exception of Russia. It was found that weak Fisher effect was valid in China, India and Indonesia. The weak form of Fisher effect in China, India and Indonesia prevented full adaptation of interest rates to the expected inflation rate and created monetary illusion. On the other hand, strong Fisher effect in Brazil, Mexico and Turkey reflects that money was super-neutral. It was also noticed that break values observed in slope coefficients had negative magnitude due to change in policies in all countries. Especially the existence of anti-inflationist policies affects the validity of Fisher effect in the long-run by weakening the relationship between inflation and interest rates. This finding was consistent with the results obtained in the study by Şimşek and Kadılar (2006) for Turkish economy.

4. CONCLUDING REMARKS

In the present study that scrutinized the relationship between inflation and interest rates, validity of Fisher effect on E-7 countries was tested using quarterly datasets. Findings demonstrated that Fisher effect was valid in weak and strong forms for country groups with the exception of Russia. Based on Arai & Kurozumi and Kejriwal & Perron procedures that consider parameter constancy, negative values in coefficient sign modulus observed in slope parameters after structural changes caused a decrease in the correlation between inflation and interest rates. Validity of the obtained findings are valid as long as they are consistent with studies that take non-linear relationships between related variables into account in different frequency structures and time intervals.

REFERENCES

1) Adachi, K. & Liu Donald J. (2009). Estimating Long-Run Price Relationship with Structural Change of Unknown Timing: An Application to the Japanese Pork Market, *American Journal of Agricultural Economics*, 91(5), 1440-1447.
2) Arai, Y., & Kurozumi, E. (2007). Testing for the Null Hypothesis of Cointegration with a Structural Break, *Econometric Reviews*, 26(6), 705-739.
3) Atkins,F.J. ve Coe, P.J. (2002). An ARDL bounds test of the long-run Fisher effect in the United States and Canada, Journal of Macroeconomics,24:255-266.
4) Bai, J., & Perron P. (1998). Estimating and Testing Linear Models with Multiple Structural Changes, *Econometrica,* 66(1), 47-78.
5) Bai, J., R., Lumsdaine L., & Stock, J. H. (1998). Testing for and Dating Common Breaks in Multivariate Time Series, *Review of Economic Studies*, 65(3), 395-432.
6) Bai, P. & Perron P. (2003). Computation and Analysis of Multiple Structural Change Models, *Journal of Apllied Econometrics*, 18(1), 1-22.

7) Barsky, R.B. (1987). The Fisher Hypothesis and the Forecastability and Persistence of Inflation, *Journal of Monetary Economics*,19: 3-24.
8) Berument, H. ve Jelassi, M. M. (2002) The Fisher hypothesis: a multi-country analysis. Applied Economics, 34:1645-1655.
9) Bonham, C.S. 1991. Correct Cointegration test of the long-run relationship between nominal interest and inflation, applied Economics, 23:1487-1492.
10) Carmichael, J. Ve Stebbing P.W.1983. Fisher's Paradox and Theory of Interest,The American Economic Review, 73(4):619-630.
11) Carneiro, Francisco, G., Angelo, J. Ve Rocha,C.H. (2002). Revisiting the Fisher hypothesis for thecases of Argentina, Brazil and Mexico, Applied Economics Letters, 9: 95-98.
12) Carrion-i-Silvestre, J. L., Sanso, A. (2004). Testing the Null of Cointegration with Structural Breaks. Mimeo.
13) Choi, W.G. (2000). The Inverted Fisher Hypothesis: Inflation Forecastability and Asset Substitution, IMF Working Paper, WP/00/194.
14) Choudhry,T. (1997). Cointegration analysis of the inverted Fisher effect: evidence from Belgium, France and Germany, Applied Economics Letters, 4(4):257-260.
15) Christopoulos, D.K. ve Leon-Ledesma M.A (2007). A Long-Run Non-Linear Approach to the Fisher Effect, Journal of Money, Credit and Banking, 39(2-3):543-559.
16) Coppock, L ve Poitras, M. (2000). Evaluating the Fisher effect in long-term cross-country averages, International Review of Economics and Finance, 9:181-192
17) Crowder, W.J. ve Hoffman, Dennis L. (1996). The Long-Run Relationship between Nominal Interest Rates and Inflation: The Fisher Equation Revisited, Journal of Money, Credit and Banking, 28(1):102-118.
18) DeJong, D.N.J., Nankervis J.C., Savin N.E. & Whiteman C.H. (1992). Integration versus Trend Stationary in Time Series, *Econometrica*, 60(2), 423-433.
19) Dutt, S.D. ve Ghosh, D. (1995). The fisher Hypothesis: Examining the Canada Experience, Applied Economics, 27(11):1025-1030.
20) Engle, R. F. & Granger C. W. J. (1987), Cointegration and Error Correction: Representation, Estimation and Testing, *Econometrica*, 55(2), 251-276.
21) Esteve, V., Navarro-Ibanez, M. & Prats, M.A. (2013). The Spanish Term Structure of Interest Rates Revisited: Cointegration with Multiple Structural Breaks, 1974-2010, *International Review of Economics and Finance*, 25(C), 24-34.
22) Evans, M.D.D. ve Karen K. Lewis (1995). Do Expected Shifts in Inflation Affect Estimates of the Long-Run Fisher Relation?, he Journal of Finance, 50(1), 225-253.
23) Fama, E.F. (1975). Short-Term Interest Rates as Predictors of Inflation, The american Economic Review, 65:269-283.
24) Fried, J. Ve Howitt, P. (1983). The effects of Inflation on Real Interst Rates, American Economic Review, 73:968-979.
25) Ghazali N.A ve Ramle, S. (2003). A long memory test of the long-run Fisher effect in the G7 countries, Applied Financial Economics, 13:763-769.
26) Granger, C. W. J. (1983). Cointegrated and Error-Correcting Models, Discussion Paper, Department of Economics, University of California, San Diego.
27) Granville, B. Ve Mallick, S. (2004). Fisher hypothesis: UK evidence over a century, Applied Economics Letters, 11(2):87-90.
28) Gregory, A. W. & Hansen B. E. (1996). Residual-Based Tests for Cointegration in Models with Regime Shifts, *Journal of Econometrics*, 70(1), 99-126.
29) Gül, E. Ve Açıkalın, S. (2008). An examination of the Fisher Hypothesis: the case of Turkey, Applied Economics, 40(24): 3227-3231.
30) Hansen, Bruce E. (1992). Tests for Parameter Instability in Regressions with I(1) Processes, *Journal of Business & Economic Statistics*, 20(1), 45-59.
31) Hansen, Bruce E. (2000). Testing for Structural Change in Conditional Models, *Journal of Econometrics*, 97(1), 93-115.
32) Hawtrey, K.M. (1997). The Fisher Effect and Australian interest rates, Applied Financial Economic,7:337-346.
33) Inder, B ve Silavapulle P. (1993). Does the Fisher effect apply in Australia?, Applied Economics, 25(6):
34) Ito, T. (2009). Fisher hypothesis in Japan: Analysis of Long Term Interest Rates under Different Monetary Policy Regimes, The World Economy, 32(7): 1019-1035.

35) Jensen, M.J. (2009). The Long-Run Fisher Effect: Can It Be Tested?, Journal of Money, Credit and Banking, 41(1):221-231.
36) Juntilla, J. (2001). Testing an augmented fisher hypothesis for a small open economy: The case of Finland, Journal of Macroeconomics, 23(4):577-599.
37) Kejriwal, M. & Perron P. (2008). The Limit Distribution of the Estimates in Cointegrated Regression Models with Multiple Structural Changes, *Journal of Econometrics*, 146(1), 59-73.
38) Kejriwal, Mohitosh & Perron Pierre (2010). Testing for Multiple Structural Changes in Cointegrated Regression Models, *Journal of Business & Economic Statistics*, 28(4), 503-522.
39) Kejriwal, Mohitosh. (2008). Cointegration with Structural Breaks: An Application to the Feldstein-Horioka Puzzle, *Studies in Nonlinear Dynamics & Econometrics*, 12(1), 1-37.
40) Koustas, Z. Ve Serletis, A. (1999). On the Fisher effect, Journal of Monetary Economics, 44:105-130
41) Kurozumi, E. & Arai, Y. (2005) Efficient Estimation and Inference in Cointegrating Regressions with Structural Change. Discussion Paper No. 2004-9, Graduate School of Economics, Hitotsubashi University.
42) Lai, Kon, S. (1997). Long-term Persistence in Real Interest Rate: Some Evidence of a Fractional Unit Root, International Journal of Finance and Economics, 2(3):225-235.
43) Lanne, M. (2001). Near Unit Root and the relationship between inflation and interest rates: A Reexamination of the Fisher Effect, Empirical Economics, 26:357-366.
44) Lee, K.F. (2007). An Empirical Study of the Fisher Effect and the Dynamic Relation Between Nominal Interest Rate and Ination in Singapore, MRPA Paper, No:12383.
45) Maddala, G.S., & Kim, I.M. (1998). Unit Roots, Cointegration, and Structural Change, Cambridge University Press, United Kingdom.
46) Malliaropulos, D. (2000). A note on nonstationarity, structural breaks, and the Fisher effect, Journal of Banking & Finance, 24(5):695-707.
47) Million, N. (2003). The Fisher Effect revisited through an efficient non linear unit root testing procedure, Applied Economics Letters, 10:951-954.
48) Million, N. (2004). Central Bank's interventions and the Fisher hypothesis: a threshold cointegration investigation, Economic Modelling, 21(6):1051-1064.
49) Mishkin, F. S. (1992) Is the Fisher effect for real? A re-examination of the relationship between inflation and interest rates, Journal of Monetary Economics, 30, 195–215.
50) Modigliani, F. and R. Cohn (1979). Inflation, Rational Valuation, and the Market, Financial Analysts Journal,35(2):24-44.
51) Ng, S. & Perron P. (1995). Unit Root Tests in ARMA Models with Data Dependent Methods for the Selection of the Truncation Lag, *Journal of the American Statistical Association*, 90(429), 268-281.
52) Ng, S. & Perron, P. (2001). Lag length Selection and the Construction of Unit Root Tests with Good Size and Power, *Econometrica*, 69(6), 1529-1554.
53) Özcan, B. Arı, A. (2015). Does the Fisher hypothesis hold for the G7? Evidence from the panel cointegration test, Economic Research,28(1):271-283.
54) Payne, J.E ve Ewing, B.T, 1997. Evidence from lesser developed countries on the Fisher hypothesis: a cointegration analysis, Applied Economic Letters,4:683-687.
55) Pelaez, R.F (1995). The Fisher Effect: Reprise, Journal of Macroeconomics, 17(2):333-346.
56) Perron, P. & Ng S. (1996). Useful Modification to Unit Root Test with Dependent Errors and Their Local Asymptotic Properties, 63(3), 435-463.
57) Perron, Pierre (1990). Testing for a Unit Root in a Time Series with a Changing Mean, *Journal of Business & Economic Statistics*, 8(2), 153-162.
58) Quintos, C. E. (1997). Stability Tests in Error Correction Models, *Journal of Econometrics*, 82(2), 289-315.
59) Saikkonen, P. (1991). Asymptotically Efficient Estimation of Cointegration Regressions, *Econometric Theory*, 7(1), 1-21.
60) Schwert, G.W. (1989). Tests for Unit Roots: A Monte Carlo Investigation, Journal of *Business and Economic Statistics*, 7(2), 147-159.
61) Seo, B. (1998). Tests for Structural Change in Cointegrated Systems Regressions, *Econometric Theory*, 14(2), 222–259.
62) Stock, J. H. & Watson, M. W. (1993). A Simple Estimator of Cointegrating Vectors in Higher Order Integrated Systems, *Econometrica*, 61(4), 783-820.

63) Stock, J.H. (1999). A Class of Tests for Integration and Cointegration, in R.F. Engle and H.White (Eds.): Cointegration, Causality and Forecasting. A Festschrift in Honour of Clive W.J. Granger, Oxford University Press, 37-167.
64) Summers, L.H. (1982). The Nonadjustment of Nominal Interest Rates: A Study of the Fisher Effect, NBER WP, WP No: 836.
65) Şimşek, M ve Kadılar, C. (2006). Fisher Etkisinin Türkiye Verileri ile Testi, Doğuş Journal, 7:99-111.
66) Thornton, J. (1996). The adjustment of nominal interest rates in Mexico: a study of the Fisher effect, Applied Economics Letters, 3(4):255-257.
67) Wallace, M.S. ve Warner J.T (1993). The Fisher Effect and the Term Structure of Interest Rates: Tests of Cointegration, The Review of Economics and Statistics, 75(2):320-324.
68) Westerlund,J. (2008). Panel cointegration tests of the Fisher effect, Journal of Applied Econometrics, 23(2):193-233.
69) Woodward, T.G. (1992). Evidence of the Fisher Effect From U.K. Indexed Bonds, The Review of Economics and Statistics, 74(2): 315-320.
70) Yılancı, V. (2009). Fisher Hipotezinin Türkiye için Sınanması: Doğrusal Olmayan Eşbütünleşme Analizi, Atatürk Üniversitesi İİBF Dergisi, 23:205-213.