Fisher Relation and Financial Crisis: Evidence for East Asian Countries

Jae-Young Kim

The Fisher relation (FR) is a key theoretical relation that underlies many important results in economics and finance. The theoretical FR implies that the two variables of the nominal interest rate and the expected inflation are cointegrated. In practice, however, real data often fail to confirm the cointegration FR. In this paper we propose that failure of confirming cointegration FR is due to nonstationary deviations in a relatively small portion of the data period. We investigate such possibility based on the notion of segmented cointegration in Kim (2003). Our analysis is to detect such a situation and to identify the segmented periods of cointegration. We analyze data from five east Asian economies in a period including the Asian financial crisis. For quarterly data for those economies, we have found that short-run nonstationary deviations cause failure of the cointegration FR. The segments of non-cointegration overlap the period of east Asian financial crisis, which started in the year of 1997, with slight difference across economies. This result implies that for data from the east Asian countries the Fisher relation prevails in the data period except the abnormal period of financial crisis.

Keywords: Fisher relation, East Asian crisis, Segmented cointegration

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I. Introduction

The Fisher relation describes that the real interest rate is equal to the nominal interest rate minus the expected inflation. The real interest rate is assumed to be a constant or a stationary variable fluctuating around a constant in many theoretical analyses. This assumption implies that the two variables of the nominal interest rate and the expected inflation, if they are integrated, are cointegrated. Often, however, the two variables fail to confirm cointgration for real data. We hypothesize that failure to confirm the 'cointegration Fisher relation' by real data is caused by nonstationary deviations from the relation in a relatively small portion of the data period. In this paper, based on this hypothesis, we perform empirical analysis for FR for data of five east Asian economies in the period containing the Asian financial crisis.

Since empirical work on the Fisher relation is initiated by Fama (1975), many researchers studied the subject. The hypothesis of constancy of the real interest is discussed in Nelson and Schwert (1977), Garbade and Wachtel (1978), Mishkin (1981, 1984), and Fama and Gibbons (1982). Nelson and Schwert (1977), Fama and Gibbons (1982). Summers (1982), Huizinga and Mishkin (1986) studied the correlation between inflation and nominal interest rates (Fisher effect). Nonstationary behavior of nominal interest rates and inflation, which necessitates a different approach for analyzing the FR, is discussed by Crowder and Hoffman (1996).

With nonstationary nominal interest rates and inflation, if the real interest rate is stationary, then common wisdom would conclude in favor of the Fisher relation (Mishkin, 1992). However, empirical results on the stationarity of the real interest rate are not uniform as can be seen in Walsh (1987), Rose (1988), Atkins (1989), and Mishkin (1992). Kim and Park (2015) studied the possibility of short-run instability of the Fisher relation for data of the U.S. and Korea by applying subsampling tests of Andrews and Kim (2008). Also, Kim and Park (2018) compares several alternative models for FR and evaluate them based on a post-data model determination method.

In this paper, we consider the possibility that weak evidence supporting the cointegration FR is caused by the short-run nonstationary deviations from the relation while the relation prevails in the remaining data period. As is well known, the concept of cointegration is often interpreted as a long-run equilibrium in a system of variables, whereas short-run deviations are allowed to a certain extent. Such short-run deviations, however, may have considerable effects on the results of statistical analysis for cointegration.

With nominal rates and inflation being nonstationary, our analysis aims to determine whether these two variables form a *segmented cointegration* that was studied by Kim (2003). The notion of segmented cointegration of Kim (2003) was introduced by noting that cointegration is often not confirmed by real data for a well-understood economic relation in the presence of nonstationary fluctuations of relatively short periods. There are various reasons for it. Prominent examples of such reasons include the possibility of structural breaks in the underlying relationship (Hansen and Johansen, 1993; Stock and Watson 1994) and the threshold effects in a likely cointegration relation (Balke and Fomby 1996). Siklos and Granger (1997) found a similar notion, where they considered the case when the given variables are cointegrated during some periods and not at others. In addition, Andrews and Kim (2008) noted that a cointegration relation may break down for short period(s), whereas the relation prevails in other periods.

We analyze data from five east Asian countries, Korea, Indonesia, Malaysia, Thailand and Singapore, for the Fisher relation by applying the methods of Kim (2003). For the data we detect the presence of segmented cointegration and identify the segmented periods of cointegration. The data period is 1989:Q1–2006:Q4 for all the countries under study. For the data, we obtain the result that the null of a unit root for the real interest rate is rejected by allowing one segmentation. The estimated non-cointegration periods overlap the period of east Asian financial crisis, which started in 1997, with slight difference across countries. Our result implies that for data from the east Asian countries the Fisher relation prevails in the data period except the abnormal period of financial crisis.

Our discussion in this paper is organized as follows. Section II explains the Fisher relation and segmented cointegration. Section III reviews methods for the inference on segmented cointegration developed by Kim (2003) with regard to the Fisher relation. Section IV explains our empirical analysis and discusses the results obtained by the methods in Section III.

II. Fisher Relation and Segmented Cointegration

A. Fisher Relation

The Fisher relation describes that the nominal interest rate has a one-for-one relation with the expected rate of inflation. In other words, the Fisher relation describes that a stable level of the “real interest rate is equal to the nominal interest rate minus the expected inflation. As is well explained in Kim and Park (2015, 2018) and others, the Fisher relation, in terms of *ex-ante* variables, is written as follows:

$$i_t = \bar{r} + \pi^e_{t+1} + \varepsilon^*_t, \quad (2.1)$$

where π^e_{t+1}, i_t, $\bar{r}$, and ε^*_t are, respectively, the expected inflation from period t to period $t + 1$, the nominal interest rate at time t, the (mean) level of real interest rate, and the *ex-ante* “disturbance. The disturbance may contain a possible liquidity effect discussed by Lucas (1990) and Fuerst (1992). In terms of *ex-post* variables, the Fisher relation becomes

$$i_t = \bar{r} + \pi_{t+1} + \varepsilon_t, \quad (2.2)$$

where π_{t+1} and ε_t are *ex-post* inflation and disturbance, respectively.

Notably, for υ_t such that $\upsilon_t = \pi_e - \pi_t$, $\varepsilon_t = \varepsilon^*_t + \upsilon_t$. Thus, if the error of the inflation expectation υ_t is a stationary variable, which is the case under rational expectations, then the *ex-ante* disturbance ε^*_t and the *ex-post* disturbance ε_t have the same statistical properties. In this case, one can analyze the Fisher relation based on the *ex-post* and *ex-ante* rates. The stationarity of the real interest rate $r_t = \bar{r} + \varepsilon_t$ implies the existence of a stable Fisher relation. Thus, typical economic theory assumes that the real interest rate is a stationary variable fluctuating around a constant mean $\bar{r}$. Examples include models of dynamic optimization and intertemporal decision-making that are used widely in economics and finance.

Although the Fisher relation is apparently simple in theory, empirical analysis of the relation is more or less complicated with mixed results. One reason for the complication in empirical work is that (nominal) interest rates and inflation usually show nonstationary properties (Crowder and Hoffman 1996). In such a situation, the data support the Fisher relation if the real interest rate is a stationary variable (Mishkin

1992). However, the stationarity of the real interest rate is not well confirmed by real data (Rose 1988; Walsh 1987).

In this study, we consider the possibility that the weak evidence favoring the Fisher relation is caused by a short-run nonstationary deviation from the Fisher relation, whereas the relation prevails in the other data periods. That is, a (relatively) short period of a regime assumes nonstationary real interest rates, whereas, in the other period, the real rate is stationary. We study such a possibility based on the notion of segmented cointegration explained by Kim (2003).

B. Segmented Cointegration

We have, in this subsection, the explanation of Kim (2003) for the concept of segmented cointegration for the Fisher relation. Thus, we consider the possibility that the real interest rate r_t or the disturbance ε_t is a unit root process in a segment of the data period but is a stationary process for the rest of the period. We reserve the notation for these two different types of periods, N_T and C_T, respectively. In view of Eq. (2.2), we define such a situation as *segmented cointegration* (SC) of the i_t and π_{t+1} with the cointegrating vector (1, −1).

Testing the cointegration for Eq. (2.2) is often based on the testing for the nullity of a unit root for ε_t. Thus, writing the disturbance process ε_t as an AR(1):

$$\varepsilon_t = \rho\varepsilon_{t-1} + \upsilon_t, \tag{2.3}$$

the hypotheses are for the value of ρ: the null of $\rho = 1$ v.s. an alternative of $\rho < 1$. Such hypotheses can be tested through several different tests, such as tests of Phillips and Perron (1988), Phillips and Ouliaris (1990), and augmented Dickey-Fuller (ADF).

In the case of possible segmented cointegration for i_t and π_{t+1}, we may consider the following hypotheses:

$$\begin{aligned} &H_0 : \rho = 1 \quad \mathit{for}\ t \in \{t = 1, \ldots , T\} \\ &H_1 : \rho = \begin{cases} 1 & \mathit{for}\ t \in N_T \\ \rho_1,\ \mathit{with}\ \rho_1 < 1 & \mathit{for}\ t \in C_T \end{cases}, \end{aligned} \tag{2.4}$$

where $N_T \subset \{1, \ldots, T\}$ and $C_T \equiv \{1, \ldots, T\} \setminus N_T$. Notably, the cointegration is a special case of segmented cointegration with $N_T = \emptyset$.

III. Inference on Segmented Cointegration

We briefly explain how to detect segmented cointegration for i_t and π_t and how to locate the period of non-cointegration (NC). Our explanation in this section is mostly from Kim (2003).

A. Detection of Segmented Cointegration

To save the notation, let C_T be a *chosen* segmented sample that may or may not coincide with the true $C_T \equiv \{1, \ldots, T\} \setminus N_T$. When the disturbance in a linear regression model such as Eq. (2.2) is heterogeneous across different periods similar to Eqs. (2.3)–(2.4), econometric theory suggests estimating the model by a weighted least square:

$$\text{argmin}_\beta\ Q(\bar{r}, w) \equiv \sum_{t=1}^{T} w_t \varepsilon_t^2(\bar{r}), \tag{3.1}$$

where $\varepsilon_t(\bar{r}) = (i_t - \bar{r} - \pi_{t+1})$, and w_t is a weight given to the t period residual and $w = (w_1, \ldots, w_T)$. For (2.3)–(2.4), the weight series $\{w_t\}$ should be as follows:

$$w_t = \begin{cases} 1 & for \quad t \in C_T \\ 0 & for \quad t \in N_T \end{cases},$$

where $w_t = 1$ for $t \in C_T$ can be replaced by $w_t =$ constant.[1]

Notice that the weighting process $\{w_t\}$ depends on the selection of C_T, $w_t = w_t(C_T)$. In addition, denote by $et(C_T)$ the residual obtained by estimating Eq. (3.1) based on the weight $\{w_t(C_T)\}$ for a segmentation C_T Let $\rho(C_T)$ be the least square estimator of the first-order autoregression coefficient ρ of et in $e_t = \rho e_{t-1} + v_t$ for $e_t = e_t(C_T)$ and $t(C_T)$ be the t test of $\rho = 1$. That is, $\rho(C_T)$ is obtained from regressing $w_t(C_T)et$ on $w_{t-1}(C_T)e_{t-1}$. Moreover, let $\hat{\sigma}^2(C_T)$ be the least square estimator of the variance of $\hat{\rho}$ and $s^2(C_T)$ be the least square estimator of the variance of the second stage residual:

[1] The specification for w_t is justified by the following: The data in N_T are not informative at all for the estimation because in N_T the relation (2.2) is a spurious regression.

$$s^2(C_T) = (T_C - 1)^{-1} \Sigma w_t(C_T)(e_t - \hat{\rho} e_{t-1})^2;$$

$$\hat{\sigma}_{\hat{\rho}}^2(C_T) = s^2(C_T)\left(\sum w_t(C_T) e_{t-1}^2\right)^{-1}$$

where T_C is the number of elements of C_T. Define

$$Z_\rho(C_T) = T_C(\hat{\rho} - 1) - \frac{1}{2}(T_C^2 \hat{\sigma}_{\hat{\rho}}^2(C_T) / s^2(C_T))(\hat{\lambda}^2 - \hat{\gamma}_0)$$

$$Z_t(C_T) = (\hat{\gamma}_0 / \hat{\lambda})^{1/2} \cdot t(C_T) - \left\{\frac{1}{2}(\hat{\lambda}^2 - \hat{\gamma}_0) / \hat{\lambda}\right\} \times \{T_C \hat{\sigma}_{\hat{\rho}}(C_T) / s(C_T)\}$$

where $\hat{\lambda}^2$ and $\hat{\gamma}_0$, respectively, are estimates of $\lambda^2 = \lambda^2 = \lim_{T_c \to \infty} T_c^{-1} E\left[(\Sigma_{t \in C_T} v_t)^2\right]$ and $\hat{\gamma}_0 = \lim_{T_c \to \infty} T_c^{-1} \Sigma_{t \in C_T} E(v_t^2)$. Consistent nonparametric estimators of λ and γ0 under H1 can be obtained as in Newey and West (1987).

The above approach is based on an AR(1) for the process *et*. Alternatively, we can consider the following variant of an ADF expression for *et*:

$$\Delta e_t = \zeta_1 \Delta e_{t-1} + \zeta_2 \Delta e_{t-2} + \cdots + \zeta_{p-1} \Delta e_{t-p+1} + w_t(C_T)(\rho - 1)e_{t-1} + \varepsilon_t, \quad (3.2)$$

where $\{\varepsilon_t\}$ is an i.i.d. sequence with mean zero, finite variance σ^2 and finite fourth moment. Notably, in Eq. (3.2), the error correction term $(\rho - 1)e_{t-1}$ is multiplied by the weight w_t. Let $\tilde{\rho}$ and $\tilde{t}_T$, respectively, be the least square estimator of ρ in (3.2) and the t test of $H0$: $\rho - 1 = 0$. Now, we define the following:

$$ADF_\rho(C_T) \equiv T_C(\tilde{\lambda} / \tilde{\sigma}_\varepsilon)(\rho - 1);$$
$$ADF_t(C_T) \equiv \tilde{t}_T \quad (3.3)$$

where $\tilde{\lambda} / \tilde{\sigma}_\varepsilon$ is consistently estimated by

$$\tilde{\lambda} / \tilde{\sigma}_\varepsilon = (1 - \tilde{\zeta}_1 - \cdots - \tilde{\zeta}_{p-1})^{-1}$$

where $\tilde{\zeta}_j$ is the least square estimator of ζ_j in (3.2).

Let T_C be a fixed subset of [0,1] such that $[T_C T] = C_T$ where $[T_C T] = \{[\tau T] : \tau \in T_C\}$ and $[\cdot]$ is the integer part of $\cdot$. Similarly, $[T_N T] = N_T$. For a given T_C, the statistics $Z_\rho(C_T)$, $Z_t(C_T)$, $ADF_\rho(C_T)$, and $ADF_t(C_T)$ converge to some limits as is shown by Kim (2003). Let $Z(C_T)$ be one of these four statistics. In practice, the period T_C or T_N is not known. In this case, we adopt the following approach: we get values of the statistic $Z(C_T)$ for all the possible candidates of C_T and take the smallest value of the statistic. The smallest value constitutes evidence against the null hypothesis. Thus, let

$$Z^*(C_T) = \inf C_T\, Z_T(C_T).$$

Our test statistics for detecting the existence of a segmented cointegration are obtained by applying the above *extremum* principle for tests of the Phillips-Perron type and ADF type: $Z^*_\rho(C_T)$, $ADF^*_\rho(C_T)$ and $Z^*_t(C_T)$, $ADF^*_t(C_T)$. Under some conditions, these statistics converge to corresponding limits under H_0 (Kim, 2003).

Kim (2003) presented critical values of the above statistics that have been obtained through simulation and an approximation method by Mackinnon (1991). Critical values of inf $Z(T_C)$ are obtained for some given upper limit of the length of T_N denoted by $\bar{l}(T_N)$.[2]

B. Identifying the Period of Non-cointegration

When a segmented cointegration is detected by the procedure discussed above, we are interested in the location of the period of non-cointegration. Let $N_T = ([\tau_0 T]+1, \ldots, [\tau_1 T])$ for some $\tau_i \in T_i$, $i = 0, 1$ where $T_0 \subset [0, 1)$ and $T_1 = (\sup T_0, 1]$, and $[\cdot]$ is the integer part. Note that we allow the possibility that the NC period lies at the beginning or the end of the sample. For a given $\tau = (\tau_0, \tau_1)$, let $\Lambda_T(\tau)$ be as in the following:

$$\Lambda_T(\tau) = \frac{[(\tau_1 - \tau_0)T] - 2\sum_{t\in N_T} et(CT)^2}{T_C^{-1}\sum_{t\in C_T} et(CT)^2}. \qquad (3.4)$$

In addition, let $\hat{\tau} = (\hat{\tau}_0, \hat{\tau}_1)$ be as

$$\hat{\tau} = \operatorname{argmax}_{\tau\in T} \Lambda_T(\tau) \qquad (3.5)$$

[2] That is, the critical values are obtained for the following statistic in Kim(2003):

$$\inf Z_\rho(T_c)_{\bar{\ell}(T_N)} = \inf_{T_C \in \{T_c : \ell(T_N) \le \bar{\ell}(T_N)\}} Z_\rho(T_C)$$

for $\bar{\ell}(T_N) = 0.3$. The objective is to reduce the computational requirements. The test (3.3) based on critical values of inf $Z\rho(T_C)_{\bar{\ell}(T_N)}$ is a conservative test when the actual test statistic $Z^*(C_T)$ in Eq. (3.4) is computed for $\ell_1(T_N)$ such that $\ell_1(T_N) \le \bar{\ell}(T_N)$.

where $T = T_0 \times T_1$. Under some conditions, the estimator $\hat{\tau}$ is a consistent estimator of τ, that is, $(\hat{\tau}_i - \tau_i) = o_p(1)$.

IV. Evidence for Segmented Cointegration

We apply the above segmented cointegration method for analyzing data from five east Asian countries, Korea, Indonesia, Malaysia, Thailand, and Singapore for the Fisher relation. The data are from International Financial Statistics for the period from 1989:Q1 to 2006:Q4, quarterly observations, for all the countries under study. We use the money market rate for the nominal interest rate and consumer price index to compute the inflation rate. We use $ADF_t(C_T)$ of (3.3) to detect the presence of SC and (3.5) to estimate the period of NC.

In practice, we adopt the following strategy: (1) We first estimate the period of NC by applying (3.5). We assume the minimum length of the NC period to be 8 quarters to insure some minimal number of observations for inference on data of the NC period. This assumption may lead to the result that the estimated NC period is longer than the true NC period, causing weaker evidence for non-cointegration in the NC period than the truth. (2) For the estimated SC period C_T obtained in (1) we get $ADF_t(C_T)$ in (*3.3*) and compare it with an appropriate critical value of the distribution of $ADF_t^*(C_T)$ provided in Kim (2003). If the value of $ADF_t(C_T)$ is less than the appropriate critical value, then the value of $ADF_t^*(C_T)$, which is less than or equal to $ADF_t(C_T)$, is obviously less than the critical value, leading to the conclusion of the presence of SC.

Tables 1 and 2 show our results. Note that the null ($H0$) of no cointegration is not rejected by the *ADFt* test for data of each country at the 5% level. However, we obtain the result that the null of a unit root for the real interest rate is rejected by allowing one segmentation for each country. Also, note the ADFt test strongly favors the non-cointegration hypothesis for the NC periods N_T (Table 1 for "Sample III). The estimated non-cointegration periods overlap the period of east Asian financial crisis, which started in 1997, with slight difference across countries. Our result implies that for data from the east Asian countries the Fisher relation prevails in the data period except the abnormal period of financial crisis.

Table 1

ADF(t) test for samples segmented by (3.5) in Table 2

Country	ADF(t)_Sample I	ADF(t)_Sample II	ADF(t)_Sample III
Indonesia	-2.8409	-3.8257	-2.5002
Korea	-1.8400	-3.1486	-1.6461
Malaysia	-1.9952	-3.0740	-1.1090
Thailand	-2.3664	-3.5742	-2.3385
Singapore	-2.7545	-3.4286	-1.8958

5% critical values: -2.895 (sample I), -2.897 (sample II), -.3.04 (sample III)
Sample I: whole sample
Sample II: segmented sample (whole sample \ NC)
Sample III: NC sample

Table 2

The Break-points estimated by (3.5)

Country	Break Points
Indonesia	1996:Q3 - 1999:Q1
Korea	1997:Q1 - 1998:Q4
Malaysia	1997:Q2 - 1999:Q1
Thailand	1997:Q1 – 1998:Q4
Singapore	1997:Q1 – 1998:Q4

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