AN EXAMINATION OF THE FISHER EFFECT IN DEVEVELOPING COUNTRY

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Acknowledgement

Abstract

This paper analyses Fisher effect in six developing countries including Arghentina, Brazil, China, India, Pakistan and Vietnam in the period 1997-2010. The unit root tests indicate mix order of cointegration for underlying variables. We employ the Johansen cointegration and ARDL bounds tests to identify the long term relationship. There is evidence which supports a weak form of Fisher effect for Argentina, India and Vietnam. The OLS results confirm a weak short run Fisher effect for Argentina. We also discuss the International Fisher effect in Vietnam.

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1. Introduction

Interest rate has always been most debating subject in the economy. Nowadays, almost all central banks implement some kinds of monetary policy to achieve their economic goals. Irving Fisher was among first people to address the concept of the real interest rate or premium for the compensation of expected loss in purchasing power. The Fisher effect states that nominal interest rate is equal to the expected inflation rate plus the expected real interest rate. If the expected real inflation rate remains constant over the period, the Fisher effect proposes a one for one relationship between nominal interest rate and expected inflation rate in the economy. Theoretically, short-run movement in nominal interest is useful to predict the future inflation. Many theoretical economic models assume Fisher effect holds. However, in practice the assumption of constant real interest is dubious. Therefore, the evidence of Fisher equilibrium can be an indication of an efficient market where nominal interest rate fully incorporates new relevant information of inflation.

Numerous empirical studies have tested the Fisher effect for many countries in different periods. Since the expected real interest rate and expected inflation can not be measured directly, the Fisher effect studies have applied several proxies for expected inflation. The early studies in the 1970s influenced by Fisher (1930) approach used distributed lag of the past on expected inflation rate. Most of these studies confirmed Fisher's findings of relationship between current nominal interest rate with the past inflation. Furthermore, Yohe and Karnosky (1969), Gibson (1972) and Lahiri (1976) typically found the evidence of a smaller time lag in expectations formation for the period after 1960 (Cooray 2002). Modern Fisher effect studies employed rational expectation assumption introduced by Fama (1975) where the current nominal interest reflects the future price level in efficient market. Mish (1992) was the first to set the trend of using cointegration test for analyzing the long term Fisher effect because nominal interest rate and inflation can be subjected to spurious regression. If the nominal interest rate and the inflation rate have first order of integration then two variables should cointegrate with a slope coefficient of unity so that the real interest rate is covariance stationary. (Beyer, Haug and Dewald 2009). The ordinary least square regression technique is only valid when two series are stationary at the level. Overall, the previous studies of long run Fisher effect showed different conclusions for countries and time period.

In the scope of this paper, we conduct the short and long-run relationship between the nominal interest rates and inflation in 6 emerging countries including Argentina, Brazil, China, India and Pakistan for the period 1997 to 2010 as the evidence of Fisher effect. Secondly, we investigate the International Fisher effect of Vietnam economy.

Six chosen countries have relatively volatile and high inflation rates with an average inflation rates are higher than 6% in the examining time period. Therefore, it is interesting to examine the movement of nominal interest rate under the influence of high inflation, market imperfection and government intervention.

We apply the rational expectations assumption to examine the Fisher effect. The real interest rate is assumed to be constant over time. We employ the Johansen coitegration procedure and autoregressive distributed lag (ARDL) bounds approach to determine the long-run relationship and Ordinary Least Square for short term relationship. The ARDL cointegration provides more reliable result compared to Johansen procedure because most underlying set of variables are mix orders of integration I(1) and I(0). Furthermore, the ARDL cointegration test is not affected by the potential structural break in unit root tests. The paper uses monthly short-term data. Short term maturity interest rate is considered to have a better predictive power of future inflation due to Vietnam high and volatile inflations within a year. However, Smant (2011) argued that short term rate is likely influenced by monetary policy. Thus, analyzing short-term real interest rates may only provide insights into government decision.

The paper is structured as follows. Section 2 is the review of Fisher effect and related literature on Fisher effect. Section 3 outlines the methodology. Section 4 describes the properties of our data. Section 5 provides the statistic test result and our interpretation. Section 6 discuss the International Fisher effect in the case of Vietnam Section 7 is the concluding remark.

2. Literature review

2.1 The Fisher effect

The relationship between nominal interest rate and real interest was firstly introduced by

Ivrin Fisher in his Appreciation and Interest (1896). In his latter work Theory of Interest (1930), the long-run Fisher relationship states that a permanent change in inflation will lead to an equal change in the nominal interest rate and the real interest rate will remain unaffected in the long run. Fisher parity is defined as the nominal interest rate equal to sum of the expected real interest rate and expected inflation

$$i_t = r_t + \pi_t^e \qquad (1)$$

where i_t is nominal interest rate, r_t is the real interest rate and π_t^e is expected rate of inflation. Here, all rates are continuously compounded variable. He also used the distribute lag structure to determine the correlation of coefficient of nominal interest and inflation. He found significant correlation coefficients between nominal interest rate and distributed lag of past inflation for Great Britain in 1820-1924 period and the United States in 1890-1927 period.

A number of studies have included more variables in their models such as wealth effect and tax. Those theoretical models explain the nominal interest might not move one to one with the expected inflation. Mundel (1963) and Tolbin (1965) argued that nominal interest rate responses less than unity to expect inflation because inflation reduces the money balance(Peng 1995).The procedure is also called "wealth effect" which the decrease in wealth leads to an increase in real saving and reduced investment demand. In other word, inflation would cause negative effect on real interest. On the other hand, Darby (1975) included the tax effect on his empirical model to investigate the Fisher effect relationship. He argued that the nominal interest must change in the range of 1.3 to 1.5 to the change in expected inflation. Crowder and Hofmann (1996) used Johansen cointegration techniques for the US in the period 1952-1991. They found that l percent increase in inflation yields 1.34 percent increase in the nominal interest rate. After adjusting for tax effect, the coefficient is insignificantly different to 1.

As previously mentioned, the modern studies of Fisher effect apply the rational expectation assumption. If the Fisher effect holds, there is a relationship between the current interest rate and the future inflation. Johansen VECM cointegration test and ARDL bounds testing techniques have become widely used techniques for long term relationship Fisher effect research. Some modern time series analysis of Fisher effect studies are presented below.

Mishkin(1992) performed common trend cointegraion test for the US monthly data in the period of 1953-1990. He found the common trend evidence for long-run relationship for post war period before Octorber 1979. However, there was no evidence of short-run relationship Fisher effect for the US. Mishkin(1995) used the treasury note in testing the Fisher hypothesis for Australia for the period of 1962-1993. He confirmed that Fisher effect only hold for the long-run but not for the short-run. He further suggested that short-run changes in interest rates reflect changes in monetary policy and longer-run levels indicate inflationary expectations.

Evans and Lewis (1995) implemented the shifts in inflation by a Markov switching model to investigate the Fisher effect for us data in the period 1947-1987. However, they did not reject the null hypothesis of one to one relationship between nominal interest rate and inflation in their stimulated models.

Shretha and Chen (1998) used Johansen cointegration technique for Canada, Japan, UK and US. They suggested the long-run Fisher effect holds for 4 countries. Nevertheless, the short-run Fisher relationship was only found in the cases of Japan and UK.

Atkins and Cole (2002) employed the ARDL bound testing approach to test the long-run Fisher relationship for Canada and the US in the period 1953-1999. They found the evidence supports the Fisher effect for both countries and the long-run coefficients are close to unity.

All previous studies focused on the developed countries. Berument and Jelassi (2002) performed a multi-country analysis of the Fisher Effect for both developed and developing countries. They suggested that the Fisher effect is likely hold for developed country than developing country. Nevertheless, their conclusion might largely influenced by the choices of data. More recently, Maghyereh and Zoubi (2006) using nonparametric test developed by Bierens (2000) to determine a robust nonlinear cotrending relationship between the interest rate and the inflation rate for Argentina, Brazil, Malaysia, Mexico, Korea and Turkey. They found strong evidence supports Fisher effect for all selected developing countries.

2.2 International Fisher effect:

The purchasing power of parity (PPP) measures the amount of resources needed to purchase one unit of good in different countries. The law of one price states that an identical good must have one price. A relative PPP implies that the foreign exchange rate will adjust to the change in price level in two countries.

$$\frac{S_{t+1}}{S_t} = \frac{1+\pi_{h,t}}{1+\pi_{f,t}} \quad \text{which can be written as} \quad \frac{S_{t+1}-S_t}{S_t} = \frac{\pi_{h,t}-\pi_{f,t}}{1+\pi_{f,t}} \qquad (2)$$

where S_t is the spot exchange rate at time t, $\pi_{f,t}$ is the inflation rate of foreign country at time t and $\pi_{h,t}$ is the home (domestic) inflation rate at time t

Apply Fisher parity (1) in to (2) where $i_{h,t} = \pi_{h,t} + r_{h,t}$ and $i_{f,t} = \pi_{f,t} + r_{f,t}$ to get International Fisher Effect (IFE) equation:

$$\frac{S_{t+1}-S_t}{S_t} = \frac{i_{h,t}-i_{f,t}}{1+i_{f,t}} \qquad (3)$$

This IFE equation proposes that change in the spot exchange rate between two countries is equal to the difference in their nominal interest rates. Another way of deriving IFE equation is based on the following interest rate parity assuming the real interest rates (yield return) are equal across countries.

$$1+i_h = \frac{S_{t+1}-S_t}{S_t}(1+i_f) \qquad (4)$$

The IFE gets much less attention from researchers. Sundqvist (2002) performed regression analysis to test IFE of US against Sweden, Japand and UK in the long run. The IFE is only significant for the countries pair of Japan and US. Utami and Inanga (2009) used the same method to test IFE of Indonesia against the Japan, Singapore, the UK and the US. They found a positive relationship between inflation rate differentials and changes in exchange rates over a five-year period, 2003-2008. However, their finding coefficients are much smaller than the hypothetical value of 1. The null hypothesis of the presence IFE was not tested.

3. Methodology

Long-run Fisher relationship

The traditional examination of long run Fisher effect involves the regression of observed inflation against observed nominal interest rate as follow:

$$\pi_t^m = \alpha_m + \beta_m i_t^m + u_t^m \qquad (4)$$

where π_t^m is the m-period (months) future of inflation rate at time t

i_t^m is the m-period future of interest rate at time t

The OLS method for examining Fisher effect includes the test of β_m. Granger and Newbold (1974) showed that the traditional Ordinary Least Square method is subjected to spurious regression which results in an inaccurate estimation of long term relationship. In Fisher proposition, either inflation or nominal interest rate can have unit roots. Mish (1992) was the first one to perform the unit roots test for each variable. He emphasized that the traditional regression analysis is not appropriate before conducting the Engle and Granger co-intergration test to examine the Fisher effect.

In this paper, I follow the Mishkin (1991) approach to test the long run and short run Fisher effect. I will first introduce the concept of stationary and order of intergration in which represt the nature of two serieses.

Stationary

A weak stationary time series has constant mean, constant variance and constant autovariances for each lag. Non-stationary or stochastic trend variable can lead to spurious regression. The two series (independent and dependent variables) must be stationary in order to perform a valid OLS estimation. We apply Augmented Dickey-Fuller (ADF) and Kwiatkowski–Phillips–Schmidt–Shin (*KPSS*) unit root test. The lag length of ADF unit root test is based on the Schwarz Info criterion with maximum lag is 12 (corresponding to possible 12 months lag). The performed KPSS is based on Bartlett Kernel method with Newey-West bandwidth selection. The null hypothesis of ADF is the variable has a unit root, in contrast the null hypothesis of KPSS assumes the series is stationary.

Order of integration and Co-integration

Order of integration indicates the minimum level of lag differences of a series to reach stationary. A stationary variable is integrated by order 0, denoted by I(0). It implies that an accurate long run estimation of OLS method should consist of all I(0) variables.

Co-integration has been intensively applied to examine the existence of long term relationship over the traditional OLS because many economic data such as inflation, interest rates and exchange rates appear to be persistent and potentially result in having unit roots. Co-integration means that underlying variables have a common stochastistic drift. According to Fisher equation, the nominal interest rate should move one on one with inflation.

The two most widely used method of testing co integration for Fisher effect are the Engle-Granger two-step method and the Johansen procedure. In practice both of these methods require all variables to be integrated in order of 1 or I(1). However, Hjalmarssonand and Österholm(2007) emphasized that stationary variable I(0) is not an issue for Johansen procedure which is consistent with Johansen (1995) arguments of "little need to pre-test the variables in the system to establish their order of integration".

The Engle-Granger (EG) approach operates on the residual of an estimated model. First, we estimate the regression (4) of inflation against nominal interest rate:

$$\pi_t^m = \alpha_m + \beta_m i_t^m + u_t^m$$

Next, we apply the ADF unit root test on the residual of previous equation. The null hypothesis is there is a unit root in the potentially co-integrating regression residual (Brook,2002). The rejection of previous null hypothesis would imply a stationary linear combination of two I (1) series. Thus, we can conclude two variables are co-integrated.

The Johansen likelihood estimation technique is more powerful co-integration test than EG approach. It allows more than one co-integrating relationship from variables. However, in this research, there can only be one co-integrating relationship since Fisher effect testing involves only two variables (inflation and nominal interest rate). Therefore, EG is still applicable. We perform EG and Johansen cointegration test for I(1) regressors when each individual has 1 order of integration. It also implies nominal interest rate and inflation should be non-stationary at the level.

The Johansen procedure is based on Vector Autoregression (Var) with p lags (order of p)

$$y_t = \mu + A_1 y_{t-1} + A_2 y_{t-2} + ... + A_p y_{t-p} + u_t \qquad (5)$$

Or $$y_t = \mu + \sum_{j=1}^{p} A_j y_{t-j} + u_t \qquad (6)$$

where y_t is 2x1 vector of two I(1) variables, inflation and nominal interest rate; 2x2 matrix A_i is the VAR parameter for lag I; u_t is 2x1 vector of white noise error term. The VAR equation (5) can be transformed into a vector error correction model (VECM) as follow:

$$\Delta y_t = \mu + \Pi y_{t-1} + \sum_{j=1}^{p-1} \Gamma_j \Delta y_{t-j} + u_t \qquad (7)$$

where $\Pi = \sum_{j=1}^{p} A_i - I$ (8) and $\Gamma = -\sum_{h=j+1}^{p} A_j$ (9)

The Johansen test is based on the rank of the coefficient Π matrix via its eigenvalues λ which contains the information of long run respones (Brook 2002). If there exist the cointegrated relationship, the number of cointegrating relationships, denoted as r < n where n is the number of variables.

The Π matrix can be represent as

$$\Pi = \alpha \beta ' \qquad (10)$$

where α and β are n x r matrices. α is the adjustment parameters in the VECM and β is a cointegrating vector. (Hjalmarssonand and Österholm,2007)

In this case, testing for long-run Fisher relationship is equivalent to prove one linearly independent column in Π. Johansen proposes the Trace test and the Maximum Eigenvalue test to determine the rank of Π as below:

$$\lambda_{trace}(r) = -T \sum_{j=r+1}^{n} \ln(1 - \hat{\lambda}_1) \qquad (11)$$

$$\lambda_{\max}(r, r+1) = -T \ln(1 - \hat{\lambda}_{r+1}) \qquad (12)$$

The null hypothesis of the trace test is that the number cointegrating vectors is less or equal to r against the alternative hypothesis of more than r cointegrating vectors. On the other hand, the null hypothesis of maximum eigenvalue test is number of cointegrating vector is r against the alternative of r+1 (Brook, 2002)

The lag intervals of Johansen cointegration test is based on Schwarz information criterion. It also is important to point out that the outcome of cointegration test can be highly sensitive to the number of selected maximum lag as well as lag length criterion. Furthermore, the cointegration relationship is more likely to occur with lower selected value of lag interval (or the order of p).

In the next section, we will explain the autoregressive distributed lag (ARDL) bounds testing approach introduced by Pesaran et al. (2001) for long-run cointegration. This technique does not require all underlying variables to have the same order of integration unlike OLS and EG/ Johansen co-integration. In our case, the inflation rate and nominal interest rate can either be I(1) and I(0) or mutually co-integrated to apply ARDL regression. We follow Atkins and Coe (2002) method to perform ARDL bound test for long run Fisher effect. In the last OLS regression, we use inflation as dependent variable instead of nominal interest.

Consider the following VECM transformation (7) of restricted VAR from equation (5)

$$\Delta y_t = \mu + \Pi y_{t-1} + \sum_{j=1}^{p-1} \Gamma_j \Delta y_{t-j} + u_t$$

Where $y_t = \left[\pi_t^m \quad i_t^m \right]'$ π_t^m is the m months future of inflation rate at time t, i_t^m is the m months future of interest rate at time t. The vector of error terms is

$u_t = \left[u_{\pi,t}^m \quad u_{i,t}^m \right]' \sim \mathrm{IN}(0, \Omega)$ where Ω is positive definite $\Omega = \begin{pmatrix} \omega_{\pi\pi} & \omega_{\pi i} \\ \omega_{i\pi} & \omega_{ii} \end{pmatrix}$

$$\Gamma = \begin{pmatrix} A_{\pi\pi,j} & A_{\pi i,j} \\ A_{i\pi,j} & A_{ii,j} \end{pmatrix} = -\sum_{h=j+1}^{p} A_j \qquad (13)$$

The long-run multiplier matrix and is given by:

$$\Pi = \begin{pmatrix} \Pi_{\pi\pi} & \Pi_{\pi i} \\ \Pi_{i\pi} & \Pi_{ii} \end{pmatrix} = \sum_{j=1}^{p} A_j - I \qquad (14)$$

I is 2x2 matrix. The diagonal elements of the matrix are unrestricted, so the selected series can be either I(0) or I(1). If $\lambda_{ii} = 0$, then *i* is I(1). In contrast, if $\lambda_{ii} < 0$, then *i* is I(0). (Atkins and Coe,2002). Under assumption 3 introduced by Pesaran *et al*. (2001), we include a restriction $\Pi_{i\pi} = 0$. As the consequence of $\Pi_{i\pi} = 0$, the inflation rate would have no long-run effect on nominal interest rate. Therefore, nominal interest rate is chosen to be long-run forcing for inflation. The VECM equation (7) under the assumption of $\Pi_{i\pi} = 0$ can be written as:

$$\Delta \pi_t = \alpha + \varphi \pi_{t-1} + \delta i_{i-1} + \sum_{j=1}^{p-1} \tau_{\pi,j} \Delta \pi_{t-j} + \sum_{j=1}^{q-1} \rho_{i,j} \Delta i_{t-j} + \omega \Delta i_t + u_t \qquad (15)$$

where $\alpha = \mu_\pi - \omega_i, \varphi = \Pi_{\pi\pi}, \delta = \Pi_{\pi i} - \omega_{ii}, \tau_{\pi,j} = A_{\pi\pi,j} - \omega A_{i\pi,j}, \rho_{i,\pi} = A_{\pi i,j} - \omega A_{ii,j}$ and $\omega = \omega_{\pi i} / \omega_{ii}$

Equation (7) can be interpreted as an ARDL(p,q) model where p and q are the number of lagged differences of inflation and nominal interest respectively. The two coefficient of φ and δ represent the long term relationship. According to Pesaran et al. (2001), the joint null hypothesis of no long-run relationship would be H_0: $\varphi = \delta = 0$ against the alternative H_1: $\varphi \neq 0, \delta \neq 0$. Under the alternative hypothesis, the stable long-run relationship between inflation and nominal interest rate can be described as:

$$\pi_t = \theta_0 + \theta_1 i_t + v_t \qquad (16)$$

Where $\theta_0 = -\alpha / \varphi$ and $\theta_1 = \delta / \varphi$ and v_t is the mean zero stationary process.

Pesaran et al. (2001) suggested F statistic test of OLS regression to examine the join significant of φ and δ . However, the obtained F-statistic values need to be compared with two sets of asymptotic critical values. The lower bounds assume all underlying variables are I(0) and the upper bounds corresponds I(1) regressors. If the computed F statistic is less than lower bounds, we can not reject the null hypothesis. If the F-statistic falls between the lower and the upper bound values, the result is conclusive. If the F-statistic if greater than the upper bound values, the null hypothesis will be rejected and inflation and nominal interest rates are cointegrated in the long-run. The (ARDL) lagged differences p and q are determined by Schwarz information criterion. The critical value bounds are selected in table C1.iii from Pesaran *et al.* (2001).

Short run Fisher relationship

According to Mishkin (1991), a short-run Fisher effect indicates that a change in the interest rate is associated with an immediate change in the expected inflation rate. The short-run relationship can be represented by the following regression:

$$\Delta \pi_t^m = \alpha + \beta \Delta i_t^m + v_t^m \qquad (17)$$

where Δ is the first difference operator of inflation and nominal interest ($\Delta \pi_t^m = \pi_t^m - \pi_{t-1}^m$ and $\Delta i_t^m = i_t^m - i_{t-1}^m$); v_t^m is the residual term

The short-run Fisher effect involves the significance coefficient test of α and β where $\alpha = 0$ and $\beta = 1$. Shretha and Chen (1998) suggested that the OLS estimation technique will be sufficient to estimate the short-run relationship between nominal interest rate and inflation. Because even if the two series are non stationary the first differenced series will be stationary. Our unit root tests also confirm this idea since all variables are either I(0) or I(1).

International Fisher effect

IFE states that the difference in nominal interest rate between two countries results in the expected change in the spot exchange rate. We apply the following regression:

$$\Delta y = \alpha + \beta \Delta x + \varepsilon \qquad (18)$$

Where $\Delta x = \frac{i_{h,t} - i_{f,t}}{1 + i_{f,t}}$ and $\Delta y = \frac{S_{t+1} - S_t}{S_t}$

The international Fisher effect will holds if the coefficients $\alpha = 0$ and $\beta = 1$. In this case, an x% increase in domestic interest rate will reflect on same percentage increase in exchange rate.

We also test the hypothesis of the nominal interest rate differential has positively relationship on change in the exchange rate $\beta > 0$.

4. Data

The data has been collected from IMF International Financial Statistic and Data Stream for the period of 1997-2010. Some series have been converted into monthly series by taking the average of daily observations.

The Argentina nominal interest rates consist of 2 and 3 months deposits, 1 month lending rate and, 6 months interbank and 1 year inter bank.

The Brazil nominal interest rates include 1month Treasury Bill, 2 months deposit and 3 months Bank Certificate Deposit (BCD).

The 3 months Treasury bill, 1 year lending rate, 1 year deposit, 3 year deposit and 5 year deposit are chosen to be China interest rate.

India nominal interest rates include 1month deposit, 6month deposit, 3 month Treasury bill, 1 year and 5 year Government bond.

KIBOR, Treasury bill and REPO are used as approximation of nominal interest rate in Pakistan. KIBOR stands for The Karachi Interbank Offered Rate.

The Vietnam nominal interest rates consist of 3-month deposit rates of four largest states owned commercial banks, 1-year treasury bill, 1month inter bank, 3 month inter bank and short-term bank lending rate. The lending rate is average short-term (less than 12 months) working capital loans of four large state-owned commercial banks. The authors assume that these loans have a maturity of 6 months. Nominal interest of foreign countries includes 3-month EU interest in France, 3-month UK Treasury bill and 3-month US Treasury bill.

The annualized continuously compounded inflation rates are calculated from monthly CPI observations. The realized m-months inflation rate is computes as

$$\pi_{t+1}^{m} = \frac{12\ln\left(CPI_{t+m} / CPI_{t}\right)}{m} x100\% \qquad (19)$$

The timing of the variables is constructed similarly to Mishkin(1992). Three months January inflation calculated from December and March CPI data.

5. Empirical result

We apply 95% confident intervals for all hypothesis testing (5% critical values). The maximum selected lag lengths are 12.

5.1 Unit root testing

First, we conduct the unit root test at the level and first difference for each series. The null hypothesis of the ADF test is that the variable contains a unit root. The null hypothesis of the KPSS test is that the variable is stationary. Perron (1989) proposed that the structural break or regime shift can reduce the power of unit root testing. In our case, the effect of structural break is neglected due to the short time span of analysis. Bearing in mind, the Johansen procedure for set of I(0) and I(1) variables can still be applicable because each I(0) variable will reveal itself as a cointegrating vector (Hjalmarssonand and Österholm,2007). Moreover, the ARDL bounds testing cointegration approach can be applied regardless the same order of integration condition.

The results are reported in the following tables (Table 1x)

* indicates the series contain a unit root at 5% critical value. The 5% critical value of ADF and KPSS tests are 2.904 and 0.463 respectively.

Table 1a: Unit root tests of nominal interest rates and inflation in Vietnam and China

Country	Variable	Level		First difference		Order of integration
		ADF t-statistic (p-value)	KPSS LM-stat	ADF t-statistic (p-value)	KPSS LM-stat	
Vietnam	Intern_1m	-3.63 (0.006)	0.35	x	x	I(0)
	Inflation_1m 1999M2 - 2010M11	-2.18* (0.216)	0.66*	-3.36 (0.014)	0.26	I(1)
	Deposit_3m	-3.71 (0.048)	0.50*	x	x	I(0)
	Inflation_3m 1997M2 - 2010M9	-2.77* (0.066)	0.57*	-5.45 (0.000)	0.03	I(1)
	Intern_3m	-3.87 (0.003)	0.36	x	x	I(0)
	Inflation_3m 1999M2 – 2010M10	-2.83* (0.056)	0.76*	-3.22 (0.021)	0.05	I(1)
	Bank lending	-3.60 (0.007)	0.29	x	x	I(0)
	Inflation_6m 1997M2 – 2010M7	-1.92* (0.325)	0.61*	-4.97 (0.000)	0.02	I(1)
	T-Bill	-4.21 (0.0009)	0.46*	x	x	I(0)
	Inflation_12m 1997M3 - 2010M1	-1.61* (0.477)	0.67*	-3.25 (0.019)	0.04	I(1)
China	TBond	-3.39 (0.012)	0.56*	x	x	I(0)
	Inflation_3m 1997M1 - 2010M12	-3.52 (0.009)	0.07	x	x	I(0)
	Lending rate 12M	-3.44 (0.01)	0.49*	x	x	I(0)
	Inflation_12m 1997M1 - 2010M12	-3.02 (0.035)	0.11	x	x	I(0)
	Deposit 2y	-3.32 (0.020)	0.42	x	x	I(0)
	Inflation_24m 1997M1 - 2009M12	-4.15 (0.001)	0.09	x	x	I(0)
	Deposit_3y	-2.90 (0.047)	0.40	x	x	I(0)
	Inflation_36m 1997M1 – 2008M12	-3.12 (0.027)	0.19	x	x	I(0)
	Deposit_5y	-3.12 (0.027)	0.64*	x	x	I(0)
	Inflation_60m 1997M5 - 2006M12	-2.87* (0.052)	0.11	x	x	I(0)

Table 1b: Unit root tests of nominal interest rates and inflation in India and Pakistan

Country	Variable	Level		First difference		Order of integration
		ADF t-statistic (p-value)	KPSS LM-stat	ADF t-statistic (p-value)	KPSS LM-stat	
India	Deposit_1m	-1.71* (0.424)	0.85*	-6.57 (0.000)	0.37	I(1)
	Inflation_1m 1997M1 - 2010M12	-9.31 (0.000)	0.32	x	x	I(0)
	T Bill_3m	-1.80* (0.378)	0.57*	-12.07 (0.000)	0.08	I(1)
	Inflation_3m 1997M1 - 2010M12	-7.11 (0.000)	0.35	x	x	I(0)
	Deposit_6m	-2.28* (0.179)	0.92*	-12.05	0.34	I(1)
	Inflation_6m 1997M1 - 2010M12	-2.27* (0.183)	0.39	x	x	I(0)
	G Bond 1y	-1.90* (0.330)	0.88*	-6.37 (0.000)	0.09	I(1)
	Inflation_12m 1997M1 – 2010M12	-2.59* (0.009)	0.48*	-9.90 (0.000)	0.05	I(1)
	G Bond_5y	-1.94* (0.312)	1.08*	-10.36 (0.000)	0.35	I(1)
	Inflation_60m 1997M5 - 2006M12	-0.26* (0.975)	0.89*	-8.00 (0.000)	0.70*	I(1)
Pakistan	KIBOR_1m	-0.27* (0.92)	1.05*	-8.00 (0.000)	0.18	I(1)
	Inflation_1m 2002M9 - 2010M12	-6.93 (0.000)	0.56*	x	x	I(0)
	KIBOR_3m	-0.61* (0.861)	1.04*	-6.59 (0.000)	0.15	I(1)
	Inflation_3m 2002M9 – 2010M12	-2.89* (0.050)	0.53*	-7.69 (0.000)	0.10	I(1)
	REPO_6m	-2.02* (0.276)	0.37	-10.77 (0.000)	0.38	I(1)
	Inflation_6m 1997M1 - 2010M3	-1.41* (0.578)	0.91*	-8.90 (0.000)	0.06	I(1)
	T Bill 6m	-1.60* (0.483)	0.35	-10.61 (0.000)	0.32	I(1)
	Inflation_6m 1997M1 – 2010M12	-1.62* (0.467)	0.92*	-9.40 (0.000)	0.04	I(1)
	KIBOR_12m	-2.55* (0.107)	1.01*	-5.64 (0.000)	0.30	I(1)
	Inflation_12m 2004M5 - 2010M12	-2.67* (0.084)	0.37	-4.06 (0.002)	0.09	I(1)

Table 1c: Unit root tests of nominal interest rates and inflation in Argentina and Brazil

Country	Variable	Level		First difference		Order of integration
		ADF t-statistic (p-value)	KPSS LM-stat	ADF t-statistic (p-value)	KPSS LM-stat	
Argentina	Lending rate 1M	0.20* (0.972)	0.14	x	x	I(0)
	Inflation_1m 1997M1 - 2010M12	-4.11 (0.001)	0.27	x	x	I(0)
	Deposit 2m	-2.63* (0.089)	0.12	-9.72 (0.000)	0.04	I(1)
	Inflation_2m 1997M1 - 2010M10	-3.73 (0.004)	0.27	x	x	I(0)
	Deposit 3m	-2.73* (0.071)	0.12	-8.36	0.03	I(1)
	Inflation_3m 1997M1 – 2010M12	-3.61 (0.007)	0.27	x	x	I(0)
	InterB 6M	-2.20* (0.206)	0.16	-7.31 (0.000)	0.06	I(1)
	Inflation_6m 1997M5 – 2010M12	-3.55 (0.008)	0.25	x	x	I(0)
	InterB 12M	-1.86* (0.351)	0.12	-12.12 (0.000)	0.06	I(1)
	Inflation_12m 1997M5 - 2010M12	-4.35 (0.001)	0.27	x	x	I(0)
Brazil	TBill	-1.18* (0.682)	1.42*	-7.52 (0.000)	0.09	I(1)
	Inflation_1m 1997M1 - 2010M12	-6.11 (0.000)	0.19	x	x	I(0)
	Deposit 2m	-2.25* (0.019)	1.24*	-8.55	0.03	I(1)
	Inflation_2m 1997M1 - 2010M12	-3.29 (0.017)	0.19	x	x	I(0)
	BCD	-1.67* (0.445)	1.05*	-4.26 (0.001)	0.06	I(1)
	Inflation_3m 2000M5 - 2010M12	-2.54* (0.107)	0.45*	-6.84 (0.000)	0.03	I(1)

* indicates the series contain a unit root at 5% critical value. The ADF test has H_0 : ym ~ $I(0)$. The KPSS test has H_1: ym~ $I(0)$. The 5% critical value of ADF and KPSS tests are 2.904 and 0.463 respectively.

In table 1a, the unit root results test show conflicting evidence for Deposit_3m and T-Bill series at the level. Therefore, we perform Phillip-Perron (PP) unit root test to determine the order of integration these two series. For instance, PP test result does not reject the null hypothesis of non-stationary for Deposit_3m at the level which implies the series has first order of integration[1]. Most of nominal interest rate series are stationary I(0) except for Deposit_3m. In addition, all inflation series contain a unit root and first order of integration I(1) excluding inflation_3m (1999M2 - 2010M7).

The ADF and KPSS tests report conflicting unit root results for there series in China. After taking into account the result of Phillip-Perron test, all nominal interest rates and inflation series are considered to be stationary at the level. Therefore, the OLS technique will be applied to verify the long run Fisher relationship in China.

For India data in table 1b, the two nominal interest rates,1 year Government bond and 5 year Government bond, exhibit the same I(1) with their respective m future of inflation. Other cases contain mixed orders of integration where nominal interest rate series are I(1) and inflation series are I(0). The PP unit root test is used to find the order of integration of *Inflation_6m* and *Inflation_60m* series.

In the case of Pakistan, the KIBOR_1m does not have the same order of integration with 1 month future inflation. Other nominal interest rate inflation series are I(1). Overall, the results of ADF tests tend to support unit root in both variables. KPSS test gives opposite unit root results for *REPO_6m, TBill_6m and Inflation_12m*.

The table 1c shows that result of unit root tests for inflation and interest rate in Argentina. All inflation rates are found to be stationary at the level. There are conflicting results of two unit roots tests for all nominal interest rate at the level. Subsequently, we perform additional Phillip-Perron (PP) unit root test to choose the order of integration. The PP test suggests most of nominal interest rate series to have first order of cointegration except for *Lending rate*.

For Brazil data, the unit root test results show all nominal interest rates to have first order of integration, I(1). The 3 month inflation series has first order of integration and other inflation series are stationary at level. In this section, the long run Fisher effect for six countries will be analyzed by various co-integration methods

[1] The p-values of PP unit root test for Deposit_3m and T-Bill are 0.0859 and 0.0101 respectively

5.2 Long run Fisher effect

Table 2a.i: Johansen maximum likelihood estimation Fisher effect in Vietnam

Variables $y_t = \left[\pi_t^m \quad i_t^m\right]'$ Sample period	Lag length of VAR (SIC)	H_0	λ_{trace} (p-value)	$\lambda_{\max}$ (p-value)	No. of co-vectors	Estimated coefficient β (Standard error)
Inflation_3m, Deposit_3m	2	r = 0	52.59	46.12	2	-1.92
1997M2 - 2010M9 (i)			(0.000)	(0.000)		(0.39)
		r ≤ 1	6.47	6.47		
			(0.011)	(0.011)		
Inflation_3m, Intern_3m	2	r = 0	46.11	36.72	2	-2.41
1999M2 – 2010M10 (ii)			(0.0000)	(0.0000)		(0.57)
		r ≤ 1	9.39	9.39		
			(0.002)	(0.002)		
Inflation_1m, Intern_1m	2	r = 0	45.23	30.95	2	-2.66
1999M2 - 2010M11 (iii)			(0.000)	(0.000)		(0.63)
		r ≤ 1	14.28	14.28		
			(0.000)	(0.000)		
Inflation_12m, T-bill	2	r = 0	18.57	17.89	1	-4.72
1997M3 - 2010M1 (iv)			(0.017)	(0.013)		(0.97)
		r ≤ 1	0.68	0.68		
			(0.409)	(0.409)		
Inflation_6m, Lending rate	7	r = 0	15.69	11.22	2_{trace}	-3.68
1997M2 – 2010M7 (v)			(0.047)	(0.143)	0_{max}	(1.12)
		r ≤ 1	4.46	4.46		
			(0.035)	(0.035)		

Table 2a.ii: ARDL Bounds Testing for long-run Fisher effect in Vietnam

Variable	Lag difference (p,q)	F-statistic	Co-integration		
			99%	95%	90%
Inflation_3m, Deposit_3m 1997M2 - 2010M9	(11,3)	4.10	X	x	inconclusive
Inflation_3m, Intern_3m 1999M2 – 2010M10	(11,1)	5.90	X	yes	yes
Inflation_1m, Intern_1m 1999M2 - 2010M11	(12,1)	3.21	X	x	x
Inflation_12m, T-bill 1997M3 - 2010M1	(7,1)	3.76	X	x	x
Inflation_6m, Lending rate 1997M2 – 2010M7	(12,2)	3.11	X	x	x

Table 2b: OLS regression for long-run Fisher effect in China

Variable	Coefficient		Long-term Fisher Relationship
	α	β	
Inflation_3m, TBond 1997M1 - 2010M12	2.25 (0.003)	-0.63 (0.002)	no
Inflation_1y, Lending 1997M1 – 2010M12	9.36 (0.000)	-1.48 (0.000)	no
Inflation_24m, Deposit_2y 1997M1 - 2019M12	2.39 (0.863)	-0.64 (0.000)	no
Inflation_36m, Deposit_3y 1997M5 - 2008M12	1.19 (0.000)	-0.26 (0.000)	no
Inflation_60m, InterB_5y 1997M5 – 2006M12	1.05 (0.000)	-0.19 (0.000)	no

Table 2c.i: Johansen maximum likelihood estimation Fisher effect in India

Variables $y_t = \left[\pi_t^m \quad i_t^m\right]'$ Sample period	Lag length of VAR (SIC)	H_0	λ_{trace} (p-value)	$\lambda_{\max}$ (p-value)	No. of co-vectors	Estimated coefficient β (Standard error)
Inflation_1m, Deposit_1m 1997M1 - 2010M12	2	r = 0	74.25 (0.000)	62.92 (0.000)	2	-2.10 (0.82)
		r ≤ 1	11.33 (0.000)	11.33 (0.000)		
Inflation_3m, T-Bill 3m 1999M1 – 2010M12	5	r = 0	32.92 (0.000)	28.83 (0.000)	2	0.23 (0.45)
		r ≤ 1	4.10 (0.043)	4.10 (0.043)		
Inflation_6m, Deposit 6m 1999M1 - 2010M12	10	r = 0	16.37 (0.037)	13.78 (0.064)	1	-0. 56 (0.55)
		r ≤ 1	2.60 (0.107)	2.60 (0.107)		
Inflation_12m, G-bond 1y 1997M1 - 2010M12	2	r = 0	8.21 (0.442)	4.84 (0.761)	0	0.25 (0.72)
		r ≤ 1	3.37 (0.066)	3.37 (0.066)		
Inflation_60m, G-Bond 5y 1997M1 – 2006M12	2	r = 0	14.19 (0.078)	12.32 (0.10)	0	-12.28 (3.59)
		r ≤ 1	1.87 (0.17)	1.87 (0.17)		

Table 2c.ii: ARDL Bounds Testing for long-run Fisher effect in India

Variable	Lag difference (p,q)	F-statistic	Co-integration 95%	Coefficient θ_0	Coefficient θ_1 or β '
Inflation_1m, Deposit_1m 1997M1 - 2010M12	(12,11)	2.52	no	6.94	0.02
Inflation_3m, T-Bill 3m 1997M1 – 2010M12	(0,4)	25.94	yes	11.59	0.78
Inflation_6m, Deposit 6m 1997M1 - 2010M12	(0,6)	18.08	yes	3.51	-0.48
Inflation_12m, G-bond 1y 1997M1 - 2010M12	(4,1)	2.45	no	9.81	0.47
Inflation_60m,G-Bond 5y 1997M1 – 2006M12	(8,1)	3.31	no	26.05	2.24

The upper bounds for critical values are 4.78, 5.73 and 7.84 for 10%, 5% and 1% respectively.
The lower bounds critical values are 4.04, 4.94 and 6.84. The lag order selected using the SIC.

Table 2d.i: Johansen maximum likelihood estimation Fisher effect in Pakistan

Variables $y_t = \left[\pi_t^m \quad i_t^m\right]'$ Sample period	Lag length of VAR (SIC)	H_0	λ_{trace} (p-value)	$\lambda_{\max}$ (p-value)	No. of co-vectors	Estimated coefficient β (Standard error)
Inflation_1m, KIBOR_1m	2	r = 0	47.07	46.57	1	-0.72
2002M9 - 2010M12			(0.000)	(0.000)		(0.24)
		r ≤ 1	0.49	0.49		
			(0.482)	(0.482)		
Inflation_3m, KIBOR_3m	2	r = 0	41.51	40.40	1	-0.69
2002M9 - 2010M12			(0.000)	(0.000)		(0.19)
		r ≤ 1	1.12	1.12		
			(0.290)	(0.290)		
Inflation_6m, REPO_6m	2	r = 0	17.77	15.24	1	-0.75
1997M1 - 2010M3			(0.022)	(0.035)		(0.38)
		r ≤ 1	2.53	2.52		
			(0.111)	(0.111)		
Inflation_6m, T-Bill_6m	2	r = 0	19.90	17.34	1	-0.58
1997M1 - 2010M12			(0.442)	(0.761)		(0.33)
		r ≤ 1	2.55	2.55		
			(0.066)	(0.066)		
Inflation_12m,KIBOR_12m	2	r = 0	12.36	7.58	0	-7.11
2004M5 – 2006M12			(0.140)	(0.422)		(2.55)
		r ≤ 1	4.78	4.78		
			(0.029)	(0.029)		

Table 2d.ii: ARDL Bounds Testing for long-run Fisher effect in Pakistan

Variable	Lag difference (p,q)	F-statistic	Co-integration 95%	Coefficient θ_0	Coefficient θ_1 or β '
Inflation_1m, KIBOR_1m 2002M9 - 2010M12	(0,0)	12.87	yes	5.34	-0.54
Inflation_3m, KIBOR_3m 2002M9 - 2010M12	(1,1)	6.32	yes	6.98	-0.36
Inflation_6m, REPO_6m 1997M1 - 2010M3	(1,6)	2.85	no	10.31	0.25
Inflation_6m, T-Bill_6m 1997M1 - 2010M12	(1,6)	2.65	no	9.81	0.47
Inflation_12m,KIBOR_12m 2004M5 – 2006M12	(1,1)	1.43	no	12.10	-0.12

The upper bounds for critical values are 4.78, 5.73 and 7.84 for 10%, 5% and 1% respectively.
The lower bounds critical values are 4.04, 4.94 and 6.84. The lag order selected using the SIC.

Table 2e.i: Johansen maximum likelihood estimation Fisher effect in Argentina

Variables $y_t = \left[\pi_t^m \quad i_t^m\right]'$ Sample period	Lag length of VAR (SIC)	H_0	λ_{trace} (p-value)	$\lambda_{\max}$ (p-value)	No. of co-vectors	Estimated coefficient β (Standard error)
Inflation_2m, Deposit_2m 1997M1 – 2010M12	12	r = 0	15.86 (0.044)	13.02 (0.078)	1	0.45 (0.33)
		r ≤ 1	2.84 (0.092)	2.84 (0.092)		
Inflation_3m, Deposit_3m 1997M1 - 2010M12	10	r = 0	15.15 (0.056)	11.15 (0.132)	0	0.53 (0.38)
		r ≤ 1	3.67 (0.056)	3.67 (0.056)		
Inflation_6m, InterB_6m 1997M5 - 2010M12	10	r = 0	11.68 (0.173)	9.21 (0.269)	0	0.19 (0.20)
		r ≤ 1	2.47 (0.116)	2.47 (0.116)		
Inflation_12m, InterB_12m 1997M5 – 2010M12	4	r = 0	13.89 (0.086)	10.03 (0.086)	0	-0.19 (0.09)
		r ≤ 1	3.86 (0.050)	3.86 (0.050)		

Table 2e.ii: ARDL Bounds Testing for long-run Fisher effect in Argentina

Variable	Lag difference (p,q)	F-statistic	Co-integration 95%	Coefficient θ_0	Coefficient θ_1 or β'
Inflation_1m, Lending rate 1997M1 - 2010M12	(1,4)	9.83	yes	7.64	0.22
Inflation_2m, Deposit_2m 1997M1 – 2010M12	(4,4)	6.47	yes	9.52	0.23
Inflation_3m, Deposit_3m 1997M1 - 2010M12	(7,4)	6.96	yes	5.03	-0.17
Inflation_6m, InterB_6m 1997M5 - 2010M12	(7,3)	2.01	x	x	x
Inflation_12m, InterB_12m 1997M5 – 2010M12	(2,0)	8.11	yes	8.02	0.03

The upper bounds for critical values are 4.78, 5.73 and 7.84 for 10%, 5% and 1% respectively. The lower bounds critical values are 4.04, 4.94 and 6.84. The lag order selected using the BIC. See Table C1.iii Case III Pesaran et al (2001) for more general values.

Table 2f.i: Johansen maximum likelihood estimation Fisher effect in Brazil

Variables $y_t = \left[\pi_t^m \quad i_t^m\right]'$ Sample period	Lag length of VAR (SIC)	H_0	λ_{trace} (p-value)	$\lambda_{\max}$ (p-value)	No. of co-vectors	Estimated coefficient β (Standard error)
Inflation_1m, TBill 1997M1 – 2010M12	2	r = 0	37.17 (0.000)	32.07 (0.000)	2	0.004 (0.08)
		r ≤ 1	5.10 (0.024)	5.10 (0.024)		
Inflation_2m, Deposit_2m 1997M1 - 2010M12	4	r = 0	23.24 (0.003)	18.71 (0.009)	1	0.05 (0.13)
		r ≤ 1	4.53 (0.033)	4.53 (0.033)		
Inflation_3m, BCD 1997M5 – 2010M12	5	r = 0	20.64 (0.008)	18.22 (0.011)	1	-0.40 (0.09)
		r ≤ 1	2.42 (0.120)	3.42 (0.120)		

Table 2f.ii: ARDL Bounds Testing for long-run Fisher effect in Brazil

Variable	Lag difference (p,q)	F-statistic	Co-integration 95%	Coefficient θ_0	Coefficient θ_1 or β '
Inflation_1m, TBill 1997M1 - 2010M12	(7,6)	3.45	no	6.15	0.03
Inflation_2m, Deposit_2m 1997M1 – 2010M12	(6,2)	5.36	inconclusive	5.52	-0.03
Inflation_3m, BCD 2000M5 – 2010M12	(6,3)	7.38	yes	1.98	-0.30

The upper bounds for critical values are 4.78, 5.73 and 7.84 for 10%, 5% and 1% respectively.
The lower bounds critical values are 4.04, 4.94 and 6.84. The lag order selected using the BIC.
See Table C1.iii Case III Pesaran et al (2001) for more general values.

In the case of Vietnam, the EG cointegration can only apply to *Deposit_3m* and 3 months future of inflation under its strictly condition of both I(1) underlying variables. The p-value of the ADF unit root test based on Akaike Information Criterion (AIC) of the regression residual is equal to 0.0258. The null hypothesis of unit root is rejected hence the two series are cointegrated. We also carry out other unit root tests PP and KPSS; the result of no cointegration from KPSS test is inconsistent with ADF and PP. Therefore, EG cointegration result does not rule out the possibility of no long-run relationship between *deposit_3m* and *inflation_3m*. Table 3a.i presents the Johansen cointegration analysis for five sets of variables. We only find the evidence which supports the existence of cointegration between the case of T-Bill and inflation_3m. In other cases (i),(ii),(iii) and (v), the null hypothesis of cointegration vectors r = 0 and r ≤ 1 are both rejected which specify 2 cointegrating vectors of 2 variables (r=n). These results suggests that these set of variables are stationary at level I(0). Hence, Johansen maximum likelihood technique is not applicable to determine their long-run relationship. Furthermore, the big negative normalized estimated coefficients β are relatively nonsense.

There are two possible errors can effect our estimation. Firstly, the low power of unit root test could not fully reflect series characteristic for inflation, I(0) instead of I(1). In table 1a, we can see that some t-statistic tests of inflation rates reject the null hypothesis of a unit root at 10% critical level. Secondly, the selected lag length of Johansen procedure can manipulate the result. We observed big suggested differences in lag interval from SIC compared to other criterions. The Johansen cointegration test with lag interval based on AIC indicate no cointergrated relation except for the case of *T_bill* at 10% critical value (see Appendix 1).We proceed to use ARDL bounds testing approach to investigate the long-run cointegrating relationship between inflations and nominal interest rates. The next table 3 presents the joint F test statistic $H_0 : \varphi = \delta = 0$ from OLS regression Eq (15)

As reported in table 2a.ii for Vietnam, there is strong evidence for long-term relationship between inflation and intern bank 3 months interest rate. The calculated coefficient of linear relationship between inter bank 3 month and inflation is relatively illogical $\theta_1 = 14.002$. The coefficient suggests that 1% change in 3 months intern bank rate would lead to 14% change in inflation. The ARDL model suggests inconclusive relationship for deposit rate and 3 month inflation at 90% confident interval which partially supports the EG cointegration outcome. The findings of three applied techniques show little evidence for long term relationship between nominal interest rate and inflation in Vietnam.

In the OLS examination of long run Fisher effect for China, all the coefficients of the variables are significantly negative. The striking results suggest that there is a negative relationship between nominal interest rate and inflation. Therefore, the Fisher effect does not hold for China.

The EG cointegration test results do not support the long run relationship of 1 year Government bond and 5 year Government bond with their corresponding future inflation in India. The p-value of the ADF unit root test of the regression residual of *G-Bond_1y* and *inflation_12m* is equal to 0.093. The p-value of the ADF unit root test of the regression residual of *G-Bond_5y* and *inflation_60m* is equal to 0.91. Johansen and ARDL bound tests also suggest the same outcome in table 2c.i and 2c.ii. We also find a significant negative relationship of *deposit_6m* and inflation which contradicts with Fisher theory. The ARDL test confirms the long-run Fisher effect for 3months Treasury Bill and inflation in India for the period of 1997M1 – 2010M12.

In the case of Pakistan, Johansen and ARDL bound tests both show a negative long run relationship between nominal interest rates and inflation in table 2d.i and 2.dii. Interestingly, the short term structure of nominal interest is likely to cointegrated with the future inflation. In general, the coefficients clearly do not support the presence of long run Fisher effect in Pakistan.

In table 2e.i, the Johansen cointegration test suggests a long term relationship between 2 months deposit rate and 2 months future inflation. The normalized estimated coefficient β is approximately equal to 0.45. As reported in table 2e.ii, the joint F-test confirms testing the long run relationship between four nominal interest rates including *Lending rate, Deposit 2m, Deposit 3m and Interbank 12m* with the corresponding future inflation. However, the coefficients β are found to be positive and much less smaller than ideal value of 1 where nominal interest rates adjust one-for-one with movements in inflation. Overall, a weak long run Fisher effect exists in Argentina for the period 1997-2010.

In the case of Brazil, we perform EG cointegration test for Bank Certificate Deposit and 3 months inflation since they are I(1) . The p-value of the ADF unit root test based on (SIC) of the regression residual is equal to 0.002. Hence, the null hypothesis of unit root is rejected and the two series are cointegrated. For other sets, the series exhibit mixed orders of intergration I(1) and I(0). All three cointegration tests confirm the long run relationship between Bank Certificated Deposit and 3 months inflation. In the case of Treasury bill, the Johansen test results reject the null hypothesis of cointegrating vectors r = 0 and r ≤ 1, then indentify 2 cointegrating vectors of 2 variables (r=n). It means that TBill and 1 month inflation are not cointegrated in the long run which is consistent with the ARDL bound result. Furthermore, the normalized estimated coefficients β approximately are much less than 1.The ARDL bounds tests gives an inconclusive result for the long run relationship between 2 months deposit and inflation. Two coefficients β are found to be negative which clearly opposes with the Fisher theory. We find little evidence for the presence of weak Fisher effect in Brazil.

5.3 Short run Fisher effect

Short run relationship of Fisher effect is estimated by the OLS regression EQ (17).
All underlying variables of the regression are stationary at the level (see Apendix 2).

Table 3: Short run Fisher effect estimation

Country	Variable	Coefficient		$H_0: \alpha = 0, \beta = 1$ F-statistic
		α	β	
Vietnam	Δ Inflation_3m, Δ Deposit_3m 1997M2 - 2010M9	-0.02 (0.930)	-1.25* (0.005)	12.97 (0.000)
	Δ Inflation_3m, Δ Intern_3m 1999M2 – 2010M10	0.03 (0.935)	-0.57 (0.219)	5.74 (0.004)
	Δ Inflation_1m, Δ Intern_1m 1999M2 - 2010M11	-0.04 (0.962)	0.32 (0.775)	0.18 (0.837)
	Δ Inflation_12m, Δ T-bill 1997M3 - 2010M1	0.19 (0.013)	-0.07 (0.625)	38.60 (0.000)
	Δ Inflation_6m, Δ Lending rate 1997M2 – 2010M7	-0.06 (0.764)	-0.86* (0.002)	22.36 (0.000)

* indicates the coefficient is significant at 5%

Table 3 (cont'd): Short run Fisher effect estimation

Country	Variable	Coefficient		$H_{0:}\ \alpha = 0, \beta = 1$ F-statistic
		α	β	
China	∆ Inflation_3m, ∆ TBond 1997M1 - 2010M12	0.02 (0.945)	-0.2 (0.737)	2.74 (0.067)
	∆ Inflation_1y, ∆ Lending 1997M1 – 2010M12	0.01 (0.882)	-0.80 (0.061)	9.27 (0.000)
	∆ Inflation_24m, ∆ Deposit_2y 1997M1 - 2019M12	0.03 (0.476)	-0.26 (0.086)	35.96 (0.000)
	∆ Inflation_36m, ∆ Deposit_3y 1997M5 - 2008M12	-0.02 (0.25)	0.03 (0.61)	95.60 (0.000)
	∆ Inflation_60m, ∆ InterB_5y 1997M5 – 2006M12	0.01 (0.312)	-0.003 (0.946)	218.51 (0.000)
India	∆ Inflation_1m, ∆ Deposit_1m 1997M1 – 2010M12	0.04 (0.970)	-2.58 (0.086)	2.88 (0.059)
	∆ Inflation_3m, ∆ T-Bill 3m 1997M1 – 2010M12	0.03 (0.933)	-0.004 (0.996)	0.82 (0.449)
	∆ Inflation_1m, Deposit 6m 1997M1 – 2010M12	0.01 (0.951)	-0.21 (0.685)	2.67 (0.071)
	∆ Inflation_12m, ∆ G-bond 1y 1997M1 - 2010M12	0.01 (0.85)	-0.18 (0.17)	39.25 (0.000)
	∆ Inflation_60m, ∆ G-Bond 5y 1997M1 – 2006M12	0.02 (0.219)	-0.05 (0.277)	214.91 (0.000)
Pakistan	∆ Inflation_1m, ∆ KIBOR_1m 2002M9 - 2010M12	0.10 (0.94)	-2.21 (0.30)	1.11 (0.333)
	∆ Inflation_3m, ∆ KIBOR_3m 2002M9 - 2010M12	0.13 (0.74)	-1.76 (0.025)	6.34 (0.003)
	∆ Inflation_6m, ∆ REPO_6m 1997M1 - 2010M3	0.03 (0.856)	-0.36 (0.09)	20.15 (0.000)
	∆ Inflation_6m, ∆ T-Bill_6m 1997M1 - 2010M12	-0.02 (0.144)	-0.49 (0.039)	20.03 (0.000)
	∆ Inflation_12m, ∆ KIBOR_12m 2004M5 – 2006M12	0.11 (0.350)	-0.83 (0.002)	25.73 (0.000)

* indicates the coefficient is significant at 5%

Table 3 (cont'd): Short run Fisher effect estimation

Country	Variable	Coefficient		H_0: $\alpha = 0, \beta = 1$ F-statistic
		α	β	
Argentina	Δ Inflation_1m, Δ Lending rate 1997M1 - 2010M12	-0.03 (0.939)	0.26* (0.008)	29.51 (0.000)
	Δ Inflation_2m, Δ Deposit_2m 1997M1 – 2010M12	0.03 (0.952)	-0.10 (0.286)	63.80 (0.000)
	Δ Inflation_3m, Δ Deposit_3m 1997M1 - 2010M12	0.05 (0.863)	-0.44* (0.000)	152.37 (0.000)
	Δ Inflation_6m, Δ InterB_6m 1997M5 - 2010M12	0.005 (0.819)	0.10* (0.004)	389.02 (0.000)
	Δ Inflation_12m, Δ InterB_12m 1997M5 – 2010M12	0.05 (0.681)	0.02 (0.267)	1190.72 (0.000)
Brazil	Δ Inflation_1m, Δ TBill 1997M1 - 2010M12	-0.15 (0.631)	0.17 (0.195)	20.39 (0.000)
	Δ Inflation_2m, Δ Deposit_2m 1997M1 – 2010M12	-0.003 (0.987)	0.14 (0.115)	48.50 (0.000)
	Δ Inflation_12m, Δ BCD 2000M5 – 2010M12	0.02 (0.931)	0.04 (0.838)	10.42 (0.000)

* indicates the coefficient is significant at 5%

The results from estimating the regression equation Eq (17) for six countries are found in Table 3. Short run Fisher effect does not hold for 4 countries including China, India, Pakistan and Brazil since all their coefficients β are not significantly different to zero. In the case of Vietnam, all statistically significant coefficients β are found negative in Vietnam which contradicts with positive correlation between nominal interest rate and inflation. We find two significant positive short-run changed coefficients β of Lending rate and 6 months interbank interest. The joint F tests for the null hypothesis of a strong form Fisher effect are greater than critical value for all cases which clearly reject a one for one movement between nominal interest rate and expected inflation rate in Argentina. We only find evidence of a weak short run relationship between nominal interest rate and inflation in Argentina. Our findings also suggest that the short-run Fisher relationship is less likely to hold compared too long- run relationship.

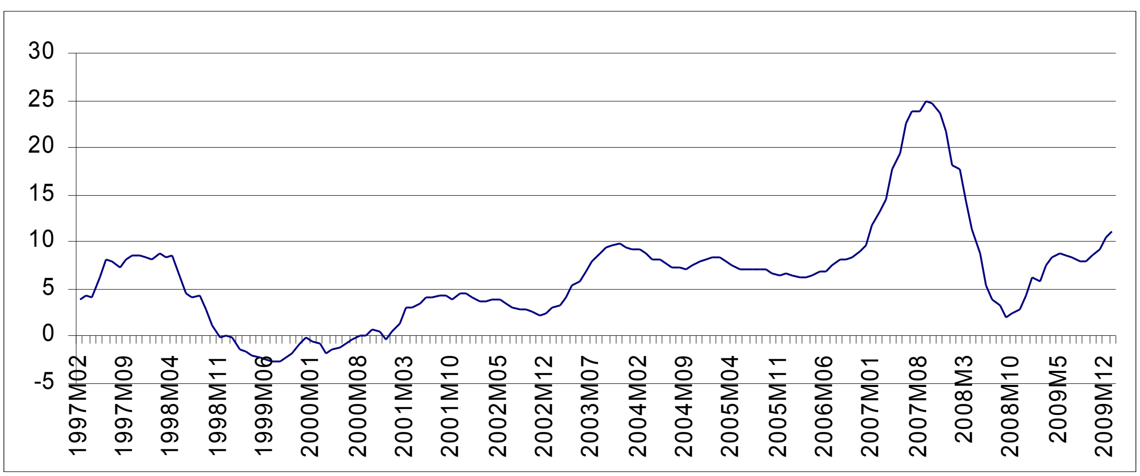
6. International Fisher effect and the case of Vietnam

6. 1Vietnam monetary policy

Vietnam has been under the process of transition toward market economy in the late 1980s. After experiencing the period of hyper-inflation and stagnation in 1987-1989, the country has been one of the fastest growing economies in the world. Before 1989, most of commercial banking sector was owned by the state. The main function of The State Bank of Vietnam (SBV) was to allocate financial resources to other working sectors in the economy. Vietnam government has implemented finical sector reform which transformed mono-bank system into two tier banking system in 1900. As the result, the SBV has official turned into a central bank which implements targeted monetary policy and supervise commercial banks and other financial institutions. The entry of commercial banking system has gradually become more open for domestic joint stock financial institutions and foreign owned bank. The Vietnam government has been reducing the state ownership to 51% in five SOBs by 2011. However, the six state-owned commercial banks (SOB) still accounted for 73% of total credits market in 2004 (Camen 2005). The Vietnam government has been reducing the state ownership to 51% in five SOBs by 2011. The increasing presence of foreign banks has put many small domestic joint stock banks under pressure. The November 2006 Law stated that all commercial banks which could not meet the minimum capital level of 3trilion VND would be forced to merge.

Many reports suggested that Vietnam government policies were keen to boost economic growth rather than macroeconomic stability. Vietnam has implemented the US-VND exchange rate targeting policy. The main advantage of exchange rate target is to facilitate the country export by reducing the exchange rate uncertainty. Figure 1 depicts the highly volatile inflation rate for the period 1997-2010. Inflations rates tend to be very high in the period December- February due to high consumer expenditure around Vietnamese New Year. Interestingly, the country price level underwent deflation in 2000. Due to the high oil price and global recession, the inflation even increased sharply to 20% in 2008. Since 2004, the SBV has always been struggled to meet their inflation targets. Long term deposits and transactions are likely to use US dollar because of the high inflation. The share of foreign currency deposit decreased from 41% in 2000 to 30% in 2004 (Camen 2005). Therefore, Vietnam could be considered as a dollarized economy with a dual monetary standard. The SBV has imposed the basic interest rate for lending and deposit rates every month. The nominal interest rates of commercial banks must not exceed 150% a basic interest rate. Recently, the policy has proved to be inefficient. In 2010, the year to year inflation rate was 21% but the nominal average nominal interest rate throughout the year was below the level. As the consequence, many commercial banks attempted to provide additional lucrative rewards beside the maximum deposit rate in order to attract more customers. It is interesting to note that the deposit rate of US dollar in Vietnam market hit the peak around 6% at the beginning of 2011 which was 4 times higher than US market itself before the new regulation has implemented to limit the USD deposit rate to maximum of 2%. Furthermore, government report indicated that the amount of Vietnamese overseas remittance increased by 11% and hit the record of 9 billion US dollar in 2011. The movements of capital inflow and nominal interest rate might suggest the presence of arbitrage opportunity.

Figure 1: Vietnam annually inflation



6.2 International Fisher effect for Vietnam

To verify the IFE theory, we will test the hypothesis of the interest rate differential has positive significant influence on change in the exchange rate. A positive relationship β indicates if there is either an increase in domestic nominal interest/inflation rate or a decrease in foreign interest/inflation rate, the domestic currency will depreciate in the exchange market. Under the assumption of efficient market hypothesis, the exchange rate should adjust quickly to the new information of nominal interest rates between two countries to prevent any potential arbitrage investment opportunities. The joint hypothesis testing of $\alpha = 0$ and $\beta = 1$ represents the condition of IFE in the perfect market.

We will first employ unit root test to decide the appropriate techniques to investigate the relationship in Eq (18). ADF and KPSS unit root tests will be applied to interest differential and the change in exchange rate.

Table 4: Unit root tests for International Fisher effect

	Level		First difference		Order of integration	Lags length
	ADF t-statistic (p-value)	KPSS LM-stat	ADF t-statistic (p-value)	KPSS LM-stat		
$\Delta Y_{US/VN}$	-10.41416 (0.0000)	0.255582	x	X	I(0)	1
$\Delta Y_{UK/VN}$	-17.00947 (0.0000)	0.194180	x	X	I(0)	0
$\Delta Y_{EU/VN}$	-9.453064 (0.0000)	0.079873	x	X	I(0)	0
ΔX_{US-VN}	0.402258 (0.9825)	0.823552	-9.140219 (0.0000)	0.331699	I(1)	12
ΔX_{UK-VN}	-0.235297 (0.9298)	0.707436	-7.015497 (0.0000)	0.309934	I(1)	12
ΔX_{EU-VN}	4.041534 (1.000)	0.787574	-2.918395 (0.0456)	0.724338	I(1) *	1

* Order of integration for ΔX_{EU-VN} is determined by PP unit root test since KPPS and ADF show conflicting results. PP suggests the series is stationary at first difference. The 5% critical value of KPSS is equal to 0.463.

The reported results in Table 5 indicate interest differential and change in nominal interest series do not have the same of order of integration. Therefore, we will perform Johansen cointegration technique and ARDL bounds testing approach to analyze the International Fisher effect in table 6 and 7 below. Both methods suggest strong evidence for co-integrated relationship between the nominal interest rate differential and the change in interest rate. All the coefficients β are significantly different to zero. Furthermore, the calculated coefficients β from two methods are quite similar except for the case of UK/VN. A negative coefficient of $\beta_{US/VN}$ means the interest differential between Vietnam and the US has a negative significant effect on the exchange rate. This result is contrary with the IFE and interest parity theory. However, negative coefficient could be caused by exchange rate targeting policy. We found positive coefficients of β for the two remaining cases. The interest differentials between Vietnam and the foreign currencies of UK and EU (France) have positive significant effect on their exchange rates. When the interest rate differential is higher, the domestic currency tends to depreciate against the foreign currency. The magnitudes of estimated coefficients β are relatively small compared to hypothetical value of 1. The results illustrates that 1% increase in nominal interest differential would lead to less than 0.05% changes in the exchange rate. Hence, the interest differential between any listed foreign currency and Vietnam has relatively weak effect on the future exchange rate. The last column of table 6 represents the F statistic tests for the null hypothesis of the presence of IFE. In general, we found no evidence supports the IFE for Vietnam

Table 5: Johansen maximum likelihood estimation for IFE

Variables Sample period	Lag length of VAR	H_0	λ_{trace} (p-value)	$\lambda_{\max}$ (p-value)	No. of co-vectors	Estimated coefficient β (Standard error)
$\Delta Y_{US/VN}, \Delta X_{US-VN}$ 1997M2 - 2010M8	11	r = 0	23.79 (0.002)	22.89 (0.002)	1	-0.02 (0.007)
		r ≤ 1	0.91 (0.341)	0.91 (0.341)		
$\Delta Y_{EU/VN}, \Delta X_{EU-VN}$ 1999M2 – 2010M10	2	r = 0	60.14 (0.000)	59.97 (0.000)	1	0.0001 (0.003)
		r ≤ 1	0.17 (0.680)	0.17 (0.680)		
$\Delta Y_{UK/VN}, \Delta X_{UK-VN}$ 1999M2 to 2010M11	10	r = 0	14.60* (0.068)	14.34 (0.048)	1	-0.001 (0.003)
		r ≤ 1	0.24 (0.623)	0.24 (0.623)		

* indicates null hypothesis is not rejected at 95% confident interval

Table 6: ARDL bounds testing approach for IFE

	ARDL bounds testing					H_0: $\alpha = 0$, $\beta = 1$
	Lag difference (p,q)	F-test statistic	Cointergration at 95% confident interval	Coefficient α	Coefficient β	
$\Delta Y_{US/VN}, \Delta X_{US-VN}$ 1997M2 - 2010M8	(1,12)	44.79	yes	0.03	-0.02	5612.88 (0.0000)
$\Delta Y_{EU/VN}, \Delta X_{EU-VN}$ 1999M2 - 2010M9	(0,1)	62.11	yes	-0.01	0.001	28220.49 (0.0000)
$\Delta Y_{UK/VN}, \Delta X_{UK-VN}$ 1997M2 - 2010M8	(0,12)	22.26	yes	0.003	0.001	694743.5 (0.0000)

7. Conclusion

In this research, we have examined the Fisher effect for six developing coutries in the period 1997-2010. First, we tested the long run and short run relationship between nominal interest rates and future inflation. Short-term interest rates can be affected by government policy due to financial crisis in 2008. After adjusting our sample, the financial crisis does not have significant effect on our conclusion.

Based on the results of Engle-Granger, Johansen and ARDL cointegration test, OLS regression, we find little evidence for the weak form of long run Fisher hypothesis in India and Vietnam. The evidence of long-run Fisher effect is much more profound in Argentina. The Fisher hypothesis does not hold for Brazil, China and Pakistan. Interestingly, our finding points out some significant negative relationship between nominal interest rates and inflation.

On the other hand, the short-run Fisher effect is less common in all six countries. We obtain little evidence of a weak short run Fisher effect in Argentina.

We also discuss the effect of interest rate differential on the change in exchange rate. We find a weak form of International Fisher effect Vietnam with two foreign currency UK pounds and France euro. However, the strong form of IFE is rejected for all cases. Nominal interest rates did not have much predictive power about future inflation. The result suggests that Vietnam financial market is far from efficient. Hence, macroeconomics stability is the key to help financial institution operate efficiently.

Further research can include the effect of marginal tax rate on income. Beyer, Haug and Dewad(2009) suggested that the rejection of long term Fisher effect might due to structural change in cointegrating vector. Identify break point or regime shift can produce more accurate result.

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Appendix 1: Johansen maximum likelihood estimation results (AIC)

Variables $y_t = \left[\pi_t^m \quad i_t^m\right]'$ Sample period	Lag length of VAR	H_0	λ_{trace} (p-value)	λ_{max} (p-value)	No. of co-vectors	Estimated coefficient β (Standard error)
Inflation_3m, Deposit_3m	11	r = 0	8.32	6.66	0	-1.78
1997M2 - 2010M9			(0.432)	(0.530)		(0.67)
		r ≤ 1	1.65	1.65		
			(0.198)	(0.198)		
Inflation_3m, Intern_3m	12	r = 0	9.34	9.46	0	-2.05
1999M2 – 2010M10			(0.335)	(0.554)		(0.92)
		r ≤ 1	2.88	2.88		
			(0.090)	(0.090)		
Inflation_1m, Intern_1m	12	r = 0	18.48	12.73	2	-4.43
1999M2 to 2010M11			(0.017)	(0.017)		(0.92)
		r ≤ 1	5.76	5.76		
			(0.016)	(0.016)		
Inflation_12m, T-bill **	8	r = 0	14.36	12.30	0	-9.62
1997M3 to 2010M1			(0.074)	(0.100)		(3.10)
		r ≤ 1	2.06	2.06		
			(0.151)	(0.151)		
Inflation_6m, Lending rate	9	r = 0	14.79	10.15	0	-2.51
1997M2 – 2010M7			(0.064)	(0.20)		(0.94)
		r ≤ 1	4.64	4.64		
			(0.031)	(0.031)		

** indicates cointegrated at 10% critical value. Lag length is based on Akaike info criterion

Appendix 2: Unit root tests short run change

	Unit root test at Level		Order of integration
	ADF t-statistic (p-value)	KPSS LM-stat	
∆ Deposit_3m	-6.36 (0.000)	0.08	I(0)
∆ Intern_3m	-6.02 (0.0000)	0.11	I(0)
∆ Intern_1m	-6.84 (0.000)	0.09	I(0)
∆ T-Bill	-4.861 (0.000)	0.07	I(0)
∆ Lending Rate	-6.74 (0.000)	0.06	I(0)
∆ Inflation_1m 1999M2 - 2010M11	-3.36 (0.014)	0.26	I(0)
∆ Inflation_3m 1997M2 - 2010M9	-5.45 (0.0000)	0.03	I(0)
∆ Inflation_3m 1999M2 – 2010M10	-3.22 (0.021)	0.05	I(0)
∆ Inflation_6m 1997M2 – 2010M7	-4.97 (0.0000)	0.02	I(0)
∆ Inflation_12m 1997M3 - 2010M1	-3.25 (0.019)	0.04	I(0)

The end!