University of Macedonia

Master of Economics

Field: Applied Economics and Finance

“ Testing the Fisher Effect in OECD countries : An empirical investigation.”

by

Karoumpa Argyro (15/09)

Supervisor : Fountas Stylianos

September 2010

Abstract

This paper tests the validity of the Fisher hypothesis for the OECD countries using a cointegration procedure developed by Gregory and Hansen (1996) that allows for the presence of a one-time endogenously determined structural break in the cointegrating vector. We are paying particular attention to the integration and cointegration properties of the variables, since meaningful Fisher effect tests critically depend on such properties. It is noteworthy that , contrary to the other empirical studies ,we test the hypothesis of stationarity of the variables using a relatively recent unit root test suggested by Ng-Perron (2001) due to its superiority to Augmented Dickey Fuller (ADF) and Phillips-Perron (P-P) tests. The results indicate that in the majority of the countries tested, their variables do not satisfy the integration properties and therefore we cannot proceed to cointegration techniques. For the rest of the countries where the integration properties are satisfied the cointegration procedure indicate that the full Fisher effect is present for only one country, Canada, while the partial Fisher effect holds for Belgium and Korea.

Contents

Ackhowledgements

I am greatly indepted to my supervisor Stelios Fountas for his helpful comments and advice. His criticisms and suggestions improved my empirical analysis tremendously. In addition , I want to thank everyone in my family for their love and support. I dedicate this work to them.

1 Introduction

The long-run Fisher hypothesis, formalized by Irving Fisher (1930) states that a permanent change in the rate of inflation will cause an equal change in the nominal interest rate so that the real interest rate is not affected by monetary shocks in the long run. For this long run relation to hold, the ex ante real interest rate, which is the difference between the nominal interest rate and the expected rate of inflation must be mean-reverting. If the nominal interest rate and the inflation rate are each integrated of order one, denoted I(1), then the two variables should cointegrate with a slope coefficient of unity so that the real interest rate is covariance stationary. The hypothesis has been studied extensively because it has important policy implications. If the hypothesis holds in the long run, the monetary policy will have no influence on the real interest rate, since, in this case, any change in the expected inflation will be an offset by a change in the nominal rate of interest, leaving the ex ante real interest rate unchanged.

Numerous empirical analyses have been done and various models have been proposed and tested (using data from both developed and developing countries) for the Fisher effect. Interestingly, the existence of the Fisher effect has been subject to debate. Its importance, however, is unarguable. From a macroeconomic perspective, the Fisher effect is the cornerstone of neutrality monetary models (i.e., money supply) and it is critical in explaining the movement of other economic fundamentals (i.e exchange rate). More importantly, because inflation is the fact of life in economies, and because of the difference between nominal and real interest rate, which affects all inter temporal savings and investment decisions in the economy, the understanting

of the Fisher link- the inflation, nominal and real interest rate- is the key to gaining knowledge about how each economy runs as a whole and how different economies interact.

A variety of empirical techniques have been used to test the Fisher hypothesis. Most recently, cointegration and error correction methodologies have been used to investigate whether the nominal interest rate and some measure of expected inflation cointegrate with a unitary coefficient which is evidence that the strong form Fisher hypothesis describes an empirically valid long-run equilibrium relation.

Recently, doubts have been raised whether these time series techniques are capable of robustness in the face of major structural changes in the economy (Gregory and Hansen 1996 , Zivot and Andrews 1992). Possible candidates for structural change that might affect empirical testing of the Fisher hypothesis are events such as a change in the monetary authority's operating procedures, or the implementation of major structural reforms such as financial deregulation. Events such as these might be expected to lead to parameter instability in the relation between the nominal interest rate and the expected rate of inflation, and thus call into doubt the findings of cointegration analysis based on a stable long-run equilibrium relation.

One response to the possibility of structural instability is to modify the cointegration procedure to make allowance for structural change. The problem with standard cointegration tests in the presence of breaks were carefully illustrated by Gregory and Hansen (1996).A break introduces spurious unit root behavior in the cointegrating relationship so that the hypothesis of no cointegration is difficult to reject.

The advantage of the testing approach developed by Gregory and Hansen (1996) is that it allows for an one-time endogenously determined structural break in the cointegrating vector.

In this paper, we provide an explanation for the apparent failure of Fisher hypothesis in earlier literature. We argue that the finding in the literature of no cointegration for Fisher hypothesis may be due to structural changes in the cointegrating vector. We use the Gregory and Hansen's cointegration test for the null hypothesis of no cointegration against the alternative hypothesis of cointegration in the presence of a structural break under both hypothesis. Their methodology allows us to determine whether the finding of no cointegration for the Fisher effect is due to structural change or not. We study the long-run Fisher effect for the OECD countries over the post-war period. Many empirical studies found no empirical evidence for the long-run Fisher effect. These findings would imply that money was not super-neutral and there was money illusion, if one assumes that money growth drives inflation, because real interest rates would be affected by inflation. We found for the most of the countries no support of cointegration due to the fact that their variables do not satisfy the integration properties. In cases where the integration properties are satisfied, we found evidence of a break in the cointegrating relationship. The full Fisher effect is present for Canada, while the partial Fisher effect holds for Belgium and Korea.

2 Literature review

Despite the relatively large number of theoretical models that assume the Fisher hypothesis holds , its validity is under investigation. The results are mixed, with different studies using different tests and different samples reaching different conclusions.

In theoretical framework, clearly, the Fisher relation holds in models (e.g. Sidrauski, 1967) in which the real interest rate is determined by a relation like the modified golden rule and therefore does not depend on monetary variables. Alternative explanations have been proposed by empirical researchers in an attempt to explain why Fisher's economically intuitive hypothesis has not held in its strictest form. Mundell (1963) and Tobin (1965) argue that nominal interest rates should adjust by less than one-for-one due to the impact of inflation on wealth and subsequently savings. Darby (1975) and Feldstein (1976) point out that the effects of tax would result in a more than one-for-one adjustment to expected inflation, while Shome, Smith and Pinkerton (1988) suggest a premium needs to be incorporated in nominal interest rates to account for covariance risk.

In empirical framework, Carmichael and Stebbing (1983) suggest a different relationship between inflation, nominal interest rates and real interest rates to that of Fisher (1930).Assuming money and financial assets to be substitutable, they hypothesize that nominal interest rates on financial assets can be considered constant over time and that the real rate of interest moves inversely with inflation. Using quarterly data for the United States (US) for the period 1953-1978 and for Australia for two periods; 1965-1981 and 1963-

1981, they are able to find evidence of the Inverted Fisher Effect. They argue that this is the reason for many empirical studies failing to find evidence for the Fisher effect in its strictest form. This so-called inverted Fisher effect, or Fisher paradox, has had little empirical support. Testing the same dataset as used by Carmichael and Stebbing (1983), Moazzami (1991) cannot find the same long run inverse relationship between the real rate of interest and expected inflation. Likewise both Choudhry (1997), using data from Belgium, France and Germany from 1955-1994, and Woodward (1992), studying monthly British data from 1982-1990, are unable to find evidence of the inverted Fisher effect.

Rose (1998) using Dickey-Fuller (DF) tests, investigates the order of integration of nominal interest rates and inflation for the US and 17 other OECD countries for a sample where the data run from 1957Q1, and finds that the inflation rates are non-stationary and nominal interest rates are stationary for these countries. He thus concludes that the real interest rates are non-stationary for these countries and thus there is no evidence for the Fisher effect .

Over the years , the long-run neutrality proposition have been investigated in a number of studies. King and Watson(1997) have contributed to the literature on testing the long-run neutrality by developing tests based on coefficient restrictions in bivariate vector autoregressive models (VAR models).They study quarterly data from US for the period 1949-1990.They show that meaningfull neutrality tests can only be constructed if both nominal and real variables satisfy certain non-stationarity conditions. In particular, they show that Fisher effect tests are possible if the inflation and interest rate series are

integrated of order one and do not cointegrate. In this study the Fisher Hypothesis is not confirmed. Similar study have been made by Koustas and Serletis (1999) paying explicit attention to the integration and cointegration properties, since meaningfull Fisher effect tests critically depend on the property of stationarity in the first differences of the variables and no-cointegration. Studying quarterly data for 11 OECD countries for a sample that begins between 1957 and 1972 and ends in 1995 are unable to find evidence of the Fisher effect. Furthermore, Engsted (1995) studying 13 OECD countries for the sample 1962-1993 rejected long-run neutrality of inflation with respect to real interest rates, using the framework of King and Watson(1997).Using the same framework, Rapach (2003) studying 14 OECD countries for a sample that begins between 1949 and 1965 and ends between 1994 and 1996 found no evidence of the Fisher Effect.

The emergence of the literature on unit roots and cointegration provided an important impulse to the empirical testing of the Fisher effect. Following the early work of Rose (1988), a number of further contributions aimed to test for the Fisher effect using cointegration techniques have subsequently appeared, with sometimes conflicting results.

Applying a simulation technique called Monte Carlo experiments, Mishkin (1992) takes the non-stationarity of inflation and nominal interest rates as a maintained hypothesis and applies the Engle-Granger (1987) methodology to test for common stochastic trends. He studies US monthly data for the period 1953-1990 and he finds that a strong Fisher effect occurs only during certain periods where inflation and interest rates have trends. He concludes that empirical evidence supports a long run Fisher effect, but not a short run

Fisher effect. Crowder and Hoffman (1996) identify the mechanism responsible for the non-stationary behavior of the system. Using a bivariate vector error correction model (VECM) for US quarterly post-war data for the period 1952-1991, they reveal a dynamic behavior of nominal interest rates and inflation. The VECM suggests a specific "causal" ordering where inflation has predictive content for the future course of the interest rates. Wesso (2000) examines the relationship between expected inflation and the nominal bond yield using South African data for the period January 1985 to February 1999. This corresponds with the South African Reserve Bank's monetary policy framework that targeted the growth in the broad money supply (M3). Using cointegration and error-correction modeling techniques, Wesso (2000) finds that long-term bond yields are largely driven by expected inflation in South Africa.

Weidmann (1997) considers a threshold co-integration (TC) model to test for the Fisher effect. Using German data for the period 1967-1996, he shows that the stochastic process governing the bivariate system of inflation and interest rates depends on the level of variables and can be designed as a TC model. The model explains the downward bias of the coefficient estimates, the country and sample sensitivity and supports the full Fisher effect. However, the TC model is based on the assumption that the Bundesbank is committed to price stability and will not allow inflation rates to become negative or persistently high. Therefore the findings help to explain the Fisher effect only in countries where Central Bank are independent and have already built a long track credibility record. Cristopoulos and León-Ledesma (2007) using a logistic and an exponential smooth transition autoregressive model (LSTR

and ESTR) for the long-run relationship study the validitity of the Fisher effect in US using quarterly data from 1960-2004.Their results are in favor of the Fisher effect.

Westerlund (2008) tested the Fisher effect in a cointegrated panel of 20 OECD countries with quarterly data from 1980 to 2004 and could not reject the Fisher hypothesis. Panels generally add power to cointegration tests due to the added cross sectional dimension, however, they also impose restrictions, particularly on the cross-sectional dependencies, that may not hold in the data. Wong and Wu (2003),using monthly data from G7 and eight Asian countries, test the hypothesis at short-run and long-run horizons using instrumental variable regressions. They find more support for the hypothesis at long-run horizons than at short-run horizons. Berument and Mehdi (2002), test whether the hypothesis holds for a sample of 26 developed and developing countries using an instrumental variable technique. They find that there is strong evidence of Fisher effect in 9 out of 12 developed countries and in 7 out of 14 developing countries.

A potential difficulty in assessing the time series properties of inflation and interest rates is the existence of structural breaks in the form of infrequent changes in the mean or the drift rate of the series due to distinct exogenous events (oil price shocks, shifts in monetary or fiscal policy regimes etc.).As Perron (1989) showed that a break in the deterministic time trend reduces dramatically the power of standard unit root tests because the possibility of a break changes the (asymptotic) distribution of the test. Therefore, if a series contains a structural break, standard unit root tests will fail to reject the null of a unit root when, in fact the, the null is false. Goldberg et al (2003) use

quarterly short-term interest rates and inflation over the period 1957q1-2000q2 for some OECD countries. They apply the Zivot-Andrews procedure, which allows for a one-time endogenously determined break and rejected the null of a unit root in the real interest rates for all the countries. Gregory et al. (1994) show that conventional cointegration tests are biased towards accepting the null of no-cointegration in the presence of structural breaks. Simirarly, recent studies, suggest that nominal interest rates, inflation rates and real interest rates may have experienced structural breaks. For instance, Garcia and Perron (1996) have retested Rose's (1988) data and found that the real interest rate is constant subject to regime shifts.

In the table below are presented in a summary the data and the results of different empirical studies that deal with the validity of the Fisher Hypothesis.

Authors	Sample	Countries	Measures of Interest Rates	Findings
Carmichael and Stebbing (1983)	1953Q1-1978Q4	US	3-month T-Bill rate	Evidence of IFH
	1965Q3-1981Q4	Australia	3-month Bank accepted/endorsed Commercial Bill rate	Evidence of IFH
	1963Q1-1981Q4		5-year industrial debendure yield	Evidence of IFH
Moazzammi (1991)	1953Q1-1978Q4	US	3-month T-Bill rate	No evidence of IFH
	1965Q2-1981Q4	Australia	3-month Bank accepted/endorsed Commercial Bill rate	No evidence of IFH
	1962Q1-1981Q4		5-year industrial debendure yield	No evidence of IFH
Choudhry (1997)	1958Q1-1994Q4	Belgium	T-bill/Long-term rate government bond	No evidence of IFH
	1958Q1-1994Q4	France	Call money rate/Long-term rate government bond	No evidence of IFH
	1955Q1-1994Q4	Germany	Call money rate/Long-term rate government bond	Evidence of long-run partial IFH
Woodward (1992)	1982M1-1990M8	Britain	indexed bonds yield	No evidence of IFH
Rose (1988)	1957Q1-end of each database	18 OECD countries	3-month T-Bill	No evidence of FH
King and Watson (1992)	1949Q1-1990Q4	US	3-month T-Bill	No evidence of FH
Koustas and Serletis (1999)	Between 1957Q1 and 1972Q1 and end in 1995Q2	11 OECD countries	short-term interest rates	No evidence of FH
Engsted (1995)	1962Q1-1993Q1	13 OECD countries	long-term government bonds yield	No evidence of FH
Rapach (2003)	Between 1949 and 1965 and end between 1994 and 1996	14 OECD countries	long-term government bonds yield	No evidence of FH
Mishkin (1992)	1953M1-1990M12	US	3-month T-Bill rate	Evidence of long-run FH
Crowder and Hoffman (1996)	1952Q1-1991Q4	US	3-month T-Bill rate	Evidence of FH
Wesso (2000)	1985M1-1999M2	South Africa	10-year government bonds yield	No evidence of FH
Weidmann(1997)	1967M1-1996M6	Germany	12-month T-Bill rate	Evidence of FH
Cristopoulos and León-Ledesma (2007)	1960Q1-2004Q4	US	3-month T-Bill rate	Evidence of FH
Westerlund (2008)	1980Q1-2004Q4	20 OECD countries	short-term interest rates	Evidence of FH
Wong and Wu(2003)	between 1958Q1 and 1966Q1 and end in 1999Q4	G7	stock indice yields	Evidence of long-run FH
	between 1970Q1 and1987Q2 and end in 1999Q4	8 Asian Countries		Evidence of long-run FH
Berument and Mehdi (2002)	between 1957M4 and 1981M6 and end in 1998	12 developed countries	T-bill and lending rates	Evidence of FH in 9 countries
	between 1957M5 and 1985M2 and end in 1998	14 developing countries		Evidence of FH in 7 countries

3 Research Methodology and Data

3.1 The Model

The Fisher hypothesis states that in the long run, inflation and nominal interest rates move together, meaning that real interest rates are stable in the long term:

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

where i_t is the nominal interest rate at time t; r_t is the real interest rate at time t; π_t^e is the expected inflation at time t. As usual we assume rational expectations so that expected inflation is equal to actual inflation (π_t).Hence we can write the Fisher equation as

$$i_t = r_t + \pi_t + e_t \tag{2}$$

When nominal interest rates and inflation behave each as an I(1) process in our samples, then they should be cointegrated with a slope coefficient a_1 .

$$i_t = a_0 + a_1\pi_t + e_t \tag{3}$$

where α_1 is the coefficient of interest. If a_1 is statistically equal to one then the strong form of the Fisher effect is implied or the full Fisher effect and thus a one-to-one relation between the nominal interest rate and inflation. The weak form of the Fisher effect or the partial Fisher effect implies that a_1 is positive and less than one. If a_1 is more than one then it is implied that the nominal interest rate is taxed and the Fisher hypothesis implies that there is a more than one-to-one relation between the nominal interest rate and inflation.(Darby 1975)

3.2 Econometric Methodology

Our analysis is testing for cointegration between the nominal interest rate and the inflation rate. The concept of cointegration involves the existence of a long-run relationship to which a system converges over time. Conversely, if the variables are not cointegrated, they tend to drift apart.

The basic problem of testing the above equation using standard tests and OLS models is that variables are not stationary that can lead to spurious results. However, if the nominal interest rate and the inflation rate are non-stationary and are integrated of order one then the Fisher hypothesis can be tested using cointegration techniques.

The most common procedure for testing cointegration is the Engle-Granger's residual based test that can easily be implemented. The procedure is based on testing the stationarity of the residuals in equation (3).If i_t and π_t are integrated of order one, the long-run Fisher hypothesis holds if the residuals are integrated of order zero. The full Fisher effect is present when $e_t \sim I(0)$ and a_1 is not significantly different from one that means that the real interest rate is mean-reverting. If, however, $e_t \sim I(0)$ and a_1 is significantly less than one , the weak form of Fisher effect is implied meaning that the real interest rate is non-stationary.

In the next step, we test for structural breaks in the cointegrating relationship by applying the Gregory and Hansen (1996) methodology. The procedure of Gregory and Hansen is an extension of the Engle-Granger's residual-based test by allowing for a one-time endogenously determined structural break in

the cointegrating vector. Gregory and Hansen consider the following three models

Model 1: level shift $i_t = a_0 + a_1\pi_t + a_2 D_{1t} + \varepsilon_t$ (4)

Model 2: level shift with trend $i_t = a_0 + a_1\pi_t + a_2 D_{1t} + a_3 t + \varepsilon_t$ (5)

Model 3: regime shift $i_t = a_0 + a_1\pi_t + a_2 D_{1t} + a_4\pi_t D_{1t} + \varepsilon_t$ (6)

The dummy variable $D_{1t} = 1$ if t× $[n\tau]$ and 0 otherwise, τ is an unknown parameter denoting the relative timing of the change point and [ ] denotes the integer part. In model 1, there is a change in the intercept, where a_0 represents the intercept before the shift and a_2 represents the change in the intercept at the time of shift. In model 2, a time trend is added to level shift model. Model 3 allows both the intercept and the slope to shift, where a_1 is the cointegrating slope coefficient before the regime shift and a_4 is the change in the slope coefficient following the regime shift. The null hypothesis in all three models is that ε_t is non-stationary or, equivalently , the nominal interest rate and inflation are not cointegrated. The alternative hypothesis is that ε_t is stationary with a one-time endogenously determined structural break in the cointegrating vector.

Gregory and Hansen (1996) propose three tests for the residuals ε_t.The tests are modifications of the tests proposed by Phillips(1987) and the ADF test, and they are designed to choose the break point that gives the least support for the null hypothesis of no-cointegration. In this paper, the ADF test is used for two reasons: first, the Gregory-Hansen procedure is commonly carried out by applying the ADF test on the residuals as Engle-Granger (1987) show that

the ADF test has the largest power. The ADF test proposed by Gregory and Hansen (1996), denoted by ADF(τ),is calculated by regressing $\Delta\hat{\varepsilon}_{t\tau\ \iota}$ in models 1,2 and 3 on $\hat{\varepsilon}_{\tau-1\tau}$ and $\Delta\hat{\varepsilon}_{t-1\tau},\ldots,\Delta\hat{\varepsilon}_{t-kt}$,where k is the number of lags on the first –differenced residuals. To determine this number of lags, Gregory and Hansen (1996) follow Perron and Vogelsang (1992) by setting a maximum of six lags and then testing backward until the last included lag is signinificant at the 5% level using critical normal values. The null hypothesis of no-cointegration and no structural break is rejected if the ADF(τ) statistic is greater, in absolute terms, than the corresponding critical value.

3.3 The Data

We retrieved quarterly data, for the consumer price index (CPI) and the short-term interest rates from the International Monetary Fund's online data base for the OECD countries: Australia, Austria, Belgium, Canada, Czech Republic, Denmark, Finland, France, Germany, Greece, Hungary, Iceland, Ireland, Italy, Japan, Korea, Luxembourg, Mexico, Netherlands, New Zealand, Norway, Poland, Portugal, Spain, Sweden, Switzerland, Turkey, United Kingdom and United States. We exclude Slovak Republic from our analysis due to inadequate data. The data have been obtained from the beginning of the base. The sample we used for each country is illustrated in tables 1 to 5.Inflation rates were calculated from the first differences of the natural logarithm of the CPI , multiplied by 400 to get annualized rates in percent. We picked for the short-term interest rate the 3-month Treasury bill where available for sufficient long spans (Canada, France, Sweden, United Kingdom and United States).Otherwise, the money rate was used (Australia, Austria, Belgium, Czech Republic, Denmark, Germany, Iceland, Italy, Japan, Korea,

the Netherlands, New Zealand, Norway, Poland, Portugal, Spain and Switzerland) or the deposit rate if the other two rates were unavailable(Finland, Hungary, Ireland, Luxembourg, Mexico, Portugal and Turkey).

4 Unit root tests

Non stationarity of the economic variables involved in the analysis leads to violation of the classical assumptions of standard regression methods and to spurious estimates. The possible endogeneity of regressors is a problem not well handled by OLS. The sample sizes available for data analysis are usually small, leading to small sample bias in estimates. Classical regression properties hold only for cases where variables are stationary (integrated of order 0).But by contrast most economic variables are integrated of order 1 or higher and hence do not satisfy these assumptions. In this case where all the variables are integrated of order one, error correction mechanisms or long run relationships may exist and therefore certain combinations of I(1) variables are likely to be I(0) and hence amenable to OLS estimation. Where this is so, the variables are said to be cointegrated and OLS estimates of such cointegrated variables may be superconsistent in the sense of collapsing to their true values more quickly than if the variables had been stationary. The first step is to determine the degree of integration of the individual series under investigation, thus the empirical analysis begins by examining this with univariate tests. Where a cointegrating relationship cannot be found, no long run relationship among the variables can be demonstrated and we have the case of spurious regression

A variable is determined to be I(1) if a unit root is found in levels and stationarity is found in first differences. Most of the variables in levels appear to possess trends of some sort, and so a linear trend is included in the unit root tests on levels.

4.1 Autoregressive Unit Root Tests

To illustrate the important statistical issues associated with autoregressive unit root tests, consider the simple AR(1) model

$y_t = \varphi y_{t-1} + \varepsilon_t$, where $\varepsilon_t \sim WN(0, \sigma^2)$

The hypotheses of interest are

H_0 : $\varphi=1$ (unit root in $\varphi(z) = 0$) $\Rightarrow y_t \sim I(1)$

H_1 : $|\varphi| < 1 \Rightarrow y_t \sim I(0)$

The test statistic is

$$t_{\varphi=1} = \frac{\hat{\varphi} - 1}{SE(\hat{\varphi})}$$

where $\hat{\varphi}$ is the least squares estimate and $SE(\hat{\varphi})$ is the usual standard error estimate. The test is a one-sided left tail test.

4.2 Dickey-Fuller Unit Root Test

The unit root tests described above are valid if the time series y_t is well characterized by an AR(1) with white noise errors. Many financial time series, however, have a more complicated dynamic structure than is captured by a simple AR(1) model. Said and Dickey (1984) augment the basic autoregressive unit root test to accommodate general ARMA(p, q) models with unknown orders and their test is referred to as the augmented Dickey-Fuller (ADF) test. The ADF test tests the null hypothesis that a time series y_t is I(1) against the alternative that it is I(0), assuming that the dynamics in the data have an ARMA structure. The ADF test is based on estimating the test regression:

$$y_t = \beta' D_t + \varphi y_{t-1} + \sum_{j=1}^{p} \psi_j \Delta y_{t-j} + \varepsilon_t \qquad (7)$$

where D_t is a vector of deterministic terms (constant, trend etc.). The p lagged difference terms, Δy_{t-j} , are used to approximate the ARMA structure of the errors, and the value of p is set so that the error ε_t is serially uncorrelated. The error term is also assumed to be homoskedastic. The specification of the deterministic terms depends on the assumed behavior of y_t under the alternative hypothesis of trend stationarity as described in the previous section. Under the null hypothesis, y_t is I(1) which implies that $\varphi = 1$. The ADF t-statistic and normalized bias statistic are based on the least squares estimates of (7) and are given by

$$ADF_t = t_{\varphi=1} = \frac{\hat{\varphi} - 1}{SE(\hat{\varphi})}$$

$$ADF_n = \frac{T - (\hat{\varphi} - 1)}{1 - \hat{\psi}_n - \ldots - \hat{\psi}_p}$$

An alternative formulation of the ADF test regression is

$$\Delta y_t = \beta' D_t + \pi y_{t-1} + \sum_{j=1}^{p} \psi_j \Delta y_{t-j} + \varepsilon_t \qquad (8)$$

where π = φ − 1. Under the null hypothesis, Δy_t is I(0) which implies that π = 0. The ADF t-statistic is then the usual t-statistic for testing π = 0 and the ADF normalized bias statistic is $T\hat{\pi}/(1 - \hat{\psi}_1 - \ldots - \hat{\psi}_p$.The test regression (8) is often used in practice because the ADF t-statistic is the usual t-statistic reported for testing the significance of the coefficient y_{t-1}.

4.3 Phillips-Perron Unit Root Test

Phillips and Perron (1988) developed a number of unit root tests that have become popular in the analysis of financial time series. The Phillips-Perron (PP) unit root tests differ from the ADF tests mainly in how they deal with serial correlation and heteroskedasticity in the errors. In particular, where the ADF tests use a parametric autoregression to approximate the ARMA structure of the errors in the test regression, the PP tests ignore any serial correlation in the test regression. The test regression for the PP tests is

$$\Delta y_t = \beta_0 D_t + \pi y_{t-1} + u_t$$

where u_t is I(0) and may be heteroskedastic. The PP tests correct for any serial correlation and heteroskedasticity in the errors u_t of the test regression by directly modifying the test statistics $T_{\pi=0}$ and $T\hat{\pi}$. These modified statistics, denoted Z_t and Z_π, are given by

$$z_t = \left(\frac{\hat{\sigma}^2}{\hat{\lambda}^2}\right)^{1/2} t_{\pi=0} - \frac{1}{2}\left(\frac{\hat{\lambda}^2 - \hat{\sigma}^2}{\hat{\lambda}^2}\right)\left(\frac{TSE(\hat{\pi})}{\hat{\sigma}^2}\right)$$

$$z_\pi = T\hat{\pi} - \frac{1}{2}\frac{T^2 SE(\hat{\pi})}{\hat{\sigma}^2}\left(\hat{\lambda}^2 - \hat{\sigma}^2\right)$$

The terms $\hat{\sigma}^2$ and $\hat{\lambda}^2$ are consistent estimates of the variance parameters

$$\sigma^2 = \lim_{T\to\infty} T^{-1} \sum_{t=1}^{T} E\left[u_t^2\right]$$

$$\lambda^2 = \lim_{T\to\infty} \sum_{t=1}^{T} E\left[T^{-1} s_T^2\right]$$

where $s_T^2 = \sum_1^T u_t$.

The sample variance of the least squares residual $\hat{u}_t$ is a consistent estimate of σ^2, and the Newey-West long-run variance estimate of u_t using $\hat{u}_t$ is a consistent estimate of λ^2.Under the null hypothesis that $\pi = 0$, the PP Z_t and Z_π statistics have the same asymptotic distributions as the ADF t-statistic and normalized bias statistics. One advantage of the PP tests over the ADF tests is that the PP tests are robust to general forms of heteroscedasticity in the

error term u_t. Another advantage is that the user does not have to specify a lag length for the test regression.

4.4 Some Problems with Unit Root Tests

The ADF and PP tests are asymptotically equivalent but may differ substantially in finite samples due to the different ways in which they correct for serial correlation in the test regression. In particular, Schwert (1989) finds that if Δy_t has an ARMA representation with a large and negative MA component, then the ADF and PP tests are severely size distorted (reject I(1) null much too often when it is true) and that the PP tests are more size distorted than the ADF tests. Recently, Perron and Ng (1996) have suggested useful modifications to the PP tests to mitigate this size distortion. In general, the ADF and PP tests have very low power against I(0) alternatives that are close to being I(1). That is, unit root tests cannot distinguish highly persistent stationary processes from non-stationary processes very well. Also, the power of unit root tests diminish as deterministic terms are added to the test regressions. That is, tests that include a constant and trend in the test regression have less power than tests that only include a constant in the test regression. For maximum power against very persistent alternatives the recent tests proposed by Elliot, Rothenberg and Stock (1996) and Ng and Perron (2001) should be used. These tests are described in the next section.

4.5 Elliot, Rothenberg and Stock Unit Root Test

Elliott, Rothenberg, and Stock (1996) introduce a potentially more powerful unit root test: the generalized least squares version of the ADF test. They find that powers of ADF tests are lower than those of the limiting power functions when deterministic components (mean or trend) are included in the data generating process.

Elliot et al. (1996), hereafter ERS, present an asymptotically efficient test of the unit root hypothesis based on the regression

$$\tilde{y}_t = \gamma \Delta \tilde{y}_{t-1} + \sum_{i=1}^{p} a_i \, \Delta \tilde{y}_{t-i} + u_t \qquad (9)$$

Where $\tilde{y}_t$ represents the quasi-differenced data obtained from the GLS regression.

$$\tilde{y}_t = y_t + x_t' \xi(\bar{c})$$

This class of test requires the choice of $\bar{c}$, the local-to-unity parameter, which following ERS is selected as

$$\bar{c} = \begin{cases} 1 - \dfrac{7}{T} \;\; , \;\; if \; x_t = \{1\} \\ 1 - \dfrac{13.5}{T}, if \; x_t = \{1, t\} \end{cases}$$

Since $\tilde{y}_t$ has already been detrended the elements of x_t need not be included in (9). The DF-GLS test for a unit root is based upon $H_0 : \gamma = 0$ in (9). The

results presented in ERS suggest that GLS local detrending yields substantial power gains over the standard ADF unit root test constructed using (8).

On the other hand, in the presence of a large negative moving average root in the residuals, the majority of unit root tests display significant size distortions resulting in over-rejections of the unit root null hypothesis (Schwert, 1989,Perron and Ng, 1996). In constructing the ADF and ERS tests it is necessary to select *p* the autoregressive truncation lag.

4.6 Ng-Perron Unit Root Test

The ADF and PP unit root tests are known (from MC simulations) to suffer potentially severe finite sample power and size problems. Firstly, the ADF and PP tests are known to have low power against the alternative hypothesis that the series is stationary (or TS) with a large autoregressive root (DeJong, et al, 1992.) Secondly, the ADF and PP tests are known to have severe size distortion (in the direction of over-rejecting the null) when the series has a large negative moving average root.

Ng and Perron (Econometrica, 2001), building on some of their own work (Perron and Ng, 1996) and work by Elliott, Rothenberg, and Stock (Econometrica, 1996), new tests to deal with both of these problems. Their tests, in contrast to many of the other “new” unit root tests that have been developed over the years, seems to have caught on as a preferred alternative to the traditional ADF and PP tests. The family of NP tests (which includes among others, modified DF and PP test statistics) share the following

features. First, the time series is de-meaned or detrended by applying a GLS estimator. This step turns out to improve the power of the tests when there is a large AR root and reduces size distortions when there is a large negative MA root in the differenced series. The second feature of the NP tests is a modified lag selection (or truncation selection) criteria. It turns out that the standard lag selection procedures used in specifying the ADF regression (or for calculating the long run variance for the PP statistic) tend to underfit, i.e., choose too small a lag length, when there is a large negative MA root. This creates additional size distortion in unit root tests. The NP modified lag selection criteria accounts for this tendency.

Ng and Perron (2001) use the GLS detrending procedure of ERS to create efficient versions of the modified PP tests of Perron and Ng (1996). These efficient modified PP tests do not exhibit the severe size distortions of the PP tests for errors with large negative MA or AR roots, and they can have substantially higher power than the PP tests especially when φ is close to unity.

Using the GLS detrended data $\tilde{y}_t$, the efficient modified PP tests are defined as

$$MZ_\alpha = (T^{-1}\tilde{y}_t^2 - f_0)/2\kappa$$

$$MSB = \left(\frac{\kappa}{f_0}\right)^{1/2}$$

$$MZ_t = MZ_a \times MSB$$

$$MPT = \begin{cases} (\bar{c}^2\kappa - \bar{c}T^{-1}\tilde{y}_t^2)/f_0 \ , & if\ x_t = \{1\} \\ (\bar{c}^2\kappa + (1-\bar{c})T^{-1}\tilde{y}_t^2)/f_0 \ , & if\ x_t = \{1, t\} \end{cases}$$

Where $\kappa = \sum_{t=2}^{T} y_{t-1}^2 / T^2$ and f_0 is an estimate of the residual spectral density at the zero frequency. The statistics MZ_α and MZ_t are efficient versions of the PP Z_α and Z_t tests that have much smaller size distortions in the presence of negative moving average errors. Again the choice of the autoregressive truncation lag, *p*, is critical for correct calculation of f_0. Here *p* is chosen using the Modified Information Criteria (MIC(*p*)) of Ng and Perron (2001) as $p = p_{MIC} = \arg\min_p MIC(p)$ where

$$\tau_\tau(p) = \left(\hat{\sigma}_p^2\right)^{-1} \hat{\gamma}^2 \sum_{t=p_{max}+1}^{T} \tilde{y}_{t-1}^2$$

$$\hat{\sigma}_p^2 = (T - p_{max})^{-1} \sum_{t=p_{max}+1}^{T} \hat{u}_t^2$$

4.7 Empirical results

We employ the Ng-Perron (2001) unit root test strategy to test the variables for stationarity. The choice of Ng-Perron procedure is propelled by its superiority to both Augmented Dickey Fuller (ADF) and Phillips-Perron (P-P) tests and, furthermore, is built on the work by Elliott, Rothenberg, and Stock (1996) that yields substantial power gains over the standard ADF unit root test.

The results of the unit root tests are indicated in the tables 1,2,3 and 4. The null hypothesis is that there is a unit root in the series. We reject the null hypothesis of the existence of a unit root when the test statistic is less than the corresponding critical value. We base our inferences on the 5% level of significance. We cannot reject the null hypothesis of a unit root for all cases when only a constant is considered in the test regression. This suggests that variables are non-stationary in level form. However, the results in table 2 show that whether MZa, MZt, MSB, MPT is used as test statistic interest rates and inflation rates exhibit random walk behavior after first difference for eight countries; Australia, Belgium, Canada, Czech Republic, Greece, Hungary, Mexico and Turkey. Once a deterministic trend is added , a unit root is no longer rejected for the interest rate of Japan. The results are shown in table 3.However, the results of Ng-Perron unit root test in first differences when a deterministic trend is considered in table 4 show that a unit root is rejected for Australia, Belgium, Canada, Mexico and Turkey as earlier and furthermore for Italy and Korea. We, therefore, taking into consideration both

cases we examine ten countries; Australia, Belgium, Canada, Czech Republic, Greece, Hungary, Italy, Korea, Mexico and Turkey.

Thus, for these countries cointegration tests can be used to see if there is a linear combination between the two variables that is stationary since the two variables for each country are non-stationary and integrated of order one. For the other countries, their variables either the interest rate or the inflation rate do not satisfy the criterion of stationarity in their first differences and therefore these countries are excluded from the cointegration tests in the next section.

Furthermore, the ADF unit root tests for nominal interest rates and inflation are provided in table 5.The inferences we make from this test are quite different from Ng-Perron. We base our inferences on the 5% level of significance. When only a constant term is considered the ADF test implies that the nominal interest rates and inflation are integrated of order one for all countries except from Czech Republic, Germany, Iceland, Mexico, Netherlands, New Zealand and Poland. When a deterministic trend is added the variables are integrated of order one for Austria, Belgium, Canada, Finland, Greece, Hungary, Iceland, Ireland, Italy and Korea.

The two procedures provide different inferences about the stationarity of the variables. As we have analyzed earlier the Ng-Perron procedure is more robust method for testing the presence of unit roots in the variables and therefore we rely on this method .

5 Cointegration tests

Macroeconomic time series are typically non-stationary, as established by Nelson and Plosser (1982). When traditional regression analysis is used on two non-stationary time series, a spurious regression may result (Granger and Newbold 1974). The non-stationary nature of the majority of macroeconomic time series has prompted the development of various non-stationary time series analysis techniques, the most prominent being cointegration analysis. This concept of cointegration, introduced by Granger (1981) and later extended by Engle and Granger (1987), is built on the premise that the linear combination of two non-stationary series results in a stationary series. Cointegration can be defined simply as the long-term, or equilibrium, relationship between two series. This makes cointegration an ideal analysis technique to validate the Fisher hypothesis: by ascertaining the existence of a long-term unit proportionate relationship between nominal interest rates and expected inflation, cointegration analysis can thereby establish if nominal interest rates are cointergrated with expected inflation. The cointegration method by Engle-Granger (1987) has become the most cited cointegration technique used in Fisherian literature, and is used in this study.

5.1 The Engle-Granger test

The Engle-Granger (1987) (EG) test is the most commonly employed (single equation) approach to the analysis of cointegration in the econometrics literature. Given two variables of interest $\{y_t, x_t\}$, the first stage of this two-step procedure involves the estimation of the following static cointegrating regression:

$$y_t = d_t + \beta x_t + \varepsilon_t \qquad (10)$$

where d_t denotes a deterministic term which may be either an intercept ($d_t = \alpha$) or an intercept and linear trend ($d_t = \alpha + \beta t$). In the second step, potential cointegration between $\{y_t, x_t\}$ is examined via analysis of the order of integration of the residuals $\{\hat{\varepsilon}_t\}$ from (10) using a Dickey-Fuller (1979) test as below:

$$\Delta\hat{\varepsilon}_t = (\rho - 1)\hat{\varepsilon}_{t-1} + \nu_t \quad (11)$$

The null hypothesis of no cointegration is examined via the t-like statistic for $(\rho - 1)$. In empirical analysis, equation (11) is augmented as necessary via the inclusion of lagged values of the dependent variable.

5.2 The Gregory-Hansen Test

The Monte Carlo analysis of Gregory et al. (1994) shows that the power of the Engle-Granger test is substantially reduced when applied to cointegrated series which experience a change or break in their cointegrating relationship. In response to this, Gregory and Hansen (1996) extend the Engle-Granger test to explicitly allow for breaks in either the intercept or the intercept and cointegrating coefficient at an unknown time. The above equation (10) for the first stage of the Engle-Granger testing procedure is therefore revised as below to provide the following three models:

Model C: Level shift $\qquad y_t = \mu_0 + \mu_1\varphi_t + ax_t + u_t$

Model C/T: Level shift with trend $y_t = \mu_0 + \mu_1\varphi_t + \beta t + ax_t + u_t$

Model C/S: Regime shift $y_t = \mu_0 + \mu_1\varphi_t + a_1x_t + a_2\varphi_t x_t + u_t$

Each of the above models therefore permits structural change via the dummy variable φ_t which is defined as:

$$\varphi_t = \begin{cases} 1 & if \;\; t > T \\ 0 & otherwise \end{cases}$$

where τ denotes the point in the sample at which a break occurs. To determine τ , Gregory and Hansen (1996) suggest the use of a grid search procedure, with all values in the central 70% of the sample being considered. For each value of τ , the above models are estimated with the resulting residuals $\{\hat{u}_t\}$ saved and employed in the following Dickey-Fuller testing equation:

$$\Delta\hat{u}_t = (\rho - 1)\hat{u}_t + \nu_t$$

which may be augmented as required by the addition of lagged values of $\Delta\hat{u}_t$ The resulting test statistic for each model is then given as the minimum value obtained for the t-ratio of (ρ-1).

5.3 Empirical results

Gregory and Hansen (1996) suggest two steps to test for cointegration. First, to apply the conventional Engle-Granger procedure and perform the ADF test on the residuals. If the null of no-cointegration is not rejected, the Gregory and Hansen procedure is applied. If the null of no-cointegration is rejected, we can conclude that the cointegrating vector has experienced a structural break. If both procedures reject the null of no-cointegration , no inference can be made that the cointegrating vector has experienced a structural break.

The results of applying the Engle-Granger procedure on the residuals in equation (2) are provided in tables 6 and 7.When nominal interest rate is the dependent variable the null hypothesis of no-cointegration is rejected for Belgium at the 10% significance level. The null hypothesis of a_1=1 is rejected for every significance level which provides evidence in favor of the partial Fisher effect for Belgium, that is a less than one-to-one relation between nominal interest rates and inflation in this country. When the nominal inflation is the dependent variable they are rejections of the null hypothesis of no-cointegration for Greece at the 10% level, for Hungary at the 5% level and for Czech Republic, Mexico and Turkey at the 1% level.

However, it is possible that these results of no-cointegration or weak evidence at the 10% level are due to structural breaks in the cointegrating vector between the nominal interest rate and inflation in these countries that caused their relationship to shift. To account for this possibility the Gregory-Hansen procedure is applied. The results are illustrated in Table 8.

Actually, a visual inspection of the Engle-Granger regression residuals in figures 1 and 2 indicate the possibility of a structural break in the cointegrating

vector in some countries. Particularly, in figure 1, the residuals of Australia, Belgium, Canada, Greece and Italy seem to exhibit a non-random behavior and therefore there is no evidence of cointegration in these countries. From the other hand, the residuals of Czech Pepublic, Hungary, Korea, Mexico and Turkey seem to imply a long-run relationship between nominal interest rates and inflation with a possibility of the existence of structural breaks in the cointegrating vectors. In figure 2, it is more obvious the possibility of the existence of cointegration between nominal interest rates and inflation with the presence of structural breaks in each country as the series of residuals seem to exhibit a random behavior for long spans.

The results in the table 8 indicate evidence of cointegration at the 10% level for Australia when inflation is the dependent variable. The estimated break date is 1982Q3 which is not related to any major event.

The estimated break date for Belgium is 1977Q1 according to model 1 and 1978Q4 according to model 2 at the 5% level when nominal interest rate is the dependent variable. The breakpoints may be related то the 1973-79 oil price shocks and the resultant shifts in international demand that sent the economy into a period of prolonged recession.

Canada has experienced a structural break in 1980Q2 according to model 2 at the 1% level when nominal interest rate is the dependent variable and in 1983Q1 according to model 2 at the 10% level when inflation rate is the dependent variable. These estimated break dates are possibly due to the major economic event of the 1980s; the 1981-1982 recession. The 1985 Canada Year Book describes the main causes of the deepest and longest recession of the Canadian economy since the Second World War.

There is evidence of cointegration for Czech Republic according to the three models when nominal inflation is the dependent variable with estimated breakpoints 1996Q4, 2005Q2 and 1997Q3 respectively. The 1996-1997 break dates are probably related to the political and financial crisis that occurred in 1997 and shattered the Czech Republic's image as one of the most stable and prosperous of post-Communist states. From the other hand we cannot explain the 2005Q2 break date since there is not any particular major event that took place in the country this time.

The estimated break date (Model 2) for Greece is 1973Q3 at the 10% level and is possibly due to the 1973-79 energy crisis and the shift in the political regime from dictatorship to democracy.

There is also evidence of cointegration for Hungary when interest rate is the dependent variable as indicated by Model 2 at the 5% level with 1988Q3 as the estimated break date and by Model 2 when inflation is the dependent variable at the 1% level with an estimated break in 1990Q1 at the 1% level. The estimated break dates in 1988Q3 and 1990Q1 may be attributed to the large falls in output that occurred in the early 1990s.

There is strong evidence of cointegration at the 1% level for Italy when inflation is the dependent variable. The estimated break dates are 1982Q2 and 1982Q4.These dates are not related to any major event.

The estimated breakpoints for Korea according to the three models whether the nominal interest rate or the nominal inflation is the dependent variable are 1997Q3 and 1984Q2 at the 1% significance level and 1990Q2,1981Q4 and 1983Q1 at the 5% level. The 1997Q3 breakpoint is possibly due to 1997/98 Asian crisis and 1984Q2 is probably related to the domestic financial

liberalization that was started in Korea in 1984 whereas the 1990Q2 breakpoint is possibly related to the current account liberalization that started in this country in 1988. The estimated break dates 1981Q4 and 1983Q1 for Korea seem to be a little bit puzzling to explain since this time does not seem to be related to any particular major event.

There is evidence of cointegration at the 1% level for Mexico when nominal inflation is the dependent variable with breakpoints 1995Q4, 1984Q1 and 1983Q2.The 1995 Q4 break point is possibly attributed to the currency crisis that occurred in Mexico during 1994-1995 while the 1983Q2 and 1984Q1 break dates are possibly due to the international debt crisis of 1982.

Lastly there is strong evidence of cointegration for Turkey as indicated by all models with estimated break dates 2002Q4, 1994Q2, 1993Q4, 2003Q3,1995Q4 and 1991Q4.The 2002Q4 and 2003Q3 breakpoints are around the 1998 currency crisis that took place in Turkey whereas 1993-1995 break dates are probably due to the sharp rise in real interest rates due to the unexpected rise in profits and dividends in 1993-1995.

6 Dynamic OLS

With regard to the estimation of cointegrating regression models, it is well known that the ordinary least squares (OLS) estimator contains the second-order bias, including the endogeneity bias and the non-centrality bias, when the I(1) regressors are endogenous and the regression errors are serially correlated. Parameter estimates can be biased in small samples as well as in the presence of dynamic effects, and this bias varies inversely with the size of the sample and the calculated R^2. Also, when the number of regressors

exceeds two there can be more than one cointegrating relationship or vector and it is difficult to give economic meaning to this finding. Then there is the problem caused by the likely endogeneity of the regressors, which would prevent OLS estimating the true values of the parameters. These difficulties associated with the OLS approach have led to the development of alternative procedures which are proposed in the literature.

One method of extracting purely the long-run coefficients is by the Johansen and Juselius procedure which is based on a maximum-likelihood approach. Another more recent and more robust method, (particularly in small samples) proposed by Stock and Watson (1993), which also corrects for possible simultaneity bias among the regressors, involves estimation of long-run equilibria via dynamic OLS (DOLS). Stock and Watson (1993) suggest a parametric approach for estimating long-run equilibria in systems which may involve variables integrated of different orders but still cointegrated. The potential of simultaneity bias and small-sample bias among the regressors is dealt with by the inclusion of lagged and led values of the change in the regressors. Their method improves on OLS by coping with small sample and dynamic sources of bias. The Johansen method, being a full information technique, is exposed to the problem that parameter estimates in one equation are affected by any mispecification in other equations. The Stock-Watson method is, by contrast, a robust single equation approach which corrects for regressor endogeneity by the inclusion of leads and lags of first differences of the regressors, and for serially correlated errors by a GLS procedure. In addition it has the same asymptotic optimality properties as the Johansen distribution. Furthermore, based on Monte Carlo evidence, Stock

and Watson (1993) show that DOLS is more favourable, particularly in small samples, compared to a number of alternative estimators of long-run parameters, including those proposed by Engle and Granger (1987), Johansen (1988) and Phillips and Hansen (1990).

In estimating the parameters, the DOLS procedure is adopted where the nominal interest rate is regressed on the level of the inflation plus the lags and leads of its first differences. The models that are used are the following

No break model:

$$i_t = a_0 + a_1\pi_t + \sum_{-k}^{k} \psi_j \Delta\pi_{t-j} + \varepsilon_t$$

Model 1: Level shift

$$i_t = a_0 + a_1\pi_t + a_2 D_{1t} + \sum_{-k}^{k} \psi_j \Delta\pi_{t-j} + \varepsilon_t$$

Model 2: Level shift with trend

$$i_t = a_0 + a_1\pi_t + a_2 D_{1t} + \alpha_3 t + \sum_{-k}^{k} \psi_j \Delta\pi_{t-j} + \varepsilon_t$$

Model 3: Regime shift

$$i_t = a_0 + a_1\pi_t + a_2 D_{1t} + a_4\pi_t D_{1t} + \sum_{-k}^{k} \psi_j \Delta\pi_{t-j} + \varepsilon_t$$

In our analysis up to four leads and lags are used and we proceed with the exclusion of the last lead and the last lag according to the General-to-specific method when the joint hypothesis of Wald test indicates that the last included

lead and lag are not significant at the 10% level. Furthermore, the standard errors are computed by Newey-West correction for heteroscedasticity and autocorrelation. The full Fisher effect holds if a_1 is not statistically different from one. In addition, the full Fisher effect with a regime shift implies that $a_1=1$ and $a_2 \neq 0$, and the partial Fisher effect with a regime shift holds if $a_1<0$ and $a_2 \neq 0$.

6.1 Empirical results

The DOLS results are indicated in Table 9.The strong form of the Fisher Hypothesis holds for only Canada since a_1 is statistically equal to one and a_2 is statistically different from zero. The partial form of the Fisher effect holds for Belgium (according to all models) and for Korea (model 1,model 3) since in these cases a_1 is positive and less than one and a_2 is statistically different from zero.

According to the three models, there is no cointegration evidence for Turkey since a_2 is not statistically different from zero. Also according to the model 2, there is no cointegration evidence for Hungary and Korea. Consequently, for these countries the Fisher Hypothesis do not hold.

7 Summary and Conclusions

This study investigates whether the Fisher hypothesis holds in the OECD countries. The hypothesis involves a long-run one-to-one relation between the nominal interest rate and expected inflation. It has important policy implications since, if it holds, the monetary policy will have no influence on real interest rate as, in this case, any change in expected inflation will be offset by a change in the nominal interest rate, leaving the real interest rate unchanged.

The paper tests whether there are structural breaks in cointegrating vectors of each country's Fisher equation and also identifies the dates of breakpoints since endogenous structural breaks exist. The Ng-Perron strategy was used for integration analysis while the Engle-Granger procedure and Gregory-Hansen approach were used for cointegration analysis. The empirical findings from the integration tests suggest that model variables became I(0) after first differences only in the minority of the countries (Australia, Belgium, Canada, Czech Republic, Greece, Hungary, Italy, Korea, Mexico and Turkey) - in contrast to the previous empirical studies that use the traditional ADF and PP methods. The Engle-Granger procedure indicates weak evidence of the Fisher hypothesis at the 10% level for Belgium. Using the Gregory- Hansen tests reveal that there exist structural breaks in the cointegrating vectors of each country's Fisher equation and particularly there is weak evidence at the 10% level for Australia, Canada and Czech Republic also some strong evidence at the 5% for Belgium, Czech Republic, Greece, Hungary and Korea and strong evidence at the 1% level for Canada, Italy, Korea, Mexico and Turkey. It is obvious that some countries (Canada, Czech Republic, Hungary

and Korea) indicate evidence of structural breaks in different significant levels and in different break dates using different models with either the nominal interest rate or the inflation rate as a dependent variable. The DOLS procedure indicates that the break that is significant for Belgium is 1978Q4, for Canada is 1980Q2 and for Korea is 1997Q3.Furthermore , the DOLS procedure indicate that the full Fisher effect is present for only Canada, while the partial Fisher effect holds for Belgium and Korea .Thus the monetary policy tool is efficient only in the cases where the partial Fisher effect holds and particularly for Belgium and Korea.

Following the seminal research of Perron (1989), it has long been recognized that the presence of structural change can substantially reduce the power of unit root tests. In response to this finding, a number of tests have been proposed which allowed for an endogenously determined single structural break.(Zivot and Andrews, 1992).More recently, Lee and Strazicich (2003) have proposed an LM unit root test which allows for two breaks. Given that we have not used unit root tests that allow for structural breaks, this can be done in future.

Table 1: Ng-Perron Unit Root Tests in Levels when only constant is considered

	Sample period		MZa	MZt	MSB	MPT
Australia	1969Q3:2010Q1	i_t	-2.436	-1.092	0.448	9.991
		p_t	-6.540***	-1.808***	0.276	3.749***
Austria	1968Q1:1998Q4	i_t	-5.294	-1.605	0.303	4.690
		p_t	-1.858	-0.935	0.504	12.828
Belgium	1968Q1:1998Q4	i_t	-5.101	-1.597	0.313	4.804
		p_t	-3.537	-1.265	0.358	6.929
Canada	1968Q1:2010Q1	i_t	-4.967	-1.390	0.280	5.382
		p_t	-6.230***	-1.729***	0.278	4.054***
Czech Republic	1993Q1:2010Q1	i_t	-0.977	-0.486	0.498	15.696
		p_t	-1.355	-0.823	0.607	18.070
Denmark	1972Q1:2010Q1	i_t	-2.504	-1.028	0.411	9.343
		p_t	-1.311	-0.756	0.576	17.168
Finland	1981Q1:2005Q3	i_t	0.694	0.558	0.804	44.792
		p_t	0.586	0.716	1.221	92.092
France	1969Q4:2010Q1	i_t	-1.827	-0.648	0.355	9.922
		p_t	-2.153	-0.941	0.437	10.594
Germany	1991Q1:2010Q1	i_t	-0.787	-0.347	0.441	14.324
		p_t	-2.944	-1.151	0.391	8.180
Greece	1968Q1:2005Q4	i_t	-1.380	-0.770	0.558	16.230
		p_t	-2.256	-1.048	0.464	10.755
Hungary	1976Q1:2010Q1	i_t	-3.722	-1.364	0.366	6.582
		p_t	-2.750	-1.158	0.421	8.860
Iceland	1987Q1:2009Q4	i_t	-0.765	-0.480	0.627	22.000
		p_t	-2.824	-1.180	0.418	8.650
Ireland	1968Q1:2006Q2	i_t	-5.781***	-1.623***	0.281	4.480
		p_t	-4.160	-1.430	0.344	5.908
Italy	1971Q1:2010Q1	i_t	-1.848	-0.849	0.459	11.873
		p_t	-0.913	-0.676	0.740	26.832
Japan	1968Q1:2010Q1	i_t	-1.185	-0.546	0.461	13.861
		p_t	-3.720	-1.331	0.358	6.603
Korea	1976Q4:2009Q4	i_t	-1.332	-0.541	0.406	11.970
		p_t	-4.648	-1.413	0.304	5.508
Luxembourg	1980Q1:1999Q1	i_t	0.388	0.208	0.536	22.505
		p_t	-5.616	-1.651***	0.294	16.170
Mexico	1977Q1:2010Q1	i_t	-4.439	-1.472	0.332	5.553
		p_t	-2.746	-1.171	0.427	8.921
Netherlands	1968Q1:2010Q1	i_t	-7.954***	-1.960***	0.246***	3.210***
		p_t	-3.844	-1.383	0.360	6.376
New Zealand	1985Q1:2010Q1	i_t	0.711	1.066	1.500	139.290
		p_t	0.394	0.635	1.612	148.395
Norway	1971Q4:2009Q3	i_t	-1.446	-0.779	0.538	15.324
		p_t	-2.463	-0.911	0.370	8.971

Poland	1991Q1:2010Q1	i_t	0.902	2.230	2.473	382.816
		p_t	0.310	0.345	1.114	72.981
Portugal	1976Q1:2000Q1	i_t	-4.765	-1.459	0.306	5.333
		p_t	-0.864	-0.638	0.738	27.100
Spain	1974Q1:2010Q1	i_t	-2.705	-1.026	0.379	8.588
		p_t	0.777	0.781	1.005	67.319
Sweden	1968Q1:2009Q1	i_t	-3.176	-1.090	0.343	7.524
		p_t	-1.356	-0.683	0.503	14.594
Switzerland	1975Q4:2010Q1	i_t	-6.301***	-1.681***	0.267***	4.202***
		p_t	-4.678	-1.460	0.312	5.390
Turkey	1989Q1:2010Q1	i_t	-0.537	-0.284	0.528	18.488
		p_t	-2.304	-1.073	0.466	10.634
United Kingdom	1968Q1:2010Q1	i_t	-2.576	-0.860	0.334	8.424
		p_t	-3.481	-1.285	0.369	7.035
United States	1974Q1:2010Q1	i_t	-4.104	-1.176	0.286	6.290
		p_t	-4.897	-1.477	0.302	5.214

***,**,* denotes significance at the 10%,5% and 1% significance levels.The critical values for the Ng-Perron test for 10%, 5% and 1% are -5.7,-8.1 and -13.8 respectively for MZa, -1.62, -1.98 and -2.58 respectively for MZt , 0.27, 0.23, 0.174 respectively for MSB and 4.45, 3.17 and 1.78 respectively for MPT.

Table 2: Ng-Perron Tests in First Differences when only constant is considered

	Sample Period		MZa	MZt	MSB	MPT
Australia	1969Q3:2010Q1	i_t	-41.571*	-4.555*	0.110*	0.600*
		p_t	-58.824*	-5.421*	0.092*	0.423*
Austria	1968Q1:1998Q4	i_t	-33.590*	-4.098*	0.122*	0.729*
		p_t	-0.064	-0.178	2.782	379.291
Belgium	1968Q1:1998Q4	i_t	-60.239*	-5.488*	0.091*	0.408*
		p_t	-44.274*	-4.698*	0.106*	0.573*
Canada	1968Q1:2010Q1	i_t	-63.146*	-5.619*	0.089*	0.389*
		p_t	-77.580*	-6.225*	0.080*	0.323*
Czech Republic	1993Q1:2010Q1	i_t	-14.485*	-2.691*	0.186**	1.692*
		p_t	-21.323*	-3.206*	0.150*	1.356*
Denmark	1972Q1:2010Q1	i_t	-0.371	-0.394	1.061	56.352
		p_t	0.589	2.419	4.109	972.544
Finland	1981Q1:2005Q3	i_t	-36.381*	-4.265*	0.117*	0.674*
		p_t	0.176	0.634	3.596	672.146
France	1969Q4:2010Q1	i_t	-2.671	-1.133	0.424	9.087
		p_t	-263.611*	-11.468*	0.044*	0.108*
Germany	1991Q1:2010Q1	i_t	-21.235*	-3.258*	0.153*	1.154*
		p_t	-0.053	-0.114	2.163	232.829
Greece	1968Q1:2005Q4	i_t	-11.864**	-2.434**	0.205**	2.071**
		p_t	-41.114*	-4.521*	0.110*	0.633*
Hungary	1976Q1:2010Q1	i_t	-40.957*	-4.520*	0.110*	0.614*
		p_t	-22.910*	-3.337*	0.146*	1.232*
Iceland	1987Q1:2009Q4	i_t	-56.718*	-5.324*	0.094*	0.434*
		p_t	-0.134	-0.221	1.651	135.432
Ireland	1968Q1:2006Q2	i_t	-69.396*	-5.890*	0.085*	0.353*
		p_t	0.786	1.903	2.422	358.351
Italy	1971Q1:2010Q1	i_t	-131.478*	-8.106*	0.062*	0.190*
		p_t	-0.827	-0.595	0.720	26.368
Japan	1968Q1:2010Q1	i_t	-46.810*	-4.836*	0.103*	0.528*
		p_t	-0.174	-0.288	1.654	134.418
Korea	1976Q4:2009Q4	i_t	-961.509*	-21.926*	0.023*	0.026*
		p_t	-0.397	-0.445	1.119	61.375
Luxembourg	1980Q1:1999Q1	i_t	-2.336	-1.080	0.462	10.486
		p_t	-0.221	-0.204	0.922	163.790
Mexico	1977Q1:2010Q1	i_t	-65.133*	-5.707*	0.088*	0.376*
		p_t	-44.145*	-4.698*	0.106*	0.555*
Netherlands	1968Q1:2010Q1	i_t	-60.799*	-5.513*	0.091*	0.405*
		p_t	0.331	0.776	2.347	302.078
New Zealand	1985Q1:2010Q1	i_t	0.490	3.200	6.531	2386.870
		p_t	0.722	2.816	3.900	904.458
Norway	1971Q4:2009Q3	i_t	-0.356	-0.392	1.104	60.601
		p_t	-0.058	-0.113	1.932	186.822

Poland	1991Q1:2010Q1	i_t	-33.268*	-4.078*	0.123*	0.739*
		p_t	0.533	1.699	3.184	578.294
Portugal	1976Q1:2000Q1	i_t	-11.167**	-2.362**	0.211**	2.199
		p_t	0.431	1.772	4.106	933.927
Spain	1974Q1:2010Q1	i_t	0.127	0.235	1.845	179.856
		p_t	-0.029	-0.119	4.042	800.822
Sweden	1968Q1:2009Q1	i_t	-22.407*	-3.294*	0.147*	1.277*
		p_t	1.005	2.052	2.041	269.324
Switzerland	1975Q4:2010Q1	i_t	-0.667	-0.471	0.705	26.713
		p_t	-0.041	-0.120	2.912	417.554
Turkey	1989Q1:2010Q1	i_t	-289.879*	-12.037*	0.042*	0.086*
		p_t	-29.590*	-3.846*	0.130*	0.828*
United Kingdom	1968Q1:2010Q1	i_t	-77.011*	-6.204*	0.081*	0.321*
		p_t	-1.529	-0.830	0.543	15.116
United States	1974Q1:2010Q1	i_t	-9.629**	-2.153**	0.224	2.710
		p_t	-4.502	-1.500	0.333	5.443

***,**,* denotes significance at the 10%,5% and 1% significance levels. The critical values for the Ng-Perron test for 10%, 5% and 1% are -5.7,-8.1 and -13.8 respectively for MZa, -1.62, -1.98 and -2.58 respectively for MZt , 0.27, 0.23, 0.174 respectively for MSB and 4.45, 3.17 and 1.78 respectively for MPT.

Table 3: Ng-Perron Unit Root Tests in Levels when constant and time trend are considered

	Sample period		MZa	MZt	MSB	MPT
Australia	1969Q3:2010Q1	i_t	-2.919	-1.135	0.389	29.235
		p_t	-7.813	-1.976	0.253	11.663
Austria	1968Q1:1998Q4	i_t	-7.569	-1.854	0.245	12.255
		p_t	-2.201	-0.996	0.453	38.747
Belgium	1968Q1:1998Q4	i_t	-8.000	-1.903	0.238	11.675
		p_t	-4.701	-1.456	0.310	18.893
Canada	1968Q1:2010Q1	i_t	-6.864	-1.777	0.259	13.366
		p_t	-8.522	-2.064	0.242	10.694
Czech Republic	1993Q1:2010Q1	i_t	-7.924	-1.976	0.249	11.539
		p_t	-3.078	-1.155	0.375	27.560
Denmark	1972Q1:2010Q1	i_t	-3.608	-1.279	0.355	24.247
		p_t	-0.649	-0.422	0.650	83.470
Finland	1981Q1:2005Q3	i_t	-10.160	-2.238	0.220	9.043
		p_t	-2.108	-0.973	0.462	40.283
France	1969Q4:2010Q1	i_t	-10.197	-2.219	0.218	9.126
		p_t	-8.875	-2.093	0.236	10.319
Germany	1991Q1:2010Q1	i_t	-2.352	-1.070	0.455	38.129
		p_t	-1.304	-0.798	0.612	68.596
Greece	1968Q1:2005Q4	i_t	-1.175	-0.550	0.468	46.964
		p_t	-3.196	-1.254	0.392	28.294
Hungary	1976Q1:2010Q1	i_t	-6.003	-1.658	0.276	15.115
		p_t	-2.830	-1.186	0.419	32.093
Iceland	1987Q1:2009Q4	i_t	-3.888	-1.369	0.352	23.107
		p_t	-7.124	-1.826	0.256	12.890
Ireland	1968Q1:2006Q2	i_t	-8.649	-2.066	0.239	10.585
		p_t	-6.144	-1.712	0.279	14.811
Italy	1971Q1:2010Q1	i_t	-2.563	-1.047	0.408	32.493
		p_t	-1.316	-0.791	0.601	66.542
Japan	1968Q1:2010Q1	i_t	-30.276*	-3.890*	0.128*	3.015*
		p_t	-5.011	-1.575	0.314	18.147
Korea	1976Q4:2009Q4	i_t	-14.083	-2.650***	0.188	6.493***
		p_t	-9.332	-2.160	0.231	9.766
Luxembourg	1980Q1:1999Q1	i_t	-4.807	-1.505	0.313	18.687
		p_t	-5.616	-1.651	0.294	16.170
Mexico	1977Q1:2010Q1	i_t	-5.274	-1.608	0.305	17.221
		p_t	-3.428	-1.305	0.381	26.512
Netherlands	1968Q1:2010Q1	i_t	-11.060	-2.302	0.208	8.498
		p_t	-3.732	-1.364	0.365	24.386
New Zealand	1985Q1:2010Q1	i_t	-1.016	-0.637	0.627	74.562
		p_t	-0.346	-0.298	0.862	142.495

Norway	1971Q4:2009Q3	i_t	-1.739	-0.795	0.457	42.052
		p_t	-14.069	-2.652***	0.189	6.479***
Poland	1991Q1:2010Q1	i_t	-1.435	-0.707	0.493	48.605
		p_t	-0.956	-0.493	0.516	55.540
Portugal	1976Q1:2000Q1	i_t	-7.132	-1.829	0.256	12.874
		p_t	-1.257	-0.734	0.584	64.171
Spain	1974Q1:2010Q1	i_t	-4.671	-1.485	0.318	19.223
		p_t	-2.451	-1.099	0.448	36.842
Sweden	1968Q1:2009Q1	i_t	-3.935	-1.300	0.330	21.927
		p_t	-1.594	-0.677	0.424	38.966
Switzerland	1975Q4:2010Q1	i_t	-7.134	-1.864	0.261	12.816
		p_t	-6.460	-1.784	0.276	14.111
Turkey	1989Q1:2010Q1	i_t	-3.115	-1.212	0.389	28.397
		p_t	-3.050	-1.227	0.402	29.686
United Kingdom	1968Q1:2010Q1	i_t	-4.634	-1.384	0.299	18.752
		p_t	-9.101	-2.121	0.233	10.064
United States	1974Q1:2010Q1	i_t	-15.637***	-2.764	0.177***	6.025***
		p_t	-12.662	-2.516	0.199	7.197

***,**,* denotes significance at the 10%,5% and 1% significance levels. The critical values for the Ng-Perron test for 10%, 5% and 1% are -14.2,-17.3 and -23.8 respectively for MZa, -2.62, -2.91 and -3.42 respectively for MZt , 0.185, 0.168, 0.143 respectively for MSB and 6.67, 5.48 and 4.030 respectively for MPT.

Table 4. Ng-Perron Unit Root tests in first differences when constant and time trend are considered

	Sample period		MZa	MZt	MSB	MPT
Australia	1969Q3:2010Q1	i_t	-76.779*	-6.194*	0.081*	1.197*
		p_t	-59.392*	-5.447*	0.092*	1.544*
Austria	1968Q1:1998Q4	i_t	-34.388*	-4.147*	0.121*	2.650*
		p_t	-0.058	-0.128	2.199	887.665
Belgium	1968Q1:1998Q4	i_t	-60.928*	-5.519*	0.091*	1.496*
		p_t	-52.202*	-5.105*	0.098*	1.765*
Canada	1968Q1:2010Q1	i_t	-87.581*	-6.617*	0.076*	1.042*
		p_t	-78.169*	-6.249*	0.080*	1.175*
Czech Republic	1993Q1:2010Q1	i_t	-3.126	-1.250	0.400	29.153
		p_t	0.098	0.128	1.307	330.628
Denmark	1972Q1:2010Q1	i_t	-1.184	-0.767	0.648	76.603
		p_t	0.073	0.242	3.290	2010.280
Finland	1981Q1:2005Q3	i_t	-37.302*	-4.317*	0.116*	2.451*
		p_t	-0.223	-0.210	0.942	170.264
France	1969Q4:2010Q1	i_t	-9.285	-2.152	0.232	9.824
		p_t	-14.840***	-2.724***	0.184***	6.142***
Germany	1991Q1:2010Q1	i_t	-21.848**	-3.303**	0.151**	4.185**
		p_t	-3.357	-1.270	0.378	26.650
Greece	1968Q1:2005Q4	i_t	-17.332**	-2.937**	0.169***	5.303**
		p_t	0.128	0.441	3.444	2220.380
Hungary	1976Q1:2010Q1	i_t	-41.532*	-4.554*	0.110*	2.208*
		p_t	-1.185	-0.761	0.642	75.471
Iceland	1987Q1:2009Q4	i_t	-49.423*	-4.971*	0.101*	1.846*
		p_t	-1.834	-0.927	0.505	47.463
Ireland	1968Q1:2006Q2	i_t	-69.964*	-5.914*	0.085*	1.303*
		p_t	-0.517	-0.506	0.978	174.347
Italy	1971Q1:2010Q1	i_t	-64.520*	-5.680*	0.088*	1.413*
		p_t	-33.873*	-4.096*	0.121*	2.801*
Japan	1968Q1:2010Q1	i_t	-51.269*	-5.062*	0.099*	1.782*
		p_t	-0.239	-0.336	1.402	359.121
Korea	1976Q4:2009Q4	i_t	-62.121*	-5.573*	0.090*	1.470*
		p_t	-56.161*	-5.298*	0.094*	1.630*
Luxembourg	1980Q1:1999Q1	i_t	-2.784	-1.172	0.421	32.504
		p_t	-0.221	-0.204	0.922	163.790
Mexico	1977Q1:2010Q1	i_t	-65.208*	-5.710*	0.088*	1.398*
		p_t	-44.151*	-4.698*	0.106*	2.065*
Netherlands	1968Q1:2010Q1	i_t	-61.325*	-5.537*	0.090*	1.486*
		p_t	-0.028	-0.060	2.138	843.437
New Zealand	1985Q1:2010Q1	i_t	-1.522	-0.793	0.521	51.973
		p_t	-0.310	-0.394	1.268	293.102
Norway	1971Q4:2009Q3	i_t	-0.703	-0.555	0.789	115.335

		p_t	0.181	0.288	1.589	487.999
Poland	1991Q1:2010Q1	i_t	-29.530*	-3.842*	0.130*	3.087*
		p_t	-0.094	-0.185	1.976	715.335
Portugal	1976Q1:2000Q1	i_t	-12.238	-2.460	0.201	7.520
		p_t	-0.380	-0.421	1.108	224.769
Spain	1974Q1:2010Q1	i_t	-3.503	-1.304	0.372	25.684
		p_t	-0.222	-0.222	1.001	190.674
Sweden	1968Q1:2009Q1	i_t	-10.525	-2.200	0.209	9.118
		p_t	0.094	0.166	1.776	597.810
Switzerland	1975Q4:2010Q1	i_t	-2.713	-1.145	0.422	32.951
		p_t	0.228	0.352	1.542	463.571
Turkey	1989Q1:2010Q1	i_t	-26.674*	-3.652*	0.137*	3.418*
		p_t	-29.570*	-3.845*	0.130*	3.082*
United Kingdom	1968Q1:2010Q1	i_t	-4.634	-1.384	0.299	18.752
		p_t	-1.818	-0.948	0.521	49.696
United States	1974Q1:2010Q1	i_t	-183.950*	-9.589*	0.052*	0.499*
		p_t	-2.352	-1.070	0.455	38.129

***,**,* denotes significance at the 10%,5% and 1% significance levels.The critical values for the Ng-Perron test for 10%, 5% and 1% are -14.2,-17.3 and -23.8 respectively for MZa, -2.62, -2.91 and -3.42 respectively for MZt , 0.185, 0.168, 0.143 respectively for MSB and 6.67, 5.48 and 4.030 respectively for MPT.

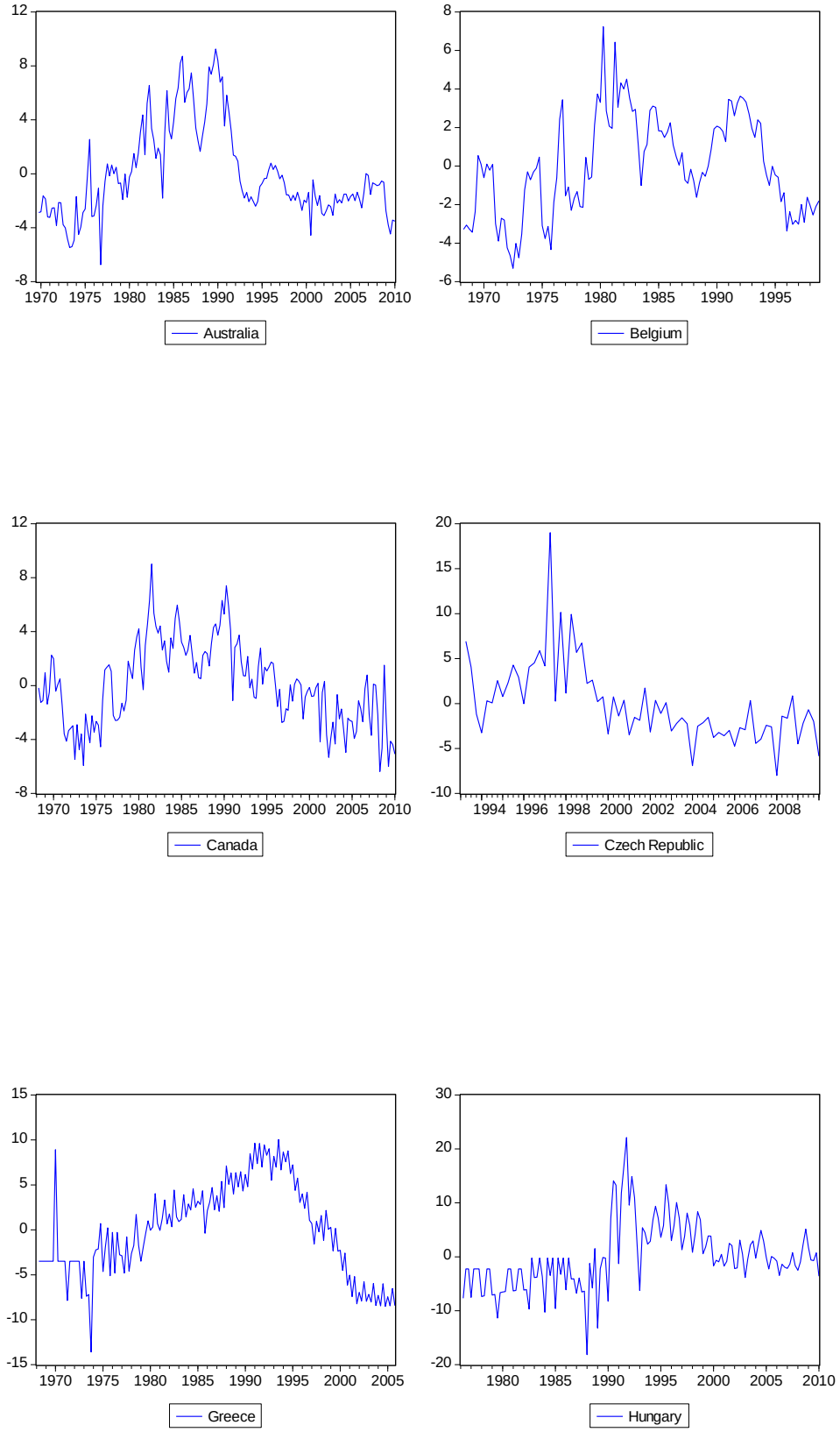
Table 5. ADF unit root test

			Level		First Difference	
	Sample period		ADF_c	ADF_t	ADF_c	ADF_t
Australia	1969Q3:2010Q1	i_t	-2.20	-2.55	-10.06*	-10.09*
		p_t	-2.09	-5.24*	-12.12*	-12.11*
Austria	1968Q1:1998Q4	i_t	-2.39	-2.41	-9.01*	-9.04*
		p_t	-1.29	-2.82	-19.00*	-19.07*
Belgium	1968Q1:1998Q4	i_t	-2.52	-2.45	-9.87*	-9.90*
		p_t	-2.27	-3.42***	-19.34*	-19.34*
Canada	1968Q1:2010Q1	i_t	-1.81	-2.47	-10.02*	-10.03*
		p_t	-1.83	-2.98	-13.78*	-13.76*
Czech Republic	1993Q1:2010Q1	i_t	-1.68	-2.12	-12.72*	-12.65*
		p_t	-3.02**	-3.97**	-8.48*	-8.39*
Denmark	1972Q1:2010Q1	i_t	-2.28	-4.21*	-13.97*	-13.97*
		p_t	-1.96	-3.58**	-7.95*	-7.93*
Finland	1981Q1:2005Q3	i_t	-0.71	-2.69	-6.17*	-6.13*
		p_t	-2.47	-2.77	-6.67*	-6.76*
France	1969Q4:2010Q1	i_t	-1.80	-3.12	-8.62*	-8.62*
		p_t	-1.34	-3.75**	-11.36*	-11.34*
Germany	1991Q1:2010Q1	i_t	-2.12	-2.82	-3.98*	-3.97*
		p_t	-6.70*	-7.39*	-11.36*	-11.41*
Greece	1968Q1:2005Q4	i_t	-0.66	-0.57	-7.94*	-8.35*
		p_t	-2.75***	-2.92	-13.90*	-13.89*
Hungary	1976Q1:2010Q1	i_t	-1.95	-1.80	-7.14*	-7.17*
		p_t	-1.98	-2.02	-8.52*	-8.55*
Iceland	1987Q1:2009Q4	i_t	-2.52	-2.49	-8.09*	-8.07*
		p_t	-2.90**	-2.72	-15.59*	-8.73*
Ireland	1968Q1:2006Q2	i_t	-1.67	-2.94	-6.04*	-9.21*
		p_t	-1.76	-2.65	-14.68*	-14.63*
Italy	1971Q1:2010Q1	i_t	-1.70	-3.25***	-8.80*	-8.87*
		p_t	-0.97	-3.33***	-4.92*	-4.93*
Japan	1968Q1:2010Q1	i_t	-2.14	-4.67*	-7.20*	-7.18*
		p_t	-1.97	-4.07*	-7.26*	-7.22*
Korea	1976Q4:2009Q4	i_t	-1.43	-2.78	-9.57*	-9.54*
		p_t	-2.44	-2.91	-15.58*	-15.52*
Luxembourg	1980Q1:1999Q1	i_t	-0.28	-1.81	-7.98*	-8.01*
		p_t	-1.58	-4.53*	-10.28*	-10.21*
Mexico	1977Q1:2010Q1	i_t	-1.59	-2.41	-10.13*	-10.14*
		p_t	-3.02**	-3.52**	-17.47*	-17.41*
Netherlands	1968Q1:2010Q1	i_t	-3.50*	-3.59**	-9.35*	-9.33*
		p_t	-1.89	-2.36	-16.63*	-16.53*
New Zealand	1985Q1:2010Q1	i_t	-2.10	-2.40	-14.50*	-14.38*
		p_t	-6.09*	-6.61*	-10.09*	-10.04*
Norway	1971Q4:2009Q3	i_t	-2.49	-3.37***	-15.30*	-15.42*
		p_t	-1.87	-3.53**	-14.09*	-14.05*

Poland	1991Q1:2010Q1	i_t	-4.91*	-8.17*	-6.18*	-6.59*
		p_t	-2.41	-1.46	-4.83*	-5.28*
Portugal	1976Q1:2000Q1	i_t	-0.83	-2.76	-4.90*	-5.22*
		p_t	-1.56	-9.43*	-10.94*	-10.87*
Spain	1974Q1:2010Q1	i_t	-1.86	-3.61**	-12.91*	-12.89*
		p_t	-1.09	-1.90	-7.07*	-7.05*
Sweden	1968Q1:2009Q1	i_t	-1.71	-2.16	-11.72*	-11.73*
		p_t	-1.30	-3.82**	-7.84*	-7.86*
Switzerland	1975Q4:2010Q1	i_t	-2.13	-2.21	-13.88*	-13.87*
		p_t	-2.87***	-3.65**	-15.53*	-15.48*
Turkey	1989Q1:2010Q1	i_t	-1.68	-2.83	-10.08*	-8.35*
		p_t	-1.98	-4.30*	-12.02*	-11.98*
United Kingdom	1968Q1:2010Q1	i_t	-2.00	-2.96	-9.79*	-9.82*
		p_t	-2.69***	-3.79**	-7.77*	-7.74*
United States	1974Q1:2010Q1	i_t	-1.75	-4.01**	-5.27*	-5.35*
		p_t	-2.74***	-3.63**	-17.00*	-16.93*

***,**,* denote significance at the 10%,5% and 10% significance levels. ADF_c is the ADF with an intercept and ADT_t with an intercept and a deterministic trend. The critical values are for MacKinnon(1996). The 10%, 5% and 10% significance levels are -2.58,-2.89, and -3.49 for ADF_c and -3.15,-3.45, and -4.05 for ADF_c.

Figure 1. Engle-Granger OLS residuals: nominal interest rate is the dependent variable



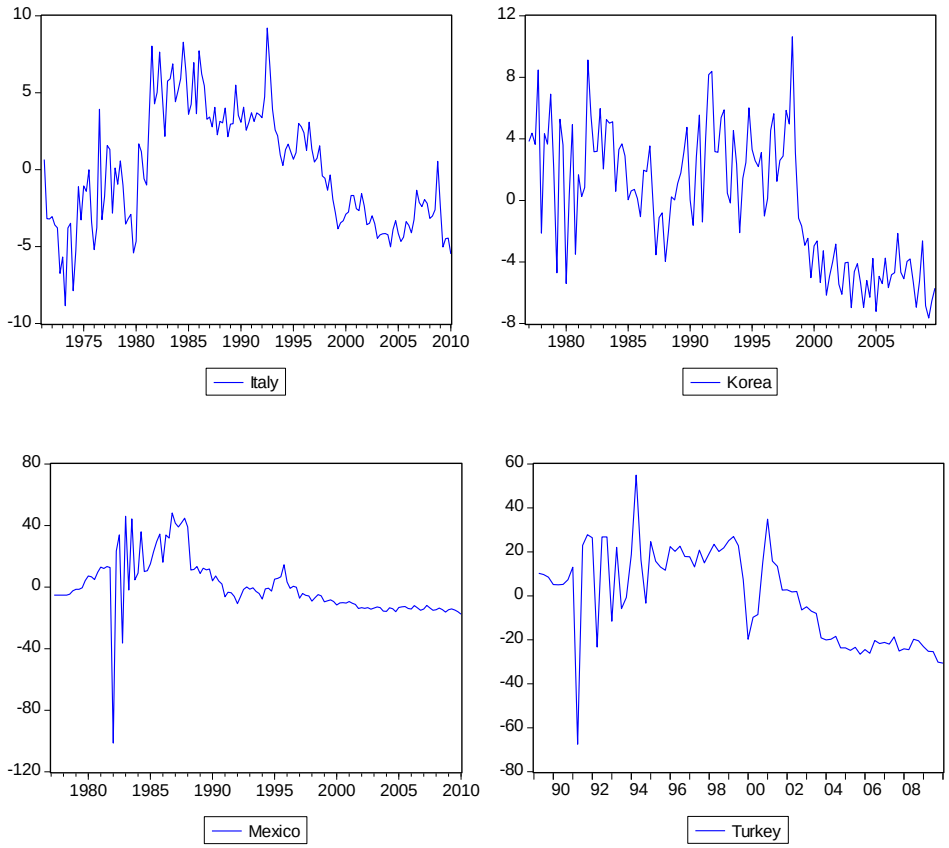
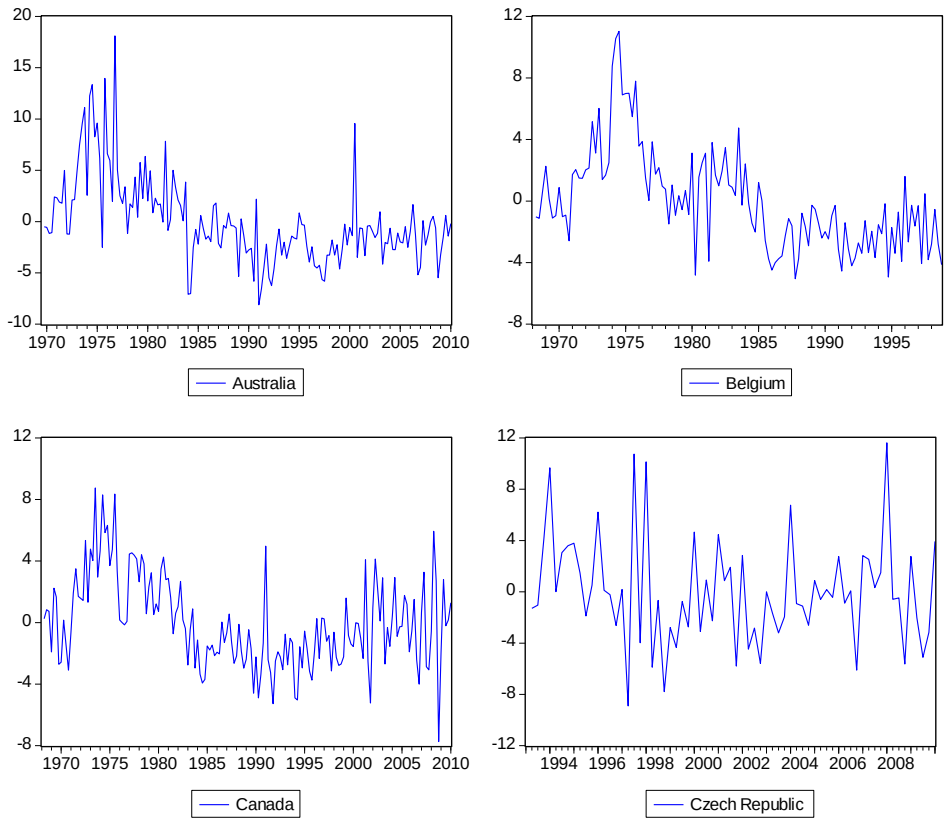
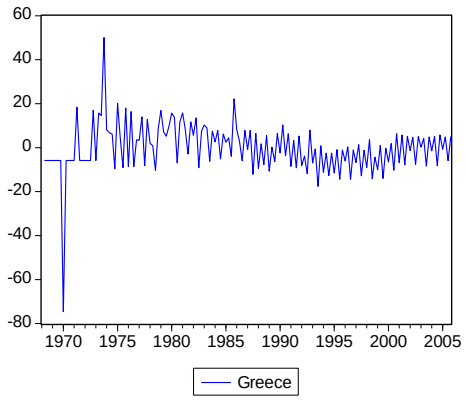
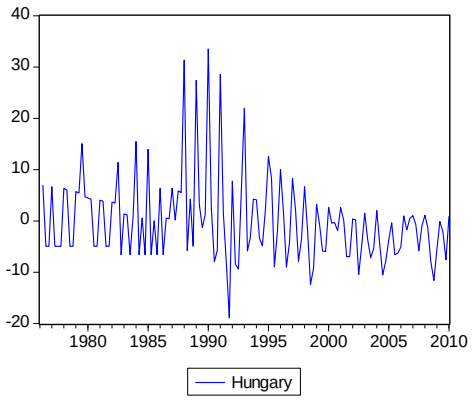
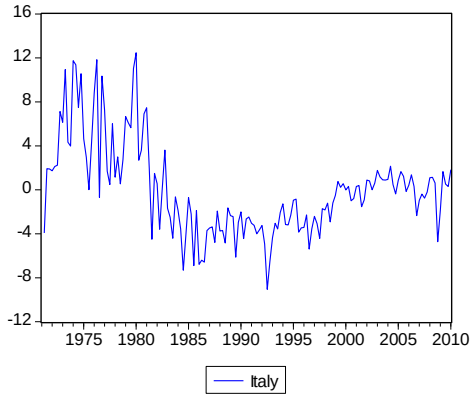


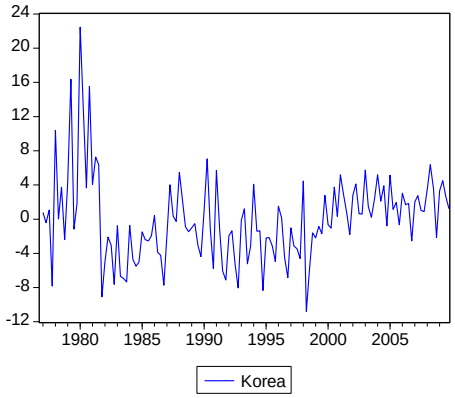
Figure 2. Engle-Granger OLS residuals: inflation is the dependent variable

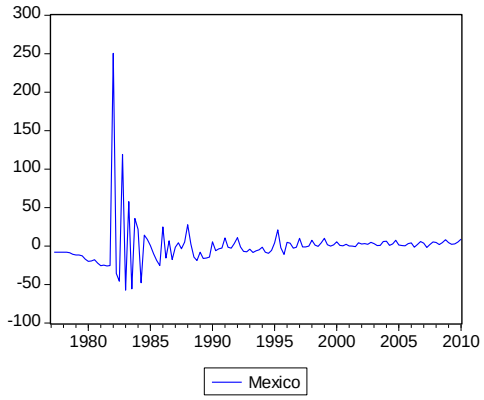



60
40
20
0
-20
-40
-60
-80
1970 1975 1980 1985 1990 1995 2000 2005
Greece


40
30
20
10
0
-10
-20
1980 1985 1990 1995 2000 2005 2010
Hungary


16
12
8
4
0
-4
-8
-12
1975 1980 1985 1990 1995 2000 2005 2010
Italy


24
20
16
12
8
4
0
-4
-8
-12
1980 1985 1990 1995 2000 2005
Korea


300
250
200
150
100
50
0
-50
-100
1980 1985 1990 1995 2000 2005 2010
Mexico

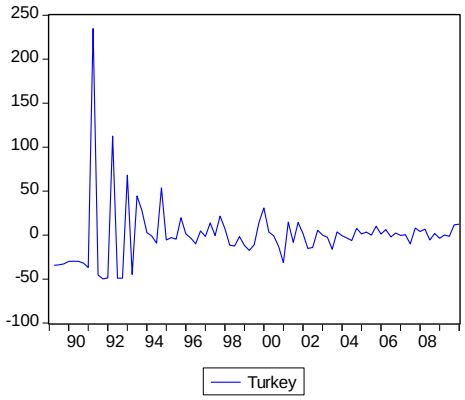

250
200
150
100
50
0
-50
-100
90 92 94 96 98 00 02 04 06 08
Turkey

Table 6: Engle-Granger cointegration test
Nominal interest rate is the dependent variable

Country	a_o	a_1	DW statistic	R^2	ADF	Wald test a_1=1
Australia	6.25	0.31	0.21	0.15	-2.95	137.14
	(14.59)*	(5.27)*				(0.00)*
Belgium	5.51	0.26	0.29	0.11	-3.21***	137.14
	(14.30)*	(3.86)*				(0.00)*
Canada	4.12	0.63	0.36	0.36	-2.28	32.99
	(11.19)*	(9.60)*				(0.00)*
Czech Republic	4.09	0.53	1.05	0.27	-2.27	18.30
	(5.62)*	(4.90)*				(0.00)*
Greece	9.29	0.18	0.32	0.16	-0.73	570.64
	(16.78)*	(5.26)*				(0.00)*
Hungary	5.23	0.46	0.72	0.37	-2.43	113.30
	(6.88)*	(8.91)*				(0.00)*
Italy	4.94	0.66	0.28	0.52	-1.34	46.24
	(10.43)*	(12.96)*				(0.00)*
Korea	7.29	0.57	0.59	0.43	-1.96	55.98
	(14.10)*	9.90				(0.00)*
Mexico	4.09	(0.53)*	1.05	0.27	-2.27	170.50
	(8.05)*	9.34				(0.00)*
Turkey	41.82	0.31	0.81	0.28	0.23	150.08
	(13.69)*	(5.60)*				(0.00)*

***, **, * denote significance at the 10%, 5% and 1% significance levels. The numbers in parenthesis are t-values for the null hypothesis that a_0 and a_1 are statistically equal to zero. The 10%, 5% and 1% critical values are 1.64 , 1.96 and 2.57 respectively. The critical values for ADF test are provided by MacKinnon (1991, Table 1).The 10%, 5% and 1% critical values are -3.09,-3.39 and -4.00 respectively. Wald test has a Chi square distribution with one degree of freedom, $\chi^2(1)$, since there is only one restriction. The numbers in parenthesis for the Wald test are the p-values.

Table 7: Engle-Granger cointegration test
Inflation Rate is the dependent variable

Country	a_o	a_1	DW statistic	R^2	ADF	Wald test a_1=1
Australia	1.87	0.48	0.89	0.15	-2.20	33.45
	(2.35)**	(5.27)*				(0.00)*
Belgium	1.75	0.42	0.63	0.11	-1.52	28.17
	(2.22)**	(3.86)*				(0.00)*
Canada	0.45	0.57	0.86	0.36	-2.59	52.23
	(0.97)	(9.60)*				(0.00)*
Czech Republic	1.33	0.50	2.30	0.27	-4.90*	24.06
	(1.58)	(4.90)*				(0.00)*
Greece	0.82	0.87	1.91	0.16	-3.39***	0.64
	(0.40)	(5.26)*				(0.43)
Hungary	2.48	0.82	1.90	0.37	-3.46**	4.00
	(2.13)**	(8.91)*				(0.05)***
Italy	-0.53	0.80	0.51	0.52	-1.73	11.10
	(-0.78)	(12.96)*				(0.00)*
Korea	-2.01	0.75	1.29	0.43	-2.41	10.36
	(-2.16)**	(9.90)*				(0.00)*
Mexico	1.33	0.50	2.30	0.27	-4.90*	0.13
	(-0.51)	(9.34)*				(0.72)
Turkey	-11.65	0.88	2.42	0.28	-8.91*	0.56
	(-1.27)	(5.60)*				(0.45)

***, **, * denote significance at the 10%, 5% and 1% significance levels. The numbers in parenthesis are t-values for the null hypothesis that a_0 and a_1 are statistically equal to zero. The 10%, 5% and 1% critical values are 1.64 , 1.96 and 2.57 respectively. The critical values for ADF test are provided by MacKinnon (1991, Table 1).The 10%, 5% and 1% critical values are -3.09,-3.39 and -4.00 respectively. Wald test has a Chi square distribution with one degree of freedom, $\chi^2(1)$, since there is only one restriction. The numbers in parenthesis for the Wald test are the p-values.

Table 8. Gregory-Hansen cointegration test

	Nominal interest rate is the dependent variable		Inflation is the dependent variable	
	ADF(τ)	Break Date	ADF(τ)	Break Date
Australia				
Model 1	-3.43	1993Q3	-4.56***	1982Q3
Model 2	-4.61	1980Q2	-4.55	1982Q3
Model 3	-3.37	1994Q1	-4.53	1981Q4
Belgium				
Model 1	-4.71**	1977Q1	-4.18	1977Q4
Model 2	-5.30**	1978Q4	-3.05	1993Q2
Model 3	-4.18	1977Q4	-3.05	1978Q1
Canada				
Model 1	-3.64	1998Q1	-4.37***	1983Q1
Model 2	-5.53*	1980Q2	-4.65	1981Q4
Model 3	-3.69	2002Q2	-4.50	1982Q2
Czech Republic				
Model 1	-3.83	2000Q2	-4.99**	1996Q4
Model 2	-3.77	2000Q2	-5.12**	2005Q2
Model 3	-3.75	2000Q2	-4.81***	1997Q3
Greece				
Model 1	-1.94	1998Q2	-4.10	1987Q2
Model 2	-3.66	1997Q4	-4.86**	1973Q3
Model 3	-1.98	1999Q1	-3.41	1987Q1
Hungary				
Model 1	-3.44	1986Q4	-2.58	1998Q3
Model 2	-5.31**	1988Q3	-5.61*	1990Q1
Model 3	-3.81	1988Q1	-2.51	1988Q2
Italy				
Model 1	-3.07	1999Q1	-7.49*	1984Q2
Model 2	-4.48	1978Q3	-4.59	1982Q3
Model 3	-3.19	1982Q4	-5.80*	1982Q4
Korea				
Model 1	-5.55*	1997Q3	-4.89**	1981Q4
Model 2	-5.17**	1990Q2	-5.10**	1983Q1
Model 3	-5.85*	1997Q3	-5.49*	1984Q2
Mexico				
Model 1	-3.16	1991Q4	-5.16*	1995Q4
Model 2	-3.74	1982Q2	-5.65*	1984Q1
Model 3	-3.79	1980Q4	-6.12*	1983Q2
Turkey				
Model 1	-5.23*	2002Q4	-7.76*	2003Q3
Model 2	-5.46*	1994Q2	-7.94*	1995Q4
Model 3	-7.06*	1993Q4	-8.76*	1991Q4

*,**,*** denote significance at the 1%, 5% and 10% significance levels. The 1%, 5%, and 10% critical values are -5.13,-4.61, and -4.34 for Model 1, -5.45,-4.99, and -4.72 for Model 2, -5.47, -4.95 and -4.68 for Model 3.The critical values are provided by Gregory and Hansen (1996, Table

Table 9. DOLS results

	a_0	a_1	a_2	a_3	a_4	Wald Test (a_1=1)
Belgium						
No break	5.33	0.31				30.42
	(8.49)*	(2.43)**				(0.00)*
Model 1	6.73	0.12	-2.7			32.44
	(8.30)*	(-0.78)	(-3.47)*			(0.00)*
Model 2	1.6	0.48	-4.99	0.06		12.41
	(-1.22)	(3.20)*	(-5.69)*	(4.57)*		(0.00)*
Canada						
Model 2	2.04	0.91	-3.41	0.02		0.23
	(-1.27)	(4.75)*	(-3.08)*	(-1.16)		(-0.63)
Hungary						
Model 2	-4.58	0.99	-1	0.06		0.02
	(-4.17)*	(12.19)*	(-0.90)	(4.47)*		(-0.88)
Korea						
Model 1	7.35	0.69	-5.51			10.8
	(7.12)*	(7.38)*	(-6.11)*			(0.00)*
Model 2	12.6	0.48	-1.95	-0.07		39.06
	(9.62)*	(5.79)*	(-1.26)	(-3.34)*		(0.00)*
Model 3	7.34	0.69	-5.26		-0.08	10.64
	(7.09)*	(7.35)*	(-5.48)*		(-1.22)	(0.00)*
Turkey						
Model 1	13.3	1.07	0.1			1.4
	(4.53)*	(17.26)*	(-0.02)			(-0.24)
Model 2	1.4	1.2	-2.1	0.17		2.52
	(-0.14)	(9.64)*	(-0.64)	(-1.19)		(-0.12)
Model 3	13.3	1.07	-0.49		0.07	328.88
	(5.65)*	(18.14)*	(-0.24)		0.54	(0.00)*

***, **, * denote significance at the 10%, 5% and 1% significance levels. The numbers in parenthesis are t-values for the null hypothesis that a_0,a_1,a_2 ,a_3 and a_4 are statistically equal to zero. The 10%, 5% and 1% critical values are 1.64 , 1.96 and 2.57 respectively. Wald test has a Chi square distribution with one degree of freedom, $\chi^2(1)$, since there is only one restriction. The numbers in parenthesis for the Wald test are the p-values.

References

Berument, Hakan, and Mohamed Mehdi J,(2002),” The Fisher hypothesis: a multi-country analysis”, Applied economics, 34 , 1645-55

Carmichael, J. and Stebbing, P.W.,(1983), “ Fisher’s paradox and the theory of interest”, American Economic Review, 73, 619-630.

Choudhry, A.,(1997), ”Cointegration analysis of the inverted Fisher effect: Evidence from Belgium, France and Germany”, Applied Economics Letters,4, 257-260.

Crowder, W.J. and Hoffman, D.L.,(1996),” The Long-run relationship between nominal interest rates and inflation: The Fisher equation revisited”, Journal of Money, Credit, and Banking, 28, 102-118.

Darby,M.R.,(1975),”The financial and tax effects of monetary policy on interest rates”, Economic Inquiry, 13, 266-276.

Dickey, D. and Fuller, W,(1979),”Distribution of the estimators for autoregressive time series with a unit root, Journal of the American Statistical Association, 74, 427-431.

Elliott, G., Rothenberg, T. and Stock, J.,(1996).”Efficient tests for an autoregressive unit root”, Econometrica, 64, 813-836.

Engle, R.F. and Granger, C.W.J., (1987) ,” Co-integration and error correction: Representation, estimation, and testing. Econometrica, 55, 251-276.

Engsted, T., (1995), “Does the long-term interest rate predict future inflation? A multi-country analysis”, Review of Economics and Statistics, 77, 42-54.

Feldstein, M.,(1976), ”Inflation, income tax and the rate of interest: A theoretical analysis”, American Economic Review, 66, 809-820.

Fisher, I., (1930),”The Theory of Interest”, New York: Macmillan.

Garcia, R., Perron, P., (1996), ” An analysis of the real interest rate under regime shifts”, Review of Economics and Statistics ,78, 111-125.

Goldberg, Lawrence G., James Lothian, and John Okunev, (2003), ”Has international financial integration increased?”,Open Economies Review,14, 299-317

Granger, C.W.J. and Newbold, P., (1974),”Spurious regression in Econometrics”, Journal of Econometrics **,** 2 , 111-120.

Gregory, A. and Hansen, B. ,(1996), “Residual-based tests for cointegration in models with regime shifts”, Journal of Econometrics, 70, 99-126.

Gregory, Allen W., James M. Nason, and David G. Watt.,(1994),”Testing for structural breaks in cointegrated relationships”,Journal of Econometrics, 71, 321-41

Johansen, S., (1988),”Statistical analysis of cointegration vectors “, Journal of Economic Dynamics and Control, 12, 231-254.

King, R.G. and Watson, M.W.,(1997),”Testing long-run neutrality”, Federal Reserve Bank of Richmond Economic Quarterly ,83, 69-101.

Koustas, Z. and Serletis, A., (1999), “On the Fisher effect”, Journal of Monetary Economics, 44, 105-130.

Lee, J., and Strazicich, M. , (2003) , “Minimum LM unit root test with two structural breaks”, Review of Economics and Statistics 85, 1082–1089.

Mishkin, F.S., (1992),”Is the Fisher effect for real? A reexamination of the relationship between inflation and interest rates”, Journal of Monetary Economics, 30, 195-215.

Moazzami, B., (1991),”The Fisher equation controversy re-examined”, Applied Financial Economics,1, 129-133.

Mackinnon, J.G.,(1991),”Critical values for cointegration tests. In Long-run economic relationships: readings in cointegration”, eds, R.F. Engle and C.W.J. Granger, 267–276, Oxford: Oxford Press.

Mundell, R.,(1963),”Inflation and real interest”,The journal of Political Economy, 71 ,280-283.

Nelson, C.R. and Plosser, C.I.,(1982),” Trends and random walks in macroeconomic time series”, Journal of Monetary Economics**,** 10, 139-162.

Ng, S., and P. Perron, (2001). “Lag Length Selection and the Construction of Unit Root Tests with Good Size and Power,” Econometrica, 69, 1519-1554.

Perron, P., (1989),” The great crash, the oil price shock, and the unit root hypothesis”, Econometrica, 57, 1361-1401.

Perron, P. and S. Ng., (1996),” Useful Modifications to Some Unit Root Tests with Dependent Errors and their Local Asymptotic Properties” ,Review of Economic Studies, 63, 435-463.

Phillips, P. C.B. and Hansen, B.E., (1990), ”Statistical inference in Instrumental variables, regression with I(1) processes”**,** Review of Economic Studies, 57,99-125.

Phillips, P.C.B. and Perron, P.,(1988),”Testing for a unit root in time series regression”. Biometrika, 75, 335-346.

Rapach, D.E. ,(2003), “International evidence on the long run impact of inflation”, Journal of Money, Credit, and Banking, 35, 23-48.

Rose, A. K. ,(1988),” Is the real interest rate stable?”, Journal of Finance, 43, 1095–112.

Said, S.E. and D. Dickey (1984). “Testing for Unit Roots in Autoregressive Moving-Average Models with Unknown Order,” Biometrika, 71, 599-607.

Shome, D.K., Smith, S.D. and Pinkerton, J.M.,(1988),”The purchasing power of money and nominal interest rates: A re-examination, Journal of Finance, 43, 1113-25.

Schwert, W., (1989). “Test for Unit Roots: A Monte Carlo Investigation”, Journal of Business and Economic Statistics, 7, 147-159.

Sidrauski, M. ,(1967),”Rational choice and patterns of growth in a monetary economy, American Economic Review Papers and Proceedings,57, 534-544

Stock, J.H. and Watson, M.W., (1993),”A simple estimator of cointegrating vector in higher order integrated system”, Econometrica, 61, 783-820.

Tobin, J., (1965),”Money and economic growth”, Econometrica,33, 671-684.

Weidmann, J., (1997),” New hope for the Fisher Effect? A reexamination using threshold cointegration”, Revised version of Sonderforschungsbereich ,303, Discussion Paper B-385, Bonn: University of Bonn.

Wesso, G.R., (2000),”Long-term yield bonds and future inflation in South Africa: a vector error-correction analysis”, Quarterly Bulletin, Pretoria: South African Reserve Bank

Westerlund, J., (2008), ”Panel cointegration tests of the Fisher effect”, Journal of Applied Econometrics, 23, 193-233

Wong, Ka-Fu, and Hai-Jun Wu, (2003), “Testing Fisher hypothesis in long horizons for G7 and eight Asian countries’, Applied Economics Letters, 10, 917–23.

Woodward, G.T., (1992),”Evidence of the Fisher effect from UK Indexed bonds”, Review of Economics and Statistics**,** 74, 315-320.