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Exchange Rates and Fundamentals: Evidence from Long-horizon Regression Tests

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Abstract

In this paper, we apply the alternative long-horizon regression approach proposed by Fisher and Seater (1993) to study the long-run relationship between nominal exchange rates and fundamentals. We find evidence supporting the explanatory power of exchange rate models. In particular, the Taylor-rule model outperforms other conventional models. We then use the inverse power function (IPF) proposed by Andrews (1989) to investigate the power of the Fisher–Seater test. The IPFs provide more evidence to support exchange rate models.

Keywords: exchange rates, fundamentals, long-horizon tests.

JEL Classification: C22, F31

1 Introduction

Since the seminal paper by Meese and Rogoff (1983), it has been puzzling that standard models fail to predict exchange rates. Exchange rates are empirically disconnected from economic fundamentals such as prices, interest rates, output, and the money supply. Recently, long-horizon regression tests have been widely used to investigate the relationship between exchange rates and fundamentals. A significant study by Mark (1995) investigates whether monetary fundamentals can forecast exchange rate movements. Having defined fundamentals as $x_t \equiv (m_t - m_t^*) - (y_t - y_t^*)$, Mark (1995) runs regressions of the form:

$$s_{t+k} - s_t = \alpha_k + \beta_k(x_t - s_t) + \varepsilon_{t+k}, \tag{1}$$

where m_t and y_t are the logs of the money supply and real output, respectively, and s_t is the log of the foreign currency price of the home currency. (Asterisks denote foreign-country variables.) If s_t exceeds x_t, the currency is predicted to appreciate. Furthermore, if the fundamentals in question help to forecast the exchange rate, one would expect β_k to be significantly positive. Mark (1995) estimates this parameter for the U.S. dollar relative to the currencies of Canada, Switzerland, Germany, and Japan from 1973Q2 to 1991Q4. Horizons, k, of 1, 4, 8, 12, and 16 quarters are chosen. Mark (1995) finds evidence that current-period deviations from the fundamental values help to predict future changes in nominal exchange rates. The results suggest increased long-horizon predictability. Both the goodness of the in-sample fit and the estimated value of β_k increase with the horizon k. In particular, at 16 quarters, Mark (1995) finds strong evidence that β_k is significant. The model that incorporates monetary fundamentals outperforms a random-walk model in terms of out-of-sample forecasts. Mark concludes that systematic exchange rate movements that are related to fundamentals become apparent in long-horizon changes, whereas quarter-to-quarter movements may be noisy. Moreover, Chinn and Meese (1995) and Chen and Mark (1996) also find that monetary fundamentals can be used to forecast nominal exchange rate returns. Mark and Sul (2001) extend Mark (1995) by using quarterly panel data on 19 currencies from 1973Q1 to 1997Q1. Using panel-data estimation increases the explanatory power of the regression. Panel-data predictive regression estimates and panel-based forecasts show that monetary fundamentals have strong predictive power. Engel et al. (2007) confirm the usefulness of economic models in forecasting exchange rates over a horizon of 16 quarters by using panel- data error-correction techniques.

The important results obtained by Mark (1995), Mark and Sul (2001) and Engel et al. (2007) suggest resolutions of the Meese and Rogoff (1983) puzzle and the exchange rate disconnect puzzle. However, Berkowitz and Giorgianni (2001) have shown that the results obtained by Mark (1995) may overstate statistical significance, and have argued that there is no justification for imposing a cointegrating relationship between nominal exchange rates and fundamentals. When Berkowitz and Giorgianni (2001) reestimate the relationship without assuming cointegration, they find no statistically significant evidence of improved long-horizon predictability. In addition, having updated Mark's (1995) data set, Kilian (1999) finds no evidence of increased predictability at longer horizons. Using a century of data, Rapach and Wohar (2002) find that simple monetary fundamentals fail to predict exchange rate movements for six out of the 10 countries they investigated: Australia, Canada, Denmark, Norway, Sweden, and the U.K. Adopting a new inference procedure developed by Clark and West (2007) in a recent study, Molodtsova and Papell (2007) finds no significant evidence that economic models have better predictive ability than random models over long horizons. Rapach and Wohar (2004) question the validity of the panel tests used by Mark and Sul (2001). Rapach and Wohar (2004) show that homogeneity restrictions must be imposed on the cointegrating coefficients to find support for the monetary model. However, these restrictions are rejected by formal tests. Thus, the support for the monetary model provided by panel estimation may be spurious. In summary, evidence of a long-run relationship between exchange rates and fundamentals seems mixed.

In this paper, to study the long-run relationship between nominal exchange rates and fundamentals, we apply the alternative long-horizon regression approach proposed by Fisher and Seater (1993). The Fisher–Seater approach has been widely used to investigate the long-run neutrality of money. In the field of international finance, their method has been applied to test for purchasing power parity by Serletis and Gogas (2004) and Wallace and Shelley (2006). To the best of our knowledge, our paper represents the first attempt to use the Fisher–Seater approach to investigate the relationship between exchange rates and fundamentals.

There are two reasons for using the Fisher–Seater approach as an alternative to adopting the standard approach of outperforming a random-walk model in an out-of-sample forecasting contest. First, as argued by Engel and West (2005) and Engel et al. (2007), judging exchange rate models based on their out-of-sample predictive power against random walks is unreliable. Hence, we conduct an in-sample investigation of the relationship between fundamentals and exchange rates according to the recommendation from Engel et al. (2007). Second, by using out-of-sample forecasting tests, Molodtsova and Papell (2007) have obtained strong evidence to support exchange rate predictability over a one-month horizon and have found some evidence to support predictability over a six-month horizon, but found no evidence to support predictability over longer horizons. Therefore, in this paper, our use of an alternative approach represents a useful opportunity to reinvestigate their evidence of decreasing predictive power over longer horizons. We would like to know whether we can reject the null no predictive power in longer horizons.

2 Empirical Model

2.1 The Fisher–Seater Approach

Fisher and Seater (1993) develop a reduced-form test from a bivariate autoregressive model on the basis that the variables in the model are integrated of the same order. Consider the following bivariate autoregressive representation:

$$\alpha_{ss}(L)\Delta^{\langle s\rangle}s_t = \alpha_{sf}(L)\Delta^{\langle f\rangle}f_t + \varepsilon_t^s, \tag{2}$$

$$\alpha_{ff}(L)\Delta^{\langle f\rangle}f_t = \alpha_{fs}(L)\Delta^{\langle s\rangle}s_t + \varepsilon_t^f, \tag{3}$$

where $\alpha_{ss}^0 = \alpha_{ff}^0 = 1$, $\Delta = 1 - L$ and L is the lag operator. s_t is the log of the nominal exchange rate and f_t is the log of fundamentals. For instance, f_t could be a monetary fundamental such as $f_t = (m_t - m_t^*) - (y_t - y_t^*)$, which represents the benchmark case in Mark (1995). $\langle x\rangle$ represents the order of integration of the variable x, where $x = \{s, f\}$. If y is integrated of order d, then $\langle x\rangle = d$ and $\langle \Delta x\rangle = \langle x\rangle - 1$. The joint distribution of ε_t^s and ε_t^f is assumed to be:

$$\begin{bmatrix} \varepsilon_t^s \\ \varepsilon_t^f \end{bmatrix} \sim^{i.i.d.} \left(\begin{bmatrix} 0 \\ 0 \end{bmatrix}, \begin{bmatrix} var(\varepsilon_t^s) & cov(\varepsilon_t^s, \varepsilon_t^f) \\ cov(\varepsilon_t^f, \varepsilon_t^s) & var(\varepsilon_t^f) \end{bmatrix} \right).$$

Hence, a test of the long-run relationship between exchange rates and fundamentals can be based on the long-run derivative of s_t with respect to a permanent change in f_t, which is defined as follows: if $\lim_{k\to\infty} \partial f_{t+k}/\partial\varepsilon_t^f \neq 0$, then

$$\mathrm{LRD}_{s,f} \equiv \lim_{k\to\infty} \frac{\partial s_{t+k}/\partial\varepsilon_t^f}{\partial f_{t+k}/\partial\varepsilon_t^f}.$$

That is, the $\text{LRD}_{s,f}$ represents the ultimate effect of an exogenous fundamental shock on the nominal exchange rate, s_t, relative to the ultimate effect of the shock to the fundamental, f_t. $\text{LRD}_{s,f}$ is undefined when $\lim_{k\to\infty} \partial f_{t+k}/\partial \varepsilon_t^f = 0$. We test the null hypothesis that $\text{LRD}_{s,f} = 0$. Rejection of the null hypothesis implies that a permanent change in the fundamentals has a long-run effect on the exchange rate; that is, fundamentals and exchange rates are connected in the long run. For instance, if we consider the purchasing power parity (PPP) fundamental, $f_t = p_t - p_t^*$, where p_t denotes the price level, then PPP requires that $\text{LRD}_{s,f} = 1 \neq 0$.

The moving average representation of the above bivariate autoregressive model can be written as:

$$s_t = \Delta^{-\langle s\rangle}[\phi_{sf}(L)\varepsilon_t^f + \phi_{ss}(L)\varepsilon_t^s], \tag{4}$$

$$f_t = \Delta^{-\langle f\rangle}[\phi_{ff}(L)\varepsilon_t^f + \phi_{fs}(L)\varepsilon_t^s], \tag{5}$$

$$\tag{6}$$

where

$$\phi_{ff}(L) = \alpha_{ss}/[\alpha_{ff}(L)\alpha_{ss}(L) - \alpha_{fs}(L)\alpha_{sf}(L)],$$
$$\phi_{ss}(L) = \alpha_{sx}/[\alpha_{ff}(L)\alpha_{ss}(L) - \alpha_{fs}(L)\alpha_{sf}(L)],$$

and so on. Fisher and Seater (1993) show that when $\langle f\rangle \geq 1$, it follows that:

$$\text{LRD}_{s,f} = \frac{(1-L)^{\langle f\rangle - \langle s\rangle}\phi_{sx}(L)|_{L=1}}{\phi_{ff}}. \tag{7}$$

Clearly, $\text{LRD}_{s,f}$ depends on $\langle f\rangle - \langle s\rangle$. Fisher and Seater (1993) considered different orders of $\langle f\rangle - \langle s\rangle$. For instance, when $\langle f\rangle - \langle s\rangle = 0$, the long-run derivative is

$$\text{LRD}_{s,f} = \frac{\phi_{sf}(1)}{\phi_{ff}(1)}.$$

Under the assumption that $cov(\varepsilon_t^s, \varepsilon_t^f) = 0$, Fisher and Seater (1993) show that $\text{LRD}_{s,f}$ can be estimated by $\lim_{k\to\infty} b_k$, where b_k is the slope coefficient from the following regression:

$$\left[\sum_{j=0}^{k} \Delta^{\langle s\rangle} s_{t-j}\right] = a_k + b_k \left[\sum_{j=0}^{k} \Delta^{\langle f\rangle} f_{t-j}\right] + e_{k,t}. \tag{8}$$

When $\langle f \rangle = \langle s \rangle = 1$, equation (8) becomes:

$$(s_t - s_{t-k-1}) = a_k + b_k(f_t - f_{t-k-1}) + e_{k,t}. \tag{9}$$

We test the null hypothesis that $b_k = 0$. If the confidence interval does not include zero when k increase, it may imply that the fundamental has predictive power on the exchange rate, that is, exchange rates and fundamentals are connected in the long-run.[1]

2.2 Fundamentals

Following Molodtsova and Papell (2007) and Engel et al. (2007), the choice of fundamentals is informed by standard exchange rate models.[2] The variables below are all in logs except for interest rates; asterisks denote foreign country variables.

1. Monetary Fundamentals:

$$f_t = (m_t - m_t^*) - \eta(y_t - y_t^*),$$

   where m_t and y_t are the money supply and real output, respectively. Note that the income elasticity of money demand, η, is assumed to be one or zero. That is, we consider $f_t = (m_t - m_t^*) - (y_t - y_t^*)$ and $f_t = m_t - m_t^*$.

2. PPP Fundamentals:

$$f_t = p_t - p_t^*,$$

   where p_t denotes the price level.

3. Taylor-rule Fundamentals:

$$f_t = 1.5(\pi_t - \pi_t^*) + 0.1(y_t - y_t^*) + 0.1(s_t + p_t^* - p_t) + s_t,$$

   where π_t represents the inflation rate. Values of the parameters in the Taylor rule are simply adopted from Engel et al. (2007), and are arguably standard.

[1] Note that predictive power does not mean the ability of current fundamentals to predict future exchange rates. In the context of the Fisher–Seater approach, we focus on how well contemporaneous movements in fundamentals predict contemporaneous movements in the exchange rate.

[2] See Molodtsova and Papell (2007) and Engel et al. (2007) for detailed derivations.

4. Uncovered Interest Rate Parity (UIP) Fundamentals:

$$f_t = (i_t - i_t^*) + s_t,$$

   where i_t is the interest rate.

5. Other Fundamentals: other fundamentals considered informally (that is, without being based on any structural model) are: the terms of trade (TOT), $(TOT_t - TOT_t^*)$, and the current account balance $(CA_t - CA_t^*)$.

3 Data and Preliminary Tests

We analyze data from the G7 countries, comprising Canada, France, Germany, Italy, Japan, the U.K., and the U.S.. The U.S. is chosen to be the numeraire. Quarterly data, typically from 1972Q1 to 2005Q4, are used. Data on nominal exchange rates (foreign currency per unit of domestic currency), the money supply (M2), the consumer price index (CPI), real GDP, the TOT, and the current account balance were obtained from Datastream and International Financial Statistics (IFS) published by the International Monetary Fund (IMF). All variables are in logarithms. Variable names and data codes are provided in Tables 1. An abnormally sharp fall in the German CPI was recorded by Datastream between 1990Q4 and 1991Q1. We thus replaced these data with CPI data from IFS. Nominal exchange rate data are from IFS.

We first performed the Augmented Dickey–Fuller (ADF) and Phillips–Perron (PP) tests on the variables. The results in Tables 2 and 3 suggest that all variables are I(1). This justifies using equation (9) to test the long-run predictive ability of economic fundamentals. Although cointegration is neither necessary nor sufficient for tests on the long-run derivative, we used the Engle–Granger and Johansen procedures to test whether exchange rates and fundamentals are cointegrated. For most countries, the test results suggest no evidence of cointegration between fundamentals and exchange rates. This is consistent with the finding of Engel and West (2005).[3]

[3] The results of the cointegration tests are available on request.

4 Empirical Results

Results of the Fisher–Seater test are presented in Figures 1 to 7. We consider values of k ranging from 1 to 30 and plot the estimates of b_k along with their 95% confidence intervals. The confidence intervals are based on Newey–West standard errors and use a t-distribution with T/k degree of freedom, where T is the number of observations.[4] Rejection of $b_k = 0$ suggests that exchange rates and fundamentals are connected in the long run.

The results for monetary fundamentals are presented in Figures 1 and 2. For three of the six countries (France, Germany, and Japan), we found strong evidence to support a long-run connection between exchange rates and monetary fundamentals. There is also evidence to support such a connection for the U.K., given that the null of $b_k = 0$ is rejected only for $k \in [1, 24]$. The result for Italy is somewhat weaker: the null hypothesis is only rejected for $k \in [14, 30]$ when $f_t = m_t - m_t^*$, and is rejected for $k \in [1, 13]$ when $f_t = (m_t - m_t^*) - (y_t - y_t^*)$. For Canada, zero is included in the confidence intervals for all horizons, and the coefficient converges to zero when k becomes large.

Figure 3 shows the results for the PPP fundamentals. In all countries except Canada and Germany, we strongly reject $b_k = 0$. Moreover, for the PPP fundamentals, it is also worth testing $b_k = 1$, which is the null of PPP. The figure shows that, for France, Italy, and the U.K., the null of no connection between exchange rates and fundamentals is strongly rejected. For these countries, long-run PPP is supported given that unity falls into the 95% confidence intervals when the horizon k increases. The exception is Japan: although the null of fundamentals having no predictive power is rejected, there is evidence of a significant deviation from PPP.

The results in Figure 4 for the Taylor-rule fundamentals are the most interesting. The null hypothesis of no connection is significantly rejected for all countries and for all horizons. This result is consistent with the results of Molodtsova and Papell (2007), who find that the out-of-sample predictive ability of Taylor-rule models exceeds that of conventional models. However, there is an obvious difference between our in-sample

[4]For the Newey–West estimation, a Bartlett kernel with a default bandwidth of $4(T/100)^{2/9}$ was used.

evidence and their out-of-sample results. Molodtsova and Papell (2007) find no significant evidence that the predictive ability of Taylor-rule models exceeds that of a random walk over long horizons. By contrast, our results suggest an overall significance against null of no predictive power of Taylor-rule models from short to longer horizons.

For the UIP fundamentals, we find some support for the predictive power of the UIP model, particularly when the horizon k is large. For Germany, we reject the null of no connection for values of k that exceed 29. For Canada and Italy, the null is rejected for $k \in [12, 28]$ and $k \in [1, 27]$, respectively.

Next, we present results for the TOT and the current account in Figures 6 and 7. When the TOT is considered as a fundamental, there is strong evidence of a connection in all six countries. For France, Germany, Italy, and Japan, the null hypothesis of no connection is rejected for all horizons. For Canada and the U.K., the null $b_k = 0$ is rejected for $k \in [1, 15]$ and $k \in [1, 21]$, respectively. Weaker evidence is obtained when considering the current account as a fundamental. For only half of the six countries (Canada, Italy, and Japan) can we can reject the null hypothesis of no connection.

To sum up, although we do not find overwhelming evidence of a long-run connection between exchange rates and fundamentals, encouraging evidence of such a connection is found for most countries by using the Fisher–Seater long-horizon test. Our in-sample results confirm that exchange models are effective in predicting exchange rates. In particular, the Taylor-rule model outperforms other conventional models in providing the strongest evidence in support of a connection between fundamentals and exchange rates.

5 Inverse Power Functions

Coe and Nason (2003) and Coe and Nason (2004) have suggested that the Fisher–Seater test may have low power and may well induce Type II errors. This may explain the relatively wide confidence intervals shown in Figures 1 to 7. To investigate the issue of low power, we follow Coe and Nason (2003) and use the inverse power function (IPF) developed by Andrews (1989) to provide information about deviations from the null hypothesis (of no long-run connection between exchange rates and fundamentals) when the

Fisher–Seater test fails to reject the null.

In Tables 4 to 9, we report estimates of b_k, tests of $b_k = 0$, and Andrews (1989) IPFs, for forecast horizons k of 5, 10, 15, 20, 25, and 30. An asterisk indicates rejection of the null hypothesis of no connection at the 5% significance level. The IPFs are constructed for a low (namely 5%) probability of a Type II error, $b_{k,0.05}$, and for a high (namely 50%) probability of a Type II error, $b_{k,0.50}$, conditional on failing to reject the null of no connection at the 5% level.

Given a 5% probability of a type II error, which corresponds to a high power of 95%, the IPF yields a region, $\Phi = \{b_k : |b_k| > b_{k,0.05}\}$, of potentially true values of b_k. An inability to reject the null implies that $|b_k|$ is less than $b_{k,0.05}$ at the 5% significance level. That is, any parameter value within Φ is not supported at the 5% significance level. Consider next an IPF based on 50% power: $b_{k,0.50}$. This yields an interval $\Theta = \{b_k : b_k \in [-b_{k,0.50}, b_{k,0.50}]\}$ containing potential true alternatives that are consistent with the data. If the true value of b_k falls into this interval, the power of the test is 0.5 or less. This is the set of potential true values of the parameter for which a test is no more powerful than flipping a coin.

Clearly, the narrower are the regions Φ and Θ, the higher is the power of the test. Hence, values of $b_{k,0.05}$ close to zero either support the no-connection hypothesis ($b_k = 0$) or suggest that $|b_k|$ is close to zero. Values of $b_{k,0.50}$ that differ greatly from zero suggest that the null of no connection and its alternative are equally likely.

Tables 4 to 9 indicate that, in most cases, the Fisher–Seater test has low power against alternatives. That is, not rejecting the null of no connection is mainly a result of low power. For instance, in Table 4, at $k = 30$ for the monetary fundamental ($f_t = (m_t - m_t^*) - (y_t - y_t^*)$), the estimates $\hat{b}_{30} = 0.012$, which are consistent with not rejecting the null of no connection at the 5% level, coincide with values of $b_{30,0.50} = 0.238$ and $b_{30,0.05} = 0.437$, respectively. This implies that $b_{30} \in [-0.437, 0.437]$ at the 5% level, but because of the low power of the test, there is no evidence against any potential true value of $b_k \in [-0.238, 0.238]$. If we interpret the test results as *inconclusive* when $b_{k,0.50}$ exceeds 0.10, Tables 4 to 9 show that for all $7 \times 6 \times 6 = 252$ cases (seven fundamentals, six horizons, and six countries), 187 (comprising 132 rejections of $b_k = 0$ and 55 that are inconclusive) of the 252 cases suggest a long-run connection between exchange rates and fundamentals.

That is, 74% of the tests strongly or weakly support exchange rate models.

Clearly, because the Fisher–Seater test may fail to detect important deviations from the null of no connection between exchange rates and fundamentals, the results of the previous section probably underestimate the explanatory power of exchange rate models. Hence, "exchange rate models are not as bad as you think" (Engel et al. (2007)).

6 Concluding Remarks

Long-horizon regression tests are widely used to investigate the relationship between exchange rates and fundamentals. However, the empirical evidence on this long-run relationship seems mixed. For instance, using out-of-sample prediction tests and random-walk models as competing models, Mark (1995) and Mark and Sul (2001) show that predictability is higher over longer horizons, whereas Kilian (1999) and Molodtsova and Papell (2007) find no evidence of better prediction over longer horizons.

Following the suggestion of Engel et al. (2007) that judging exchange rate models on the basis of their out-of-sample predictive power relative to a random-walk model is unreliable, we refrain from adding one more horse to the race against a random walk. Instead, to study the long-run connection between nominal exchange rates and fundamentals, we applied an alternative long-horizon regression approach proposed by Fisher and Seater (1993).

Our in-sample results from the Fisher–Seater test confirmed the effectiveness of exchange rate models in predicting exchange rates. In particular, the Taylor-rule model outperformed other conventional models in terms of providing the strongest supporting evidence for a long-run connection between exchange rates and fundamentals.

We used the inverse power functions proposed by Andrews (1989) to investigate the power of the Fisher–Seater test. We showed that the Fisher–Seater test may fail to detect important deviations from the null of no connection between exchange rates and fundamentals. This implies that our results underestimate the ability of exchange rate models to predict exchange rates. Hence, they provide more support for exchange rate models. That is, exchange rate models are not as 'bad' as they are thought to be.

Table 1: Data

Data	Code in Datastream
M2	USOCFMONB, JPM2CDF.B, UKOCM2MNB, BDM2C...B
	ITM2....A, CNOCM2MNB, FRM2....A,
CPI	USI64...F, JPI64...F, UKI64...F, ITI64...F, CNI64...F, FRI64...F
	13464.D.ZF...(1972Q1-1991Q4)
	13464...ZF...(1991Q1-2005Q2) for Germany (IFS)
Real GDP	USOCFGDPD, JPOCFGDPD, UKOCFGDPD, BDOCFGDPD
	ITOCFGDPD, CNOCFGDPD, FROCFGDPD,
Interest Rates	CNOCFISTR, FROCFISTR, BDOCFISTR, ITOCFISTR
	JPOCFISTR, UKOCFISTR, USOCFISTR
Terms of Trade	CNQ..NEUE, FROL0808H, BDTOTPRCF, ITOL0808H
	JPTOTPRCF, UKTOTPRCF, USTOTPRCF
Current Account	BDI78ALDA, CNI78ALDA, FRI78ALDA, ITI78ALDA
	JPI78ALDA, UKI78ALDA, USI78ALDA
Nominal Exchange	156..AE.ZF..., 132..AE.ZF..., 134..AE.ZF..., 136..AE.ZF...
Rates (IFS)	158..AE.ZF..., 112..AG.ZF...

Figure 1: Plots of the b_k coefficient and the 95% confidence interval: $f_t = (m_t - m_t^*) - (y_t - y_t^*)$.

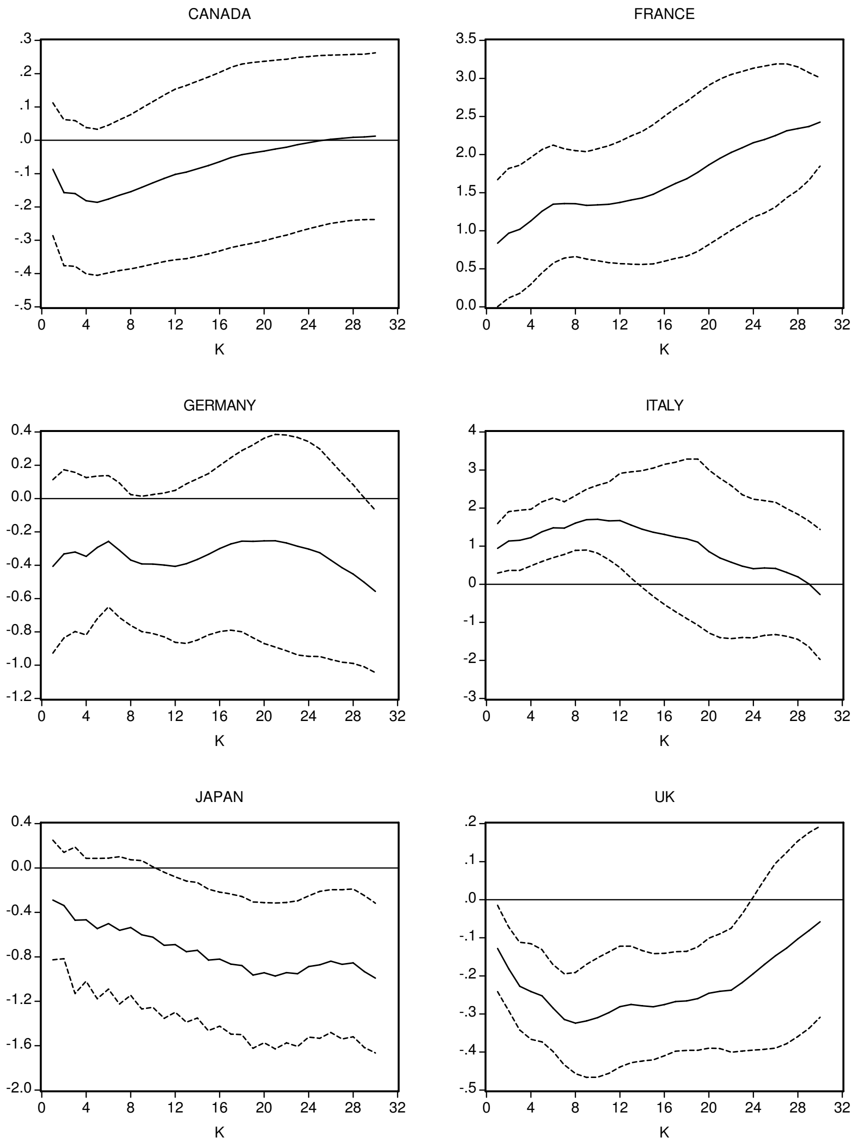

Figure 2: Plots of the b_k coefficient and the 95% confidence interval: $f_t = m_t - m_t^*$.

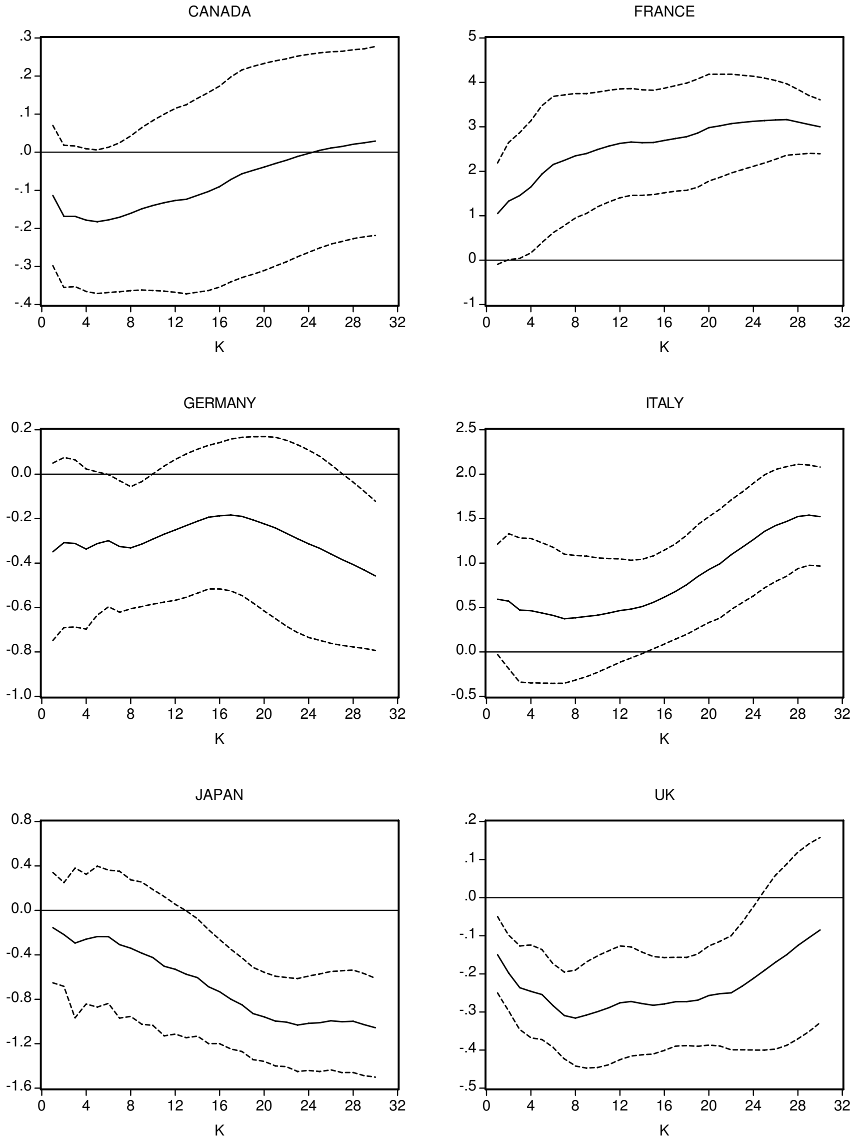

Figure 3: Plots of the b_k coefficient and the 95% confidence interval: $f_t = p_t - p_t^*$.

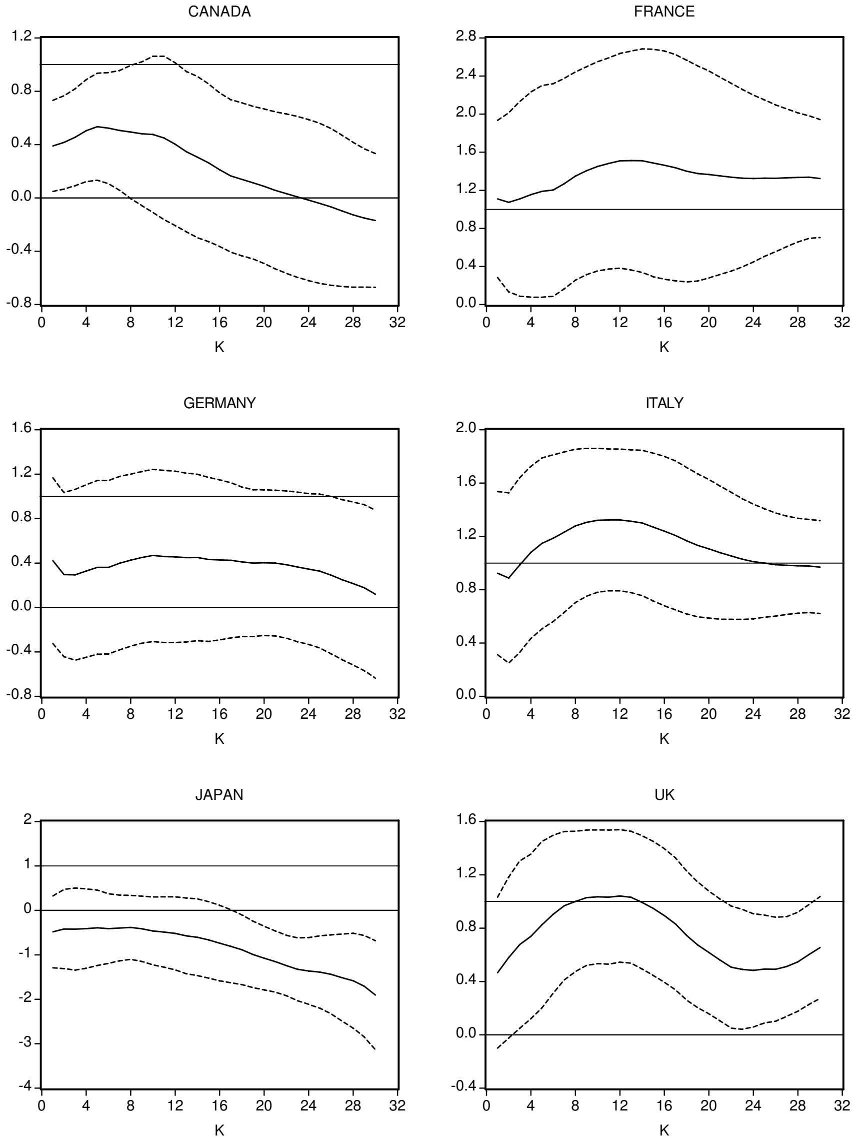

Figure 4: Plots of the b_k coefficient and the 95% confidence interval: $f_t = 1.5(\pi_t - \pi_t^*) + 0.1(y_t - y_t^*) + 0.1(s_t + p_t^* - p_t) + s_t$.

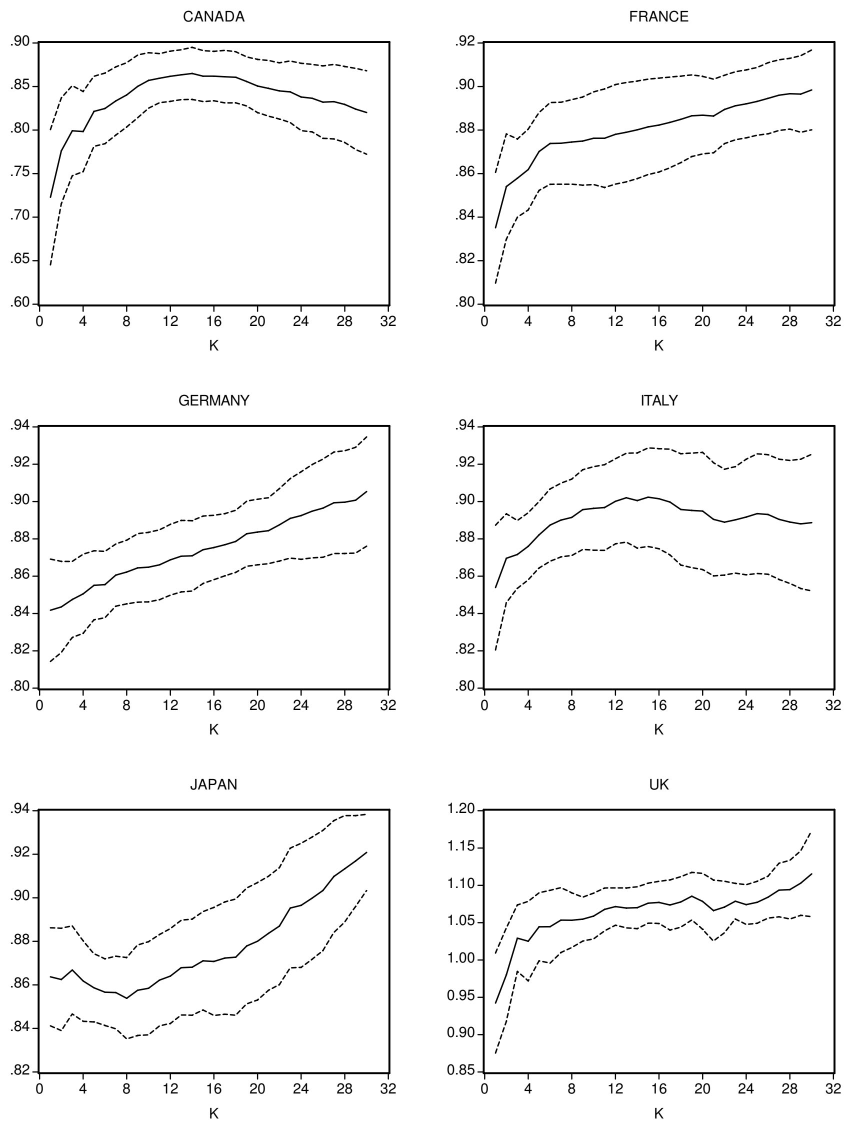

Figure 5: Plots of the b_k coefficient and the 95% confidence interval: $f_t = i - i^* + s$.

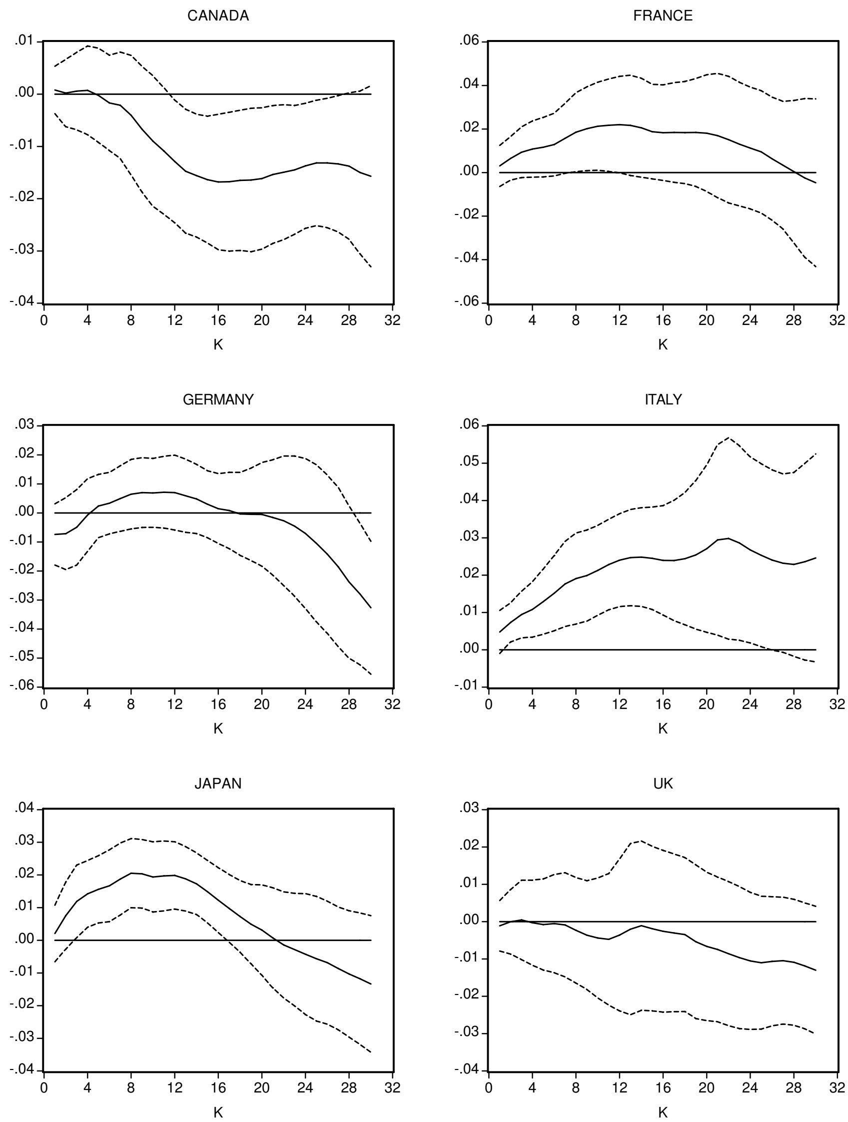

Figure 6: Plots of the b_k coefficient and the 95% confidence interval: $f_t = TOT_t - TOT_t^*$.

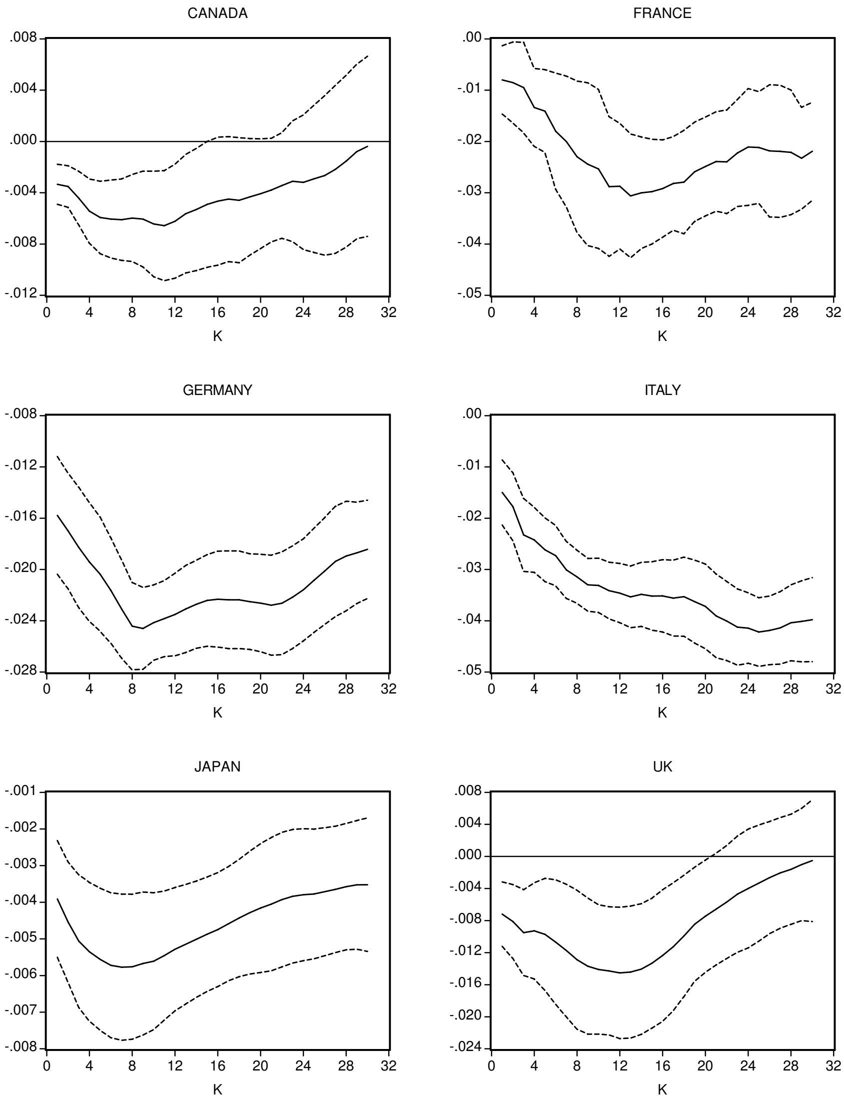

Figure 7: Plots of the b_k coefficient and the 95% confidence interval: $f_t = CA_t - CA_t^*$.

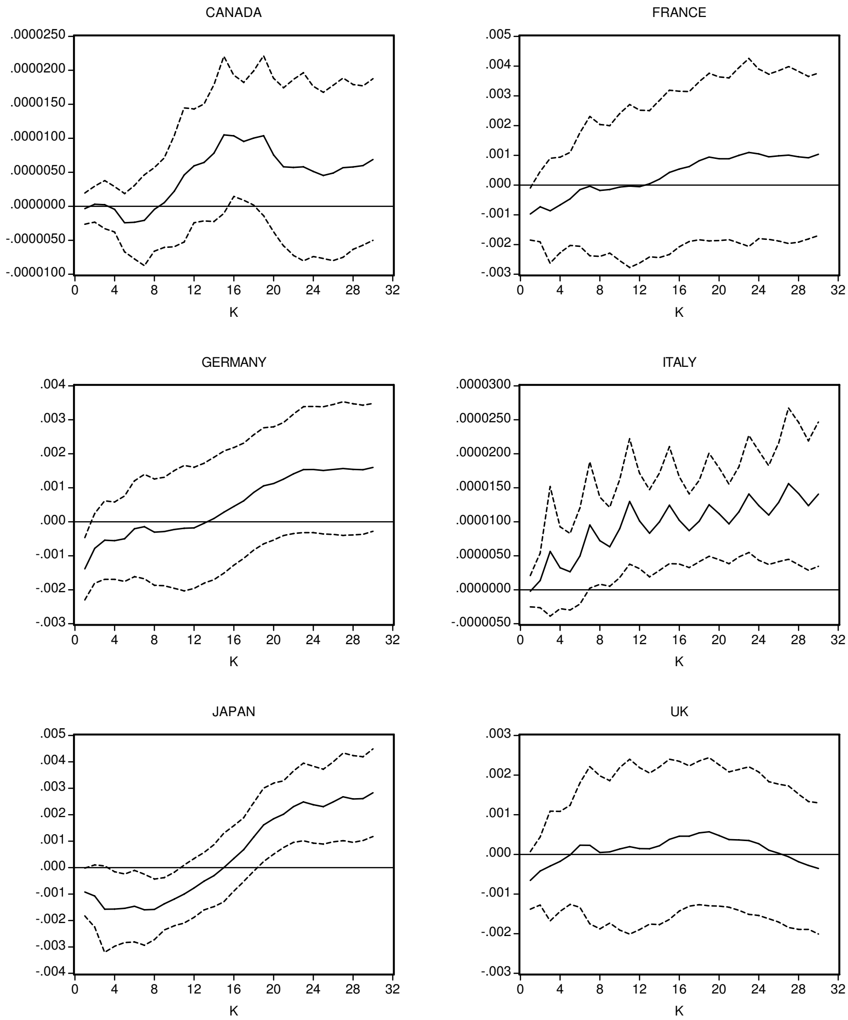

Table 2: Augmented Dickey–Fuller Test

Series in Level	Canada	France	Germany	Italy	Japan	U.K.
S	-1.88	-2.56	-3.18	-2.44	-2.32	-3.36
$p - p^*$	-0.60	-0.22	-2.67	0.34	-3.07	-1.40
$CA - CA^*$	-0.25	0.29	0.85	-2.51	0.12	0.76
$m - m^*$	-2.04	-2.68	-1.55	-2.71	-0.41	-1.82
$TOT - TOT^*$	-1.23	-0.29	-2.54	-2.51	-2.24	-1.93
$(m - m^*) - (y - y^*)$	-1.60	-2.24	-1.71	-2.24	-1.20	-1.81
$1.5(\pi - \pi^*) + 0.1(y - y^*) + 0.1q + s$	-1.81	-2.68	-3.11	-3.16	-2.45	-3.95
$i - i^* + s$	-2.77	-3.38	-4.29	-3.69	-2.58	-4.36

Series in first difference	Canada	France	Germany	Italy	Japan	U.K.
S	-3.34	-4.63	-4.53	-10.30	-5.61	-3.50
$p - p^*$	-3.98	-4.04	-3.53	-4.07	-3.80	-6.30
$CA - CA^*$	-4.37	-2.31	-2.20	-5.21	-2.17	-2.06
$m - m^*$	-3.38	-5.28	-5.16	-4.64	-3.03	-11.78
$TOT - TOT^*$	-10.31	-7.46	-8.83	-5.47	-7.18	-6.26
$(m - m^*) - (y - y^*)$	-4.32	-5.99	-4.58	-4.92	-3.83	-12.08
$1.5(\pi - \pi^*) + 0.1(y - y^*) + 0.1q + s$	-3.64	-5.79	-5.80	-4.64	-5.66	-6.64
$i - i^* + s$	-5.63	-9.35	-4.72	-7.95	-6.01	-5.30

Note: $q = s + p^* - p$. The null hypothesis is that the series has a unit root. An intercept and time trend are included in the testing equation. Lag length is selected by using the Akaike Information Criterion. Critical values for the tests are -3.992156 (1%), -3.426433 (5%), and -3.136443 (10%).

Table 3: Phillips–Perron Test

Series in Level	Canada	France	Germany	Italy	Japan	U.K.
S	-0.94	-2.07	-2.81	-1.75	-2.15	-2.60
$p - p^*$	-0.72	0.06	-2.28	0.84	-1.94	-0.11
$CA - CA^*$	-2.37	0.30	1.05	-5.66	0.14	0.65
$m - m^*$	-1.30	-2.30	-1.49	-1.75	-1.53	-2.15
$TOT - TOT^*$	-1.23	-1.38	-2.35	-2.47	-2.04	-2.37
$(m - m^*) - (y - y^*)$	-0.73	-1.64	-1.51	-1.64	-2.66	-2.06
$1.5(\pi - \pi^*) + 0.1(y - y^*) + 0.1q + s$	-1.26	-2.14	-2.96	-2.62	-2.31	-2.71
$i - i^* + s$	-2.65	-3.08	-2.52	-2.60	-2.99	-3.48
Series in first difference	Canada	France	Germany	Italy	Japan	U.K.
S	-11.13	-10.31	-10.85	-10.38	-10.27	-10.02
$p - p^*$	-9.30	-9.05	-8.78	-7.61	-10.10	-10.70
$CA - CA^*$	-15.38	-11.67	-12.20	-18.48	-10.29	-13.60
$m - m^*$	-7.00	-5.24	-5.09	-8.84	-17.79	-11.84
$TOT - TOT^*$	-10.27	-11.23	-8.88	-11.16	-7.13	-10.02
$(m - m^*) - (y - y^*)$	-9.61	-6.18	-7.69	-8.17	-17.68	-12.08
$1.5(\pi - \pi^*) + 0.1(y - y^*) + 0.1q + s$	-12.70	-10.66	-11.04	-9.02	-10.54	-11.68
$i - i^* + s$	-11.55	-9.26	-11.67	-8.14	-8.20	-11.08

Note: $q = s + p^* - p$. The null hypothesis is that the series has a unit root. An intercept and time trend are included in the testing equation. The HAC variance–covariance matrix is estimated by using the Newey–West method based on the Bartlett kernel function. Critical values for the tests are -3.992156 (1%), -3.426433 (5%), and -3.136443 (10%).

Table 4: Selected estimates of b_k, and Andrews (1989)'s IPF bounds: Canada

Fundamentals	k=5	k=10	k=15	k=20	k=25	k=30
$p - p^*$	0.534**	0.476	0.264	0.087	-0.041	-0.169
$b_{k,0.05}$	–	1.197	1.163	1.089	1.093	0.877
$b_{k,0.50}$	–	0.650	0.632	0.597	0.594	0.477
$CA - CA^*$	-2.42E-06	-2.22E-06	1.05E-05	7.54E-05	4.53E-06	6.88E-06
$b_{k,0.05}$	4.93E-06	9.06E-06	1.23E-05	1.16E-05	1.20E-05	1.12E-05
$b_{k,0.50}$	9.06E-06	1.67E-05	2.26E-05	2.13E-05	2.21E-05	2.06E-05
$m - m^*$	-0.183	-0.140	-0.103	-0.039	0.005	0.029
$b_{k,0.05}$	0.399	0.456	0.511	0.515	0.466	0.434
$b_{k,0.50}$	0.217	0.248	0.278	0.280	0.254	0.236
$TOT - TOT^*$	-0.006**	-0.006**	-0.005**	-0.004	-0.003	0.000
$b_{k,0.05}$	–	–	–	0.008	0.010	0.012
$b_{k,0.50}$	–	–	–	0.004	0.006	0.007
$(m - m^*) - (y - y^*)$	-0.186	-0.128	-0.076	-0.032	-0.001	0.012
$b_{k,0.05}$	0.464	0.498	0.522	0.510	0.466	0.437
$b_{k,0.50}$	0.252	0.271	0.284	0.277	0.253	0.238
$1.5(\pi - \pi^*) + 0.1(y - y^*)$	0.821**	0.857**	0.862**	0.851**	0.837**	0.820**
$+0.1(s + p^* - p) + s$						
$b_{k,0.05}$	–	–	–	–	–	–
$b_{k,0.50}$	–	–	–	–	–	–
$(i - i^*) + s$	-0.000	-0.009	-0.016**	-0.016**	-0.013**	-0.015
$b_{k,0.05}$	0.019	0.025	–	–	–	0.030
$b_{k,0.50}$	0.010	0.014	–	–	–	0.016

Rejection of $b_k = 0$ at the 5% asymptotic level is indicated by **

Table 5: Selected estimates of b_k, and Andrews (1989)'s IPF bounds: France

Fundamentals	k=5	k=10	k=15	k=20	k=25	k=30
$p - p^*$	1.189**	1.451**	1.486**	1.366**	1.328**	1.324**
$b_{k,0.05}$	–	–	–	–	–	–
$b_{k,0.50}$	–	–	–	–	–	–
$CA - CA^*$	0.000	-6.86E-05	0.000	0.000	0.001	0.001
$b_{k,0.05}$	0.003	0.005	0.005	0.005	0.005	0.005
$b_{k,0.50}$	0.002	0.003	0.003	0.003	0.003	0.003
$m - m^*$	1.937**	2.490**	2.647**	2.981**	3.141**	3.000**
$b_{k,0.05}$	–	–	–	–	–	–
$b_{k,0.50}$	–	–	–	–	–	–
$TOT - TOT^*$	-0.014	-0.025	-0.030	-0.025	-0.021	-0.022
$b_{k,0.05}$	0.016	0.029	0.018	0.015	0.016	0.012
$b_{k,0.50}$	0.009	0.016	0.010	0.008	0.008	0.007
$(m - m^*) - (y - y^*)$	1.255**	1.340**	1.478**	1.864**	2.197**	2.427**
$b_{k,0.05}$	–	–	–	–	–	–
$b_{k,0.50}$	–	–	–	–	–	–
$1.5(\pi - \pi^*) + 0.1(y - y^*)$	0.870**	0.876**	0.881**	0.887**	0.893**	0.898**
$+0.1(s + p^* - p) + s$						
$b_{k,0.05}$	–	–	–	–	–	–
$b_{k,0.50}$	–	–	–	–	–	–
$(i - i^*) + s$	0.012	0.021**	0.019	0.018	0.009	-0.005
$b_{k,0.05}$	0.029	–	0.043	0.051	0.051	0.067
$b_{k,0.50}$	0.016	–	0.023	0.027	0.027	0.036

Rejection of $b_k = 0$ at the 5% asymptotic level is indicated by **

Table 6: Selected estimates of b_k, and Andrews (1989)'s IPF bounds: Germany

Fundamentals	k=5	k=10	k=15	k=20	k=25	k=30
$p - p^*$	0.361	0.468	0.433	0.403	0.327	0.119
$b_{k,0.05}$	1.655	1.582	1.452	1.241	1.260	1.321
$b_{k,0.50}$	0.488	0.454	0.503	0.609	0.585	0.434
$CA - CA^*$	0.000	0.000	0.000	0.001	0.002	0.002
$b_{k,0.05}$	0.003	0.004	0.004	0.003	0.003	0.003
$b_{k,0.50}$	0.001	0.002	0.002	0.002	0.002	0.002
$m - m^*$	-0.313	-0.292	-0.194	-0.224	-0.334	-0.458
$b_{k,0.05}$	0.682	0.598	0.636	0.743	0.755	0.587
$b_{k,0.50}$	0.371	0.325	0.346	0.404	0.410	0.319
$TOT - TOT^*$	-0.020**	-0.024**	-0.022**	-0.023**	-0.021**	-0.018**
$b_{k,0.05}$	-	-	-	-	-	-
$b_{k,0.50}$	-	-	-	-	-	-
$(m - m^*) - (y - y^*)$	-0.293	-0.393	-0.334	-0.254	-0.325	-0.557
$b_{k,0.05}$	0.898	0.836	0.925	1.121	1.076	0.798
$b_{k,0.50}$	0.488	0.454	0.503	0.609	0.585	0.434
$1.5(\pi - \pi^*) + 0.1(y - y^*)$ $+0.1(s + p^* - p) + s$	0.855**	0.865**	0.874**	0.884**	0.895**	0.905**
$b_{k,0.05}$	-	-	-	-	-	-
$b_{k,0.50}$	-	-	-	-	-	-
$(i - i^*) + s$	0.002	0.007	0.003	-0.000	-0.010	-0.033**
$b_{k,0.05}$	0.023	0.024	0.023	0.034	0.049	-
$b_{k,0.50}$	0.013	0.013	0.012	0.018	0.027	-

Rejection of $b_k = 0$ at the 5% asymptotic level is indicated by **

Table 7: Selected estimates of b_k, and Andrews (1989)'s IPF bounds: Italy

Fundamentals	k=5	k=10	k=15	k=20	k=25	k=30
$p-p^*$	1.147**	1.320**	1.270**	1.106**	1.000**	0.970**
$b_{k,0.05}$	–	–	–	–	–	–
$b_{k,0.50}$	–	–	–	–	–	–
$CA-CA^*$	2.64E-06	9.00E-06	1.24E-05**	1.12E-05**	1.10E-05**	1.41E-05**
$b_{k,0.05}$	1.19E-05	1.47E-05	–	–	–	–
$b_{k,0.50}$	6.46E-06	7.99E-06	–	–	–	–
$m-m^*$	0.438	0.413	0.557	0.927	1.356	1.522
$b_{k,0.05}$	1.667	1.306	1.020	1.114	1.138	0.954
$b_{k,0.50}$	0.906	0.710	0.555	0.606	0.619	0.519
$TOT-TOT^*$	-0.026**	-0.033**	-0.035**	-0.037**	-0.042**	-0.040**
$b_{k,0.05}$	–	–	–	–	–	–
$b_{k,0.50}$	–	–	–	–	–	–
$(m-m^*)-(y-y^*)$	1.380**	1.707**	1.366**	0.860**	0.428**	-0.269**
$b_{k,0.05}$	–	–	–	–	–	–
$b_{k,0.50}$	–	–	–	–	–	–
$1.5(\pi-\pi^*)+0.1(y-y^*)$ $+0.1(s+p^*-p)+s$	0.882**	0.896**	0.902**	0.895**	0.894**	0.889**
$b_{k,0.05}$	–	–	–	–	–	–
$b_{k,0.50}$	–	–	–	–	–	–
$(i-i^*)+s$	0.013	0.021	0.025	0.027	0.025	0.025**
$b_{k,0.05}$	–	–	–	–	–	0.049
$b_{k,0.50}$	–	–	–	–	–	0.026

Rejection of $b_k = 0$ at the 5% asymptotic level is indicated by **

Table 8: Selected estimates of b_k, and Andrews (1989)'s IPF bounds: Japan

Fundamentals	k=5	k=10	k=15	k=20	k=25	k=30
$p - p^*$	-0.392	-0.465	-0.665	-1.075	-1.388	-1.905
$b_{k,0.05}$	1.794	1.558	1.691	1.356	1.478	2.138
$b_{k,0.50}$	0.975	0.847	0.920	0.737	0.803	1.163
$CA - CA^*$	-0.002**	-0.001**	4.00E-06	0.002**	0.002**	0.003**
$b_{k,0.05}$	-	-	0.003	-	-	-
$b_{k,0.50}$	-	-	0.001	-	-	-
$m - m^*$	-0.236	-0.424	-0.687	-0.958	-1.011	-1.056
$b_{k,0.05}$	1.342	1.248	1.008	0.759	0.800	0.778
$b_{k,0.50}$	0.730	0.679	0.548	0.413	0.435	0.423
$TOT - TOT^*$	-0.006**	-0.006**	-0.005**	-0.004**	-0.004**	-0.004**
$b_{k,0.05}$	-	-	-	-	-	-
$b_{k,0.50}$	-	-	-	-	-	-
$(m - m^*) - (y - y^*)$	-0.546	-0.623	-0.829	-0.943	-0.872	-0.992
$b_{k,0.05}$	1.337	1.292	1.253	1.195	1.205	1.179
$b_{k,0.50}$	0.727	0.702	0.681	0.650	0.650	0.641
$1.5(\pi - \pi^*) + 0.1(y - y^*)$ $+0.1(s + p^* - p) + s$	0.859**	0.858**	0.871**	0.880**	0.900**	0.921**
$b_{k,0.05}$	-	-	-	-	-	-
$b_{k,0.50}$	-	-	-	-	-	-
$(i - i^*) + s$	0.017**	0.019**	0.015**	0.003	-0.005	-0.013
$b_{k,0.05}$	-	-	-	0.026	0.035	0.037
$b_{k,0.50}$	-	-	-	0.014	0.019	0.020

Rejection of $b_k = 0$ at the 5% asymptotic level is indicated by **

Table 9: Selected estimates of b_k, and Andrews (1989)'s IPF bounds: the U.K.

Fundamentals	k=5	k=10	k=15	k=20	k=25	k=30
$p-p^*$	0.826**	1.035**	0.948**	0.618**	0.493**	0.655**
$b_{k,0.05}$	–	–	–	–	–	–
$b_{k,0.50}$	–	–	–	–	–	–
$CA-CA^*$	-1.26E-05	0.000	0.000	0.000	0.000	0.000
$b_{k,0.05}$	0.003	0.004	0.004	0.003	0.003	0.003
$b_{k,0.50}$	0.001	0.002	0.002	0.002	0.002	0.002
$m-m^*$	-0.254**	-0.299**	-0.283**	-0.257**	-0.191**	-0.085**
$b_{k,0.05}$	–	–	–	–	–	–
$b_{k,0.50}$	–	–	–	–	–	–
$TOT-TOT^*$	-0.010**	-0.014**	-0.013**	-0.007**	-0.003	-0.001
$b_{k,0.05}$	–	–	–	–	0.013	0.013
$b_{k,0.50}$	–	–	–	–	0.007	0.007
$(m-m^*)-(y-y^*)$	-0.252**	-0.310**	-0.281**	-0.246**	-0.170**	-0.058**
$b_{k,0.05}$	–	–	–	–	–	–
$b_{k,0.50}$	–	–	–	–	–	–
$1.5(\pi-\pi^*)+0.1(y-y^*)$ $+0.1(s+p^*-p)+s$	1.044**	1.059**	1.076**	1.078**	1.077**	1.115**
$b_{k,0.05}$	–	–	–	–	–	–
$b_{k,0.50}$	–	–	–	–	–	–
$(i-i^*)+s$	-0.000	-0.004	-0.002	-0.007	-0.011	-0.013
$b_{k,0.05}$	0.026	0.033	0.043	0.038	0.032	0.030
$b_{k,0.50}$	0.014	0.018	0.024	0.020	0.018	0.016

Rejection of $b_k = 0$ at the 5% asymptotic level is indicated by **

References

Andrews, Donald W K (1989), "Power in econometric applications", *Econometrica*, 57(5), 1059–1090.

Berkowitz, Jeremy and Giorgianni, Lorenzo (2001), "Long-horizon exchange rate predictability?", *Review of Economics and Statistics*, 83(1), 81–91.

Chen, Jian and Mark, Nelson C (1996), "Alternative long-horizon exchange-rate predictors", *International Journal of Finance and Economics*, 1(4), 229–250.

Chinn, Menzie D and Meese, Richard A (1995), "Banking on currency forecasts: How predictable is change in money?", *Journal of International Economics*, 38(1-2), 161–178.

Clark, Todd E. and West, Kenneth D. (2007), "Approximately normal tests for equal predictive accuracy in nested models", *Journal of Econometrics*, 138(1), 291–311.

Coe, Patrick J. and Nason, James M. (2003), "The long-horizon regression approach to monetary neutrality: How should the evidence be interpreted?", *Economics Letters*, 78(3), 351–356.

——— (2004), "Long-run monetary neutrality and long-horizon regressions", *Journal of Applied Econometrics*, 19(3), 355–373.

Engel, Charles, Mark, Nelson C., and West, Kenneth D. (2007), "Exchange rate models are not as bad as you think", *Working Paper, University of Wisconsin.*

Engel, Charles and West, Kenneth D. (2005), "Exchange rates and fundamentals", *Journal of Political Economy*, 113, 485–517.

Fisher, Mark E. and Seater, John J. (1993), "Long-run neutrality and superneutrality in an arima framework", *American Economic Review*, 83(3), 402–415.

Kilian, Lutz (1999), "Exchange rates and monetary fundamentals: What do we learn from long-horizon regressions?", *Journal of Applied Econometrics*, 14(5), 491–510.

Mark, Nelson C. (1995), “Exchange rates and fundamentals: Evidence on long-horizon predictability”, *American Economic Review*, 85(1), 201–218.

Mark, Nelson C. and Sul, Donggyu (2001), “Nominal exchange rates and monetary fundamentals evidence from a small post-bretton woods panel”, *Journal of International Economics*, 53, 29–52.

Meese, Richard and Rogoff, Kenneth (1983), “Empirical exchange rate models of the 1970’s: Do they fit out of sample?”, *Journal of International Economics*, 14, 3–24.

Molodtsova, Tanya and Papell, David H. (2007), “Out-of-sample exchange rate predictability with taylor rule fundamentals”, *Working Paper, University of Houston.*

Rapach, David E and Wohar, Mark E. (2002), “Testing the monetary model of exchange rate determination: New evidence from a century of data”, *Journal of International Economics*, 58(2), 359–385.

——— (2004), “Testing the monetary model of exchange rate determination: A closer look at panels”, *Journal of International Money and Finance*, 23(6), 867–895.

Serletis, Apostolos and Gogas, Periklis (2004), “Long-horizon regression tests of the theory of purchasing power parity”, *Journal of Banking and Finance*, 28, 1961–1985.

Wallace, Frederick H. and Shelley, Gary L (2006), “An alternative test of purchasing power parity”, *Economics Letters*, 92(2), 177–183.

赴國外研究心得報告

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計畫名稱	匯率與經濟基本面: 長區間檢定
出國人員姓名 服務機關及職稱	陳旭昇/台大經濟系/助理教授
出國時間地點	2007 年 5 月 24 日-- 5 月 27 日/ 日本
國外研究機構	青山學院大學

工作記要：

本人於 2007 年 5 月 24 日-- 5 月 27 日赴日本青山學院大學與論文合作人 Kenshi Taketa 教授商討合作之論文。除了 24 與 27 日為赴日旅行與返台旅行，我與 Taketa 教授於 25 日及 26 日在青山學院大學討論已完成之論文：

（1）Exchange Rate Pass-Through and Import Prices across Different Episodes

（2）An Assessment of Weymark's Measures of Exchange Market Intervention: The Case of Japan

（3）Cross-Country Assessment of Weymark's Measures of Exchange Market Intervention

之修改與寄送期刊審查事宜，並商討下一步合作之主題。