An Empirical Investigation of Exchange Rates and The Term Structure of Interest Rates

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Abstract

In this paper, we investigate the relationship between exchange rates and the term structure of interest rates, using cointegrating regressions for selected developed countries. We find a stylized fact that the long-run effect of the short-term real interest rate differential on the real exchange rate is the opposite from what the standard exchange rate model predicts when the effect of the long-term real interest rate differential is controlled: when the domestic short-term real interest rate rises relative to the foreign short-term interest rate while the long-term real interest differential stays the same, the domestic currency tends to depreciate in the long-run. We present an exchange rate model, which is consistent with this stylized fact.
1. Introduction

In this paper, we investigate the long-run relationship between exchange rates and the term structure of interest rates, using cointegrating regressions for selected developed countries. We find a stylized fact that the long-run effect of the short-term real interest rate differential on the real exchange rate is the opposite from what the standard exchange rate model predicts when the effect of the long-term real interest rate differential is controlled: when the domestic short-term real interest rate rises relative to the foreign short-term interest rate while the long-term real interest differential stays the same, the domestic currency tends to depreciate in the long-run.

This paper extends Meese and Rogoff's (1988) cointegrating regression by regressing the log of the real exchange rate (expressed in terms of a basket of domestic goods) on both the short-term real interest rate differential and the long-term real interest rate differential. For this regression to be a cointegrating regression, the real interest rate differentials need to be unit root nonstationary. It seems, however, more reasonable to assume that they are stationary but have autoregressive roots which are near one. In that case, the point estimates can still be expected to be fairly precise even though standard errors and p-values of test statistics we calculate are imprecise (see Elliott (1998)).

We obtain a negative coefficient for the long-term interest rate differential, which implies that the foreign currency tends to depreciate when the domestic long-term real interest rate rises. The negative coefficient is predicted by standard exchange rate models which make two assumptions. The first assumption is that the uncovered interest parity
(UIP) condition holds and the second assumption is that the real exchange rate is expected to revert toward a long-run equilibrium value. We typically obtain a positive coefficient for the short-term interest rate differential for most of exchange rates we examined. Most of these coefficients are statistically significant. Ogaki and Santaella (1999) obtain similar results for Mexico.

There are many empirical studies in which exchange rates are regressed on interest rate differentials. However, there has been little empirical work on exchange rates and the term structure of interest rates. Our study is the first in which real exchange rates are regressed on both real short-term and long-term interest rate differentials. Many studies use either short-term or long-term interest rate differential as a regressor. Some use the real short-term interest rate differential and the nominal long-term interest rate differential (see, e.g., Frankel (1979) and Edison (1985)) in an exchange rate regression. Boughton (1988) uses the nominal short-term interest rate and the real long-term interest rate.

The work in this paper is motivated by stylized facts regarding the uncovered interest parity for short-term and long-term interest rates. For short-term interest rates and forward exchange rates, the uncovered interest parity is typically rejected (see, e.g., Hodrick (1988) and Engel (1997) for recent surveys). As Engel (1997) emphasizes, one form of the rejection found in many recent papers is that regressions of the future depreciation on the current forward premium (which is equal to the short-term interest rate differential under the covered interest parity) yield negative estimates of the slope coefficient. This is called the forward premium anomaly (also see Backus, Foresi, and Telmer (1998) for a recent
For long-term bonds, more favorable evidence for the uncovered interest parity has been found. Direct evidence is found in recent papers by Meredith and Chinn (1988) and Alexius (1998, 1999). Meredith and Chinn (1988) and Alexius (1998) find that regressions of the future depreciation over a long-horizon on the current long-term interest rate differential typically yield significantly positive estimates of the slope coefficient. Alexius (1999) finds similar results for returns on long-term bonds over short investment horizons.

Indirect evidence has been found in the form of more favorable evidence for standard exchange rate models which assume the uncovered interest parity and the long-run purchasing power parity (see, e.g., Meese and Rogoff (1988), Edison and Pauls (1993) and Baxter (1994)) with long-term interest rate differentials than with short-term interest rate differentials. Similarly, implications of standard exchange rate models hold better in long-horizon data than in short-horizon data (see, e.g., Mark (1995)).

It is difficult to find an economic explanation for the forward premium anomaly because standard consumption-based asset pricing model with risk averse investors cannot explain it (see, e.g., Mark and Wu (1998)). A model which is consistent with the stylized fact without relying on high risk aversion coefficients is constructed in the companion paper, Ogaki (1999).\footnote{Alvarez, Atkeson, and Kehoe (1999) constructs a model of segmented asset markets which can be consistent with the forward premium anomaly. Ogaki's (1999) model gives an alternative explanation which is not based on transactions costs. McCullum (1994) and Meredith and Chinn (1998) provide an explanation for the forward premium anomaly based on policy reactions. Ogaki's (1999) explanation is complementary to theirs because their models require an error term which is correlated with the current interest rate differential in order for the exchange rate to deviate from the level implied the uncovered interest rate parity.}
The model is a three-asset model with domestic short-term bonds, domestic long-term bonds, and foreign bonds. In the model, foreign bonds and domestic long-term bonds are stronger substitutes than foreign and domestic short-term bonds. An important assumption of the model is that the investment horizon of the economic agents is short-run. This assumption may be a good approximation of the reality if exchange rates and long-term bond prices are mainly determined by the traders with short-run investment horizons.

The intuition behind this result arises from a consideration of the risk structure of interest rates under the assumption of risk aversion and Ogaki's (1990) concept of indirect complementarity. If domestic short-term interest rates unexpectedly rise, holders of domestic long-term bonds suffer a capital loss. If the domestic currency appreciates as a result of this increase, then holders of foreign bonds also suffer a capital loss. Therefore, as long as an increase in short-term interest rates is associated with an appreciation of the domestic currency, risk averse agents have an incentive to avoid holding both domestic long-term bonds and foreign bonds. Hence these two assets are likely to be strong substitutes. Since domestic short-term and long-term bonds are also likely to be strong substitutes, and since a substitute of a substitute is an indirect complement, domestic short-term bonds and foreign bonds must be strong indirect complements.

For example, suppose that the domestic short-term interest rate rises. This will have two effects, each of which will shift the demand for foreign bonds in the opposite directions. The first effect, which arises from the direct substitution between domestic short-term bonds and foreign bonds,
reduces the demand for foreign bonds. The second effect, caused by the indirect complementarity, increases the demand for foreign bonds. This is because the strong substitutability between domestic short-term and long-term bonds implies that the demand for domestic long-term bonds decreases. Since domestic long-term bonds and foreign bonds are strong substitutes, this decrease causes an indirect effect which increases the demand for foreign bonds.\(^2\) Whether domestic short-term bonds and foreign bonds are substitutes or complements depends on which of these effects is stronger. In the rational expectations equilibrium of the model, the indirect complementarity dominates the direct substitutability with some reasonable configurations, so that foreign bonds and domestic short-term bonds are complements.

When the indirect complementarity dominates the direct substitutability, the relationship between exchange rates and interest rate differentials is very complicated. Suppose that the short-term real interest rate in the domestic country rises relative to foreign real short-term interest rates. If the domestic long-term real interest rate also rises relative to their foreign counterparts, then the domestic currency appreciates as predicted by standard models. However, if the domestic long-term real interest rate does not rise, then the domestic currency depreciates. This is because economic agents substitute away from domestic long-term bonds, and this is a factor to increase the demand for foreign bonds. In order to achieve equilibrium for foreign bonds, the domestic currency depreciates now, which is a factor to decrease the demand for

\(^{2}\)This intuitive argument is true even when there are more than three goods (or assets); a substitute of a substitute is always an indirect complement (see Ogaki (1990)).
foreign bonds because the depreciation gives rise to an expectation that the
domestic currency will appreciate in the future.

Our empirical result is consistent with the view that the indirect
complementarity dominates the direct substitutability. This view is in
sharp contrast to the conventional view that short-term capital is more
internationally mobile than long-term capital. The 1960's Operation Twist,
in which the Federal Reserve and the Treasury attempted to raise short-term
interest rates relative to long-term interest rates, was evidently based on
the view. However, empirical work of Fukao and Okubo (1984) suggests that
the relationship between domestic long term interest rates and foreign
interest rates is not limited to that which exists as a result of the
relationship between domestic short term interest rates and foreign interest
rates. Popper (1989) presents empirical evidence that long-term capital is
as internationally mobile as short-term capital.

The rest of the paper is organized as follows. Section 2 describes the
data. Section 3 explains the econometric model used to investigate the
relationship between the real exchange rates and the short-term and long-
term real interest rate differentials. Section 4 explains the econometric
procedure we use for estimation and testing of the econometric model.
Section 5 presents the empirical results. Section 6 describes an exchange
rate model which is consistent with the stylized fact found in Section 5.
Section 7 concludes.

2. The Data

The data are obtained monthly from the CITIBASE, OECD Main Economic
Indicator (MEI), and International Financial Statistics (IFS) over April
1973-April 1995 period. Exchange rates are monthly average of daily rates and are converted to the unit of domestic currency (dollar) per unit foreign currency. The short-term interest rates are three-month T-bills rate for the U.S., Great Britain, and Switzerland, three-month Fibor for the Germany, three-month Gensaki for the Japan, 90 days deposit receipts for the Canada, three-month Fibor for the France, and three-month call money rate for the Italy. The long-term rates are yield on 10+ years Treasury composite for the U.S., yield on 20 years central government bond for the Great Britain, yield on 7-15 years public sector bonds for the Germany, yield on central government bonds for the Japan, yield on federal government bonds (10 years +) for the Canada, yield on public and semi public sector bonds for the France, yield on long-term government bond for the Italy, and yield on government bond for Switzerland. Prices are measured by seasonally unadjusted CPIs.

The log of the real exchange rates are defined by the log of the nominal spot exchange rate (expressed in terms of the domestic currency) plus the log of the foreign price level minus the log of the domestic price level. To construct the *ex ante* short- and long-term real interest rates, r_s and r_l, one needs to compute a measure of expected inflation. We compute one-period-ahead expected inflation rate with three-month realized inflation rate assuming it follows an autoregressive process of order p and then annualize it\(^3\). We use that three-month expected inflation rate as a proxy for short-term expected inflation. We also compute five-year expected inflation as a measure of long-term expected inflation but the results are

\(^3\) We determine the order of AR process using the BIC criterion proposed by Hannan and Rissanen (1982).
very similar to those with three-month expected inflation as a measure of long-term expected inflation. Therefore, we only report the results with three-month expected inflation as a measure of expected short- and long-term inflation in this paper.

The null hypothesis of unit root nonstationarity for these real interest differentials are reported in Table 1. Typically, the null hypothesis is not rejected, even though the test results often depend on the test used as reported in Table 1. Thus these variables may be persistent enough that the unit root nonstationarity is a good approximation.

3. The Econometric Model

In order to investigate the relationship between exchange rates and the term structure of interest rates, we extend Meese and Rogoff's (1988) cointegrating regression by regressing the log of the real exchange rate on both the short-term real interest rate differential and the long-term real interest rate differential. Let \( s_t \) be the log of the real exchange rate, expressed in terms of a basket of domestic good, \( r_{t,1} \) be the domestic real long-term interest rate, and \( r_{t,1}^* \) be the foreign real long-term interest rate, \( r_{s,t} \) be the domestic real short-term interest rate, and \( r_{s,t}^* \) be the foreign real long-term interest rate. Then we consider the following econometric model:

\[
(1) \quad s_t = \mu + \gamma t + \alpha (r_{t,1} - r_{t,1}^*) + \beta (r_{s,t} - r_{s,t}^*) + \nu_t
\]

where we assume that \( \nu_t \) is a stationary random variable and \( s_t, (r_{t,1} - r_{t,1}^*), \) and \( (r_{s,t} - r_{s,t}^*) \) are unit root nonstationary, so that (1) is a cointegrating regression. The assumption that the dominant autoregressive root of each of the real interest rate differentials is not exactly one but
is close to one seems more reasonable. In that case, the point estimates of the regression can still be expected to be fairly precise even though the standard errors and test statistics we calculate are imprecise (see Elliott (1988)).

In (1), \( \gamma \) is zero if the deterministic cointegration restriction is satisfied in the terminology of Ogaki and Park (1998). If \( X_t \) is a vector unit root nonstationary process with drift, \( X_t \) is said to be stochastically cointegrated if \( X_t' \theta \) is trend stationary for a vector of real numbers \( \theta \), and is said to satisfy the deterministic cointegration restriction if \( X_t' \theta \) is stationary. If \( X_t' \theta \) is trend stationary, the vector \( \theta \) (which is called a cointegrating vector) eliminates the stochastic trends in \( X_t \). If the deterministic cointegration restriction is satisfied, the cointegrating vector also eliminates the deterministic trends which arise from the drift.

In (1), \( \alpha \) and \( \beta \) are identified when \( (r_{1,t} - r_{1,t-1}) \) and \( (r_{s,t} - r_{s,t-1}) \) are not stochastically cointegrated. As reported in Byeon (1996), this hypothesis is not rejected for our data. However, because short-term and long-term interest rates tend to move together, they may be close to be stochastically cointegrated, which can cause problems similar to multicollinearity in the stationary case. For this reason, we also try an alternative measure for the short-term interest rate, which we call the normalized short-term interest rate \( (r_{s,r}^N) \). The normalized short-term interest rate is the negative term premium for the long-term bond, and is defined as

\[
(2) \quad r_{s,t}^N = \frac{1}{m} \sum_{k=0}^{m-1} r_{s,t+3k} - r_{L,t+3k}
\]

When the expectation hypothesis of the term structure of interest rates
holds, the short-term and long-term interest rates have a common stochastic trend and are stochastically cointegrated. The normalized short-term interest rate captures the stochastic trend in the short-term interest rate which deviate from the stochastic trend of the long-term interest rate. The economic model in Section 5 also motivates this measure of the short-term.

When we use the normalized short-term interest rate differential, the regression is

\[(1') \quad s_t = \mu + \gamma t + \alpha (r_{t,i} - r_{t,i}^*) + \beta (r_{s,t}^N - r_{s,t}^N) + v_t.\]

4. The Econometric Procedure

This section describes our econometric procedure for the estimation of the cointegrating regression. We employ Park's (1992) Canonical Cointegrating Regression (CCR). The CCR estimator is one of many asymptotically efficient estimators for cointegrated systems. This procedure allows us to test the null hypothesis of stochastic cointegration and the deterministic cointegration restriction.

The CCR procedure requires us to transform the data before running a regression so as to correct for endogeneity and serial correlation. Let \( X_t \) be the 2-dimensional vector of the stochastic regressors: \( X_t = \begin{bmatrix} (r_{t,i} - \bar{r}_{1,i}) \quad (r_{s,i}^N - \bar{r}_{s,i}^N) \end{bmatrix}' \) for (1) and \( X_t = \begin{bmatrix} (r_{t,i} - \bar{r}_{1,i}) \quad (r_{s,i}^N - \bar{r}_{s,i}^N) \end{bmatrix}' \) for (1'). Let \( w_t = \nu_t \Delta X_t \). Define \( \Phi(i) = E(\nu(t)\nu(t-i)') \), \( \Sigma = \Phi(0) \), \( \Gamma = \sum_{i=0}^{\infty} \Phi(i) \), and \( \Omega = \sum_{i=0}^{\infty} \Phi(i) \). Here \( \Omega \) is the long run covariance matrix of \( w_t \).

Partition \( \Omega \) as

\[
(3) \quad \Omega = \begin{pmatrix} \Omega_{11} & \Omega_{12} \\ \Omega_{21} & \Omega_{22} \end{pmatrix} \begin{pmatrix} 1 \\ 2 \end{pmatrix}
\]
and partition \( \Gamma \) conformably. Define

\[
\Omega_{12} = \Omega_{11} - \Omega_{12} \Omega_{22}^{-1} \Omega_{21},
\]

and \( \Gamma_2 = (\Gamma'_{12}, \Gamma'_{22})' \). The CCR procedure assumes that \( \Omega_{22} \) is positive definite.

The OLS estimator for a cointegrating regression is super-consistent in that the estimator converge to \( \beta \) at the rate of \( T \) (sample size) even when \( \Delta X_i \) and \( y_i \) are correlated. The OLS estimator, however, is not asymptotically efficient. Consider transformations

\[
\begin{align*}
\tilde{y}_t^* &= y_t + \Pi^*_y w_t \\
\tilde{X}_t^* &= X_t + \Pi^*_x w_t.
\end{align*}
\]

Because \( w_t \) is stationary, \( \tilde{y}_t^* \) and \( \tilde{X}_t^* \) are cointegrated with the same cointegrating vector \( (1, -\beta) \) as \( y_t \) and \( X_t \) for any \( \Pi_y \) and \( \Pi_x \). The idea of the CCR is to choose \( \Pi_y \) and \( \Pi_x \), so that the OLS estimator is asymptotically efficient when \( \tilde{y}_t^* \) is regressed on \( \tilde{X}_t^* \). \(^4\) This requires

\[
\begin{align*}
\Pi_y &= \Sigma^* \Gamma_2 y + (0 \Omega_{12}^{-1} \Omega_{22}^{-1})' \\
\Pi_x &= \Sigma^* \Gamma_2 x.
\end{align*}
\]

In practice, long-run covariance parameters in these formulas are estimated, and estimated \( \Pi_y \) and \( \Pi_x \) are used to transform \( y_t \) and \( X_t \). As long as these parameters are estimated consistently, the resultant CCR estimator is asymptotically efficient.

The CCR estimators have asymptotic distributions that can essentially

\(^4\)Under general conditions, a sequence of functions \((1/T)^{1/2} w(t)\) converges in distribution to a vector Brownian motion \( B \) with covariance matrix \( \Omega \). The OLS estimator converges in distribution to ???
be considered normal, implying that their standard errors have the usual interpretation. An important property of the CCR procedure is that linear restrictions can be tested by $\chi^2$ tests, which are free from nuisance parameters. We use $\chi^2$ tests in a regression with spurious deterministic trends added to (1) or (1') in order to test for stochastic and deterministic cointegration. For this purpose, the CCR procedure is applied to the regression

$$s_t = \eta_0 + \sum_{i=1}^{q} \eta_i d^i + \gamma X_t + \nu_t$$

Let $H(p,q)$ denote the standard Wald statistic under the hypothesis $\eta_{p+1} = \eta_{p+2} = \ldots = \eta_q = 0$ with the estimate of the variance of $\nu_t$ replaced by $\Omega_{1:2}$ (see Park (1990) for details). Then $H(p,q)$ converges in distribution to a $\chi^2_q$ random variable under the null of cointegration. In particular, the $H(0,1)$ statistic tests the deterministic cointegrating restriction. On the other hand, the $H(1,q)$ statistic tests stochastic cointegration.

5. Cointegrating Regression Results

Table 2 reports the CCR results for (1), and Table 3 reports the results for (1'). If the coefficient of the trend term is not significant

5 The CCR estimators are asymptotically efficient, but there are other asymptotically efficient estimators such as those developed by Stock and Watson (1993) among others. Johansen’s (1988) estimators are often used, but Johansen assumes a Gaussian VAR structure, which is not compatible with our economic model with nonlinear short-run dynamics.

6 We used Ogaki’s (1993) GAUSS CCR Package for the CCR estimations. The CCR procedure requires an estimate of the long run covariance of the disturbances in the system. We used Park and Ogaki’s (1991) method with Andrews and Monahan’s (1992) prewhitened HAC estimator with the QS kernel. VAR of order one was used for prewhitening. We followed footnote 4 of
at even 10 percent significance level, we run the CCR without trend term.

In Table 2, the point estimates for $\alpha$ are significant at least at the 10 percent level in all cases and have the theoretically correct negative sign. Our estimates indicate that high long-term real interest rate in the United States relative to foreign long-term real interest rate causes the real appreciation of the dollar. The estimates for $\beta$ are significant at least at the 10 percent level except U.S./Germany, US/Canada, and U.S./Switzerland. The point estimates for $\beta$ are positive except for U.S./Canada. The $H(1,q)$ statistics test stochastic cointegration and we fail to reject the null of stochastic cointegration at the 1 percent significance level in all cases except for the $H(1,4)$ statistic for U.S./Canada. The $H(0,1)$ statistic tests deterministic cointegration restriction and we cannot reject deterministic cointegration restrictions at the conventional significance levels in U.S./Germany and U.S./Switzerland.

In Table 3, the point estimates for $\alpha$ are significant at least 10 percent level in all cases and have the theoretically correct negative sign. The point estimates for $\beta$ are significant except U.S./Germany and U.S./Switzerland. They are positive except for U.S./Canada and U.S./Switzerland. The $H(1,q)$ statistics test stochastic cointegration and we fail to reject the null of stochastic cointegration at 1 percent significance level in all cases except $H(1,3)$ and $H(1,4)$ statistics for U.S./Canada. The $H(0,1)$ statistic tests deterministic cointegration restriction and we cannot reject deterministic cointegration restrictions at

Andrews and Monahan and the maximum absolute value of the elements of $\Delta$ in their notation was set to 0.99. Andrews's (1991) automatic bandwidth estimator, $S_P$, was constructed from fitting AR(1) to each disturbance. We report third stage CCR estimates and $H(p,q)$ test statistics with the bounds.
even 10 percent significance level in U.S./Germany and U.S./Switzerland. Contrary to many recent papers, this results appear to show that there is the long-run relationship between real exchange rates and real interest rate differentials except U.S./Canada.

6. An Exchange Rate Model

This section explains Ogaki's (1996) partial equilibrium model of exchange rate determination as one model which is consistent with the stylized fact found in motivates our empirical work. For simplicity, the price level is assumed to be constant. Alternatively, all variables can be considered to be measured in real terms. Investors are assumed to live for two periods, and the same number of investors are assumed to be born every period. Let $B_{s,t}$, $B_{l,t}$, and $B_{f,t}$ denote domestic short-term bonds, domestic long-term bonds, and foreign bonds, respectively.

The long-term and short-term bonds are discount bonds which will pay one unit of the domestic currency after two periods and one period, respectively. Let $q_t$ be the price of long-term bonds during period $t$ and $r_t$ be the domestic short-term interest rate. Then the rate of return on holding long-term bonds for one period is

$$r_{L,t} = (1 + r_{t+1})q_t/q_t.$$

Let $R_t$ denote the current long-term interest rate:

$$q_t = 1/(1 + R_t)^2.$$

From (10) and (11), we obtain

$$r_{L,t} = (1 + R_t)^2/(1 + r_{t+1}) - 1 = 2R_t - r_{t+1}.$$ 

Define the risk premium for long-term bonds, $p_{L,t}$, to be the difference
between the expected rate of return for long-term bonds and that for short-term bonds:

\[(13) \quad \rho_{L,t} = E_t(r_{L,t}) - r_t = 2 \left[ R_t - \frac{1}{2} \{ r_t + E_t(r_{L,t+1}) \} \right], \]

where \( E_t \) is the expectation operator conditional on the information set in period \( t \), \( \Omega_t \). We assume that \( \Omega_t \) includes the current and past values of \( r_t \), \( R_t \), \( r^*_t \), and \( s_t \), where \( r^*_t \) is the foreign short-term interest rate and \( s_t \) is the natural log of the exchange rate expressed in terms of the domestic currency.

Let \( r^*_t \) be the foreign short-term interest rate. Then the one-period holding rate of return on foreign bonds in terms of the domestic currency is

\[(14) \quad r_{F,t} = r^*_t + s_{t+1} - s_t.\]

Let the risk premium for the foreign bonds, \( \rho_{F,t} \), be defined as the difference between the expected return on foreign bonds and the domestic short-term interest rate:

\[(15) \quad \rho_{F,t} = r^*_t + E_t(s_{t+1}) \cdot s_t - r_t.\]

The representative investor in period \( t \) maximizes his expected utility

\[(16) \quad E_t(u(w_{t+1}))\]

subject to the budget constraint,

\[(17) \quad B^d_{S,t} + B^d_{L,t} + B^d_{F,t} = W_t.\]

The superscript \( d \) denotes demand, \( W_t \) is the value of his assets in the beginning of the period \( t \), and \( W_{t+i} \) satisfies

\[(18) \quad W_{t+i} = B^d_{S,t} (1+r_t) + B^d_{L,t} (1+r_{L,t}) + B^d_{F,t} (1+r_{F,t}).\]

Here \( r_t \) is the short-term interest rate, and \( r_{L,t} \) and \( r_{F,t} \) are one-period
holding rates of return for domestic long-term bonds and foreign bonds, respectively.

In the partial equilibrium model, the stochastic processes for the interest rates are exogenously given, and the utility function is parameterized. The equilibrium exchange rate satisfies the condition that $B_{F,t}^d$ is equal to $B_{F,t}^s$, where $B_{F,t}^s$ is the supply of the foreign bonds to the domestic residents. The supply of foreign bonds is equal to the cumulative current account balance, and follows the dynamic equation:

\[(19) \quad B_{F,t}^s = B_{F,t-1}^s + C_t,\]

where $C_t$ is the current account balance in the period $t$. Neglecting interest received by holders of foreign bonds, $C_t$ is assumed to satisfy

\[(20) \quad C_t = -a + b \, s_t + u_t,\]

where $s_t$ is the log of the exchange rate expressed in terms of the domestic currency, $b$ is a positive number, and $u_t$ is the trade shock which is assumed to be white noise with variance $\sigma_u^2$.

Suppose that $\omega_{t+1}$ is normally distributed conditional on $\Omega_t$, and that the measure of the absolute risk aversion of $u_t, -u_t', u_t'$, is a positive constant $\alpha$. Then the demand function for foreign bonds and the demand function for long-term bonds are

\[(21) \quad B_{F,t}^d(E(r_{F,t}), r_t, E(r_{L,t})) = \psi \rho_{F,t} \cdot \psi \phi_{F,t},\]

\[(22) \quad B_{L,t}^d(E(r_{F,t}), r_t, E(r_{L,t})) = \psi \frac{\sigma^2}{\sigma_t^2} \rho_{L,t} \cdot \psi \phi_{F,t},\]

where
\[ \psi = \frac{1}{\alpha \sigma_s^2 (1 - \text{cor}^2)} \]

\[ \phi = -\text{cov} / \sigma_f^2 \]

\[ \sigma_s^2 = E_t \{ f(s_{t+1} - E_t(s_{t+1}))^2 \} \]

\[ \sigma_f^2 = E_t \{ f(r_{t+1} - E_t(r_{t+1}))^2 \} \]

\[ \text{cov} = E_t \{ f(s_{t+1} - E_t(s_{t+1}))(r_{t+1} - E_t(r_{t+1})) \} \]

\[ \text{cor} = \text{cov} / (\sigma_s^2 \sigma_f^2). \]

As in (21), the demand function for foreign bonds depends on \( \text{cov} \), the covariance conditional on \( \Omega_t \) between the exchange rate and the short-term interest rate, and \( \sigma_s^2 \), the conditional variance of the exchange rate. At the same time, the stochastic process of the exchange rate and \( \text{cov} \) depend on the demand function for foreign bonds. Therefore, it is necessary to solve for a rational expectations equilibrium in which the values of \( \text{cov} \) and \( \sigma_s^2 \) that the investors expect are consistent with the stochastic process of the exchange rate implied by the demand function consistent with these values of \( \text{cov} \) and \( \sigma_s^2 \).

The stochastic processes of short-term interest rates are assumed to be as follows:

\[ r_t = \mu + e_t + e_{t-1} + \epsilon_t \]

\[ R_t = \mu + e_t + \frac{1}{2} e_{t-1} \]

\[ r_{t}^* = \mu \]

where \( e_t \) and \( e_{t-1} \) are an interest rate shock with a duration of two periods.
and a temporary interest rate shock with a duration of one period, respectively. It is assumed that $e_t$ and $\epsilon_t$ are white noise with variance $\sigma_e^2$ and $\sigma_\epsilon^2$, respectively, and that they are independent of each other and of $u_t$.

The conditional expectation is assumed to coincides with the best linear prediction. Because (29) and (30) give a fundamental representation in the sense of linear prediction theory (see, e.g., Rozanov (1967)), observing the current and past values of $R_t$ is equivalent to observing the current and past values of $e_t$. It follows that

$$E_t(r_{t+1}) = \mu + e_t,$$

and from (13) and (30),

$$\rho_{L,t} = -e_t.$$

As shown in (29) and (30), the assumption employed here is that only the interest rate shock with the two period duration is transmitted to the long-term interest rate, so that the risk premium is equal to the temporary interest rate shock.

Define $\eta = \sigma_\epsilon^2 / \sigma_e^2$, which may be called the measure of substitution between short-term and long-term bonds. If $\eta=0$, then the risk premium for long-term bonds will always be zero, implying that short-term and long-term bonds are perfect substitutes. The greater is $\eta$, the smaller the degree of the substitution.

Solving the market equilibrium conditions as a difference equation system of $E_t(s_{t+\tau})$, with respect to $\tau$, provides the unique saddle point solution,

$$s_t = \bar{s} - \left[ \frac{1 - \lambda}{\beta} u_t \right] - \left[ \lambda (1 + \lambda) e_t + \lambda e_{t+\tau} \right] + \left[ \lambda (\phi - 1) e_t - \left( \frac{1 - \lambda}{\beta} \right) R_{t+\tau} \right].$$
where \( \tilde{z} = a/b \) is the long-run equilibrium exchange rate which clears the current account, and

\[
\lambda = 1 + \frac{b}{2\psi} - \frac{b}{2\psi} \sqrt{1 + 4\psi/b}.
\]

Here \( 0 < \lambda < 1, \partial \lambda / \partial \psi > 0, \lim_{\psi \to 0} \lambda = 0, \) and \( \lim_{\psi \to \infty} \lambda = 1. \)

Equation (34) shows that the investors' expected values of \( \text{cov} \) and \( \sigma_\varepsilon^2 \) affect the exchange rate dynamics through \( \lambda \) and \( \phi. \) On the other hand, the exchange rate dynamics in (34) imply certain values of \( \text{cov} \) and \( \sigma_\varepsilon^2, \) which need to be consistent with the investors' expected values in the rational expectations equilibrium.

The equilibrium is analyzed in two steps. First, the rational expectation of \( \text{cov} \) is solved. Ogaki (1998) shows that \( \text{cov} \) in the rational expectations equilibrium is given by

\[
\text{cov} = -\lambda(1 + \lambda + \eta)(1 + \eta)\sigma_\varepsilon^2/(1 + \eta + \eta \lambda) < 0.
\]

Second, the uniqueness and existence of the rational expectations equilibrium are shown by solving for the rational expectation of \( \text{var}. \) It turns out that the condition for the rational expectations equilibrium value for \( \text{var} \) is given by

\[
\frac{1}{\alpha} = g(\lambda),
\]

where

\[
g(\lambda) = \frac{\lambda \sigma_\varepsilon^2}{b} + \frac{b \lambda^2 (2 + \lambda)^2 \eta (1 + \eta) \sigma_\varepsilon^2}{(1 - \lambda)^2 (1 + (i + \lambda) \eta)^2}.
\]

It can easily be checked that \( g'(\lambda) > 0, \lim_{\lambda \to 0} g(\lambda) = 0, \) and \( \lim_{\lambda \to \infty} g(\lambda) = \infty. \)

Hence there exists a unique rational expectations equilibrium.

We note some features of (34). The discrepancy between the actual and
the long-run equilibrium exchange rate is explained by the trade shock (in the first brackets), the interest rate shock that is transmitted to the long-term interest rate (the second brackets), the interest rate shock that is not transmitted to the long-term interest rate (the third brackets), and the cumulative current account balance (in the fourth brackets). The trade shock, which tends to give rise to current account surplus, makes the domestic currency appreciate. Prolonged increases in the short-term interest rate make the domestic currency appreciate, though the shock which arises during this period, $e_i$, has a greater effect than the one from the previous period, $e_{i-1}$. As the cumulative current account balance becomes greater, the domestic currency’s appreciation increases: for investors to have incentives to hold more foreign bonds, the domestic currency must appreciate at present, so that investors can anticipate that it will depreciate in the future.

All of these effects are in the expected directions. The interest shock that is not transmitted to the long-term interest rate, $e_i$, however, has a perverse effect if the relative magnitude of the indirect complements relationship, $\phi$, is greater than 1. This is because the indirect complementarity of short-term and foreign bonds will exceed the direct substitutability if $\phi > 1$.

In the rational expectations equilibrium, the term $\phi$ is given by

\begin{equation}
\phi = \frac{\lambda(I + \lambda + \eta)}{(1 + \eta + \eta\lambda)} > 0.
\end{equation}

In the rational expectations equilibrium, the conditional covariance between the exchange rate and the short-term interest rate, $cov$, is negative. Hence, foreign bonds and long-term bonds are substitutes as explained in the
previous section. The measure of the relative magnitude of the indirect complementarity relationship, $\phi$, is positive. The main issue for the purpose of this paper is whether $\phi$ is greater or less than one. In order to determine this, we will investigate the sign of

$$
(40) \quad \phi-1 = \{(\lambda^2 + \lambda - 1) - \eta\}/(1 + \eta + \eta \lambda).
$$

Equation (40) shows that $\phi$ can be either greater or less than one, depending on the parameter values. One interesting case arises when the investors are close to being risk neutral. For a very small $\alpha$, an approximate formula for (40) with $\lambda = 1$ is

$$
(41) \quad \phi-1 \approx (1-\eta)/(1+2\eta).
$$

Hence the measure of the relative magnitude of the indirect complementarity, $\phi$, is greater than one if the measure of the degree of substitution between short-term and long-term bonds, $\eta$, is smaller than one. If the investors are close to being risk neutral and $\alpha$ is close to zero, the degree of substitution between short-term and long-term bonds must be high, and consequently, $\eta$ should be very small. Therefore, $\phi$ is greater than one, and the indirect complementarity dominates the direct substitutability for reasonable parameter configurations of the model.

6. Conclusions

This paper extends Mccsc and Rogoff's (1988) cointegrating regression by regressing the log of real exchange rate on the short-term real interest rate differential and the long-term real interest rate differential. When both interest rate differentials are included, we obtain a negative coefficient for the long-term interest rate differential that is predicted
by standard exchange rate models for all exchange rates we examine. However, we typically obtain a positive coefficient for the short-term interest rate differential for most of exchange rates we examined. Most of these coefficients are statistically significant. Ogaki and Santaella (1997) obtain similar results for Mexico.

These results are consistent with Ogaki’s (1995) model with risk averse agents when the indirect complementarity between the domestic short-term bonds and the foreign bonds dominates the direct substitutability of these assets. The model in this paper suggests that a much more complicated relationship may exist between short-term and long-term interest rates and exchange rates than is implied by exchange rate models with risk neutral agents. The model also applies to the relationship between exchange rates and the term structure of various short-term interest rates if the investment horizon is very short (e.g., 1 month or shorter). In this sense, the model could help explain Clarida and Taylor’s (1997) finding that the information in term structure of 1-month to 12-month forward premiums is useful in predicting future exchange rates.

The empirical results obtained in this paper show that attention must be given to the term structure of interest rates when considering exchange rate objectives. When the indirect complementarity dominates, the relationship between the domestic short-term interest rates and the exchange rate is complicated. If a central bank increases the domestic short-term interest rate, its effect on the exchange rate depends on how domestic long-term interest rates react. If the long-term interest rates rise in response to the rise in the short-term interest rate, a rise in the short-term interest rate will cause the domestic currency to appreciate as predicted by
the standard theory with risk neutral agents. If the long-term interest rate does not rise, a rise in the short-term interest rate will cause the domestic currency to depreciate.

REFERENCES


Baxter, M., 1994, "Real exchange rates and real interest differentials: Have we missed the business-cycle relationship?" Journal of Monetary Economics 33, 5-37.


Meredith, G. and M.D. Chinn, 1998, "Long-Horizon Uncovered Interest Parity <


Table 1. Tests for the Null of the Difference Stationary

<table>
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<th></th>
<th>s</th>
<th>r^*_t</th>
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Note. a. Critical values for the 1%, 5%, and 10% significance levels are 0.1166, 0.2937, and 0.4544. These are from Park and Choi(1988).
b. Critical values for the 1%, 5%, and 10% significance levels are -3.4456, -2.8418, and -2.5731 for U.K., Germany, Canada, France, and Italy. Those for Japan and Switzerland are -3.4926, -2.8760, and -2.5688 and 3.4642, 2.9124, and 3.4642, respectively.
c. Significant at the 1% level.
d. Significant at the 5% level.
e. Significant at the 10% level.
Table 2  Canonical Cointegrating Regression Results

(A) \[ s_t = \mu + \alpha (r_{l,t} - r^*_{l,t}) + \beta (r_{s,t} - r^*_{s,t}) + \varepsilon_t \]

(B) \[ s_t = \mu + \gamma t + \alpha (r_{l,t} - r^*_{l,t}) + \beta (r_{s,t} - r^*_{s,t}) + \varepsilon_t \]

<table>
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<th>( \hat{\alpha} )</th>
<th>( \hat{\beta} )</th>
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<th>H(1,2)</th>
<th>H(1,3)</th>
<th>H(1,4)</th>
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<td>(2)</td>
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<td>(4)</td>
<td>(5)</td>
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<td>(0.497)</td>
<td>(0.800)</td>
<td>(0.683)</td>
<td>(0.854)</td>
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Note: In cols. 1, 2, and 3, standard errors are in parentheses.

Column 4 is a \( \chi^2 \) test statistic with one degree of freedom for the deterministic cointegration restriction.

Cols 5, 6, and 7 is a \( \chi^2 \) test statistic with one, two, and three degree of freedom for the stochastic cointegration restriction, respectively. P-values are in parentheses.

a. Significant at 1% level.
b. Significant at 5% level.
c. Significant at 10% level.
e. For the pound, CCR results with 10 years government bond rate are \( \hat{\alpha} = -10.252\(^a\), \hat{\beta} = 2.642\), p-value of H(1,2) = 0.023, that of H(1,3) = 0.063, and that of H(1,4) = 0.122.
Table 3. Canonical Cointegrating Regression Results

\[
\begin{align*}
(A) & \quad s_t = \mu + \alpha(r_{t,t} - r^*_{t,t}) + \beta(r^N_{t,t} - r^N_{s,t}) + \varepsilon_t \\
(B) & \quad s_t = \mu + \gamma t + \alpha(r_{t,t} - r^*_{t,t}) + \beta(r^N_{t,t} - r^N_{s,t}) + \varepsilon_t
\end{align*}
\]

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<th>(H(1,4))</th>
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Note: In cols. 1, 2, and 3, standard errors are in parentheses.

Column 4 is a \(\chi^2\) test statistic with one degree of freedom for the deterministic cointegration restriction.

Cols 5, 6, and 7 is a \(\chi^2\) test statistic with one, two, and three degree of freedom for the stochastic cointegration restriction, respectively. P-values are in parentheses.

a. Significant at 1% level.
b. Significant at 5% level.
c. Significant at 10% level.