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Abstract

Many researchers have found that changes in real exchange rates based on general price indexes such as Consumer Price Indexes (CPI) or Gross Domestic Product (GDP) deflators are very persistent, implying that Purchasing Power Parity (PPP) does not hold for these general price indexes even in the long run. If PPP holds for tradable goods but not for nontradable goods in the long run, and if the weights placed on tradable and nontradable good prices in general price indexes are stable, then relative prices between these nontradables and tradables will move together with real exchange rates based on general price indexes. In this paper, we identify time periods, countries and relative price measures for which these comovements between real exchange rates and relative prices of nontradables and tradables are observed.

I. Introduction

Purchasing Power Parity (PPP) is said to hold in the long run when a measure of the real exchange rate, defined as the nominal exchange rate multiplied by a domestic price index divided by a foreign price index, is stationary. If the real exchange rate is stationary, then it has a tendency to revert to its mean value at least in the long run. Many researchers have found empirical evidence against the hypothesis that PPP holds in the long run when general price indexes such as Consumer Price Indexes (CPI) for total consumption expenditures and/or Wholesale Price Indexes (WPI) are used as the price indexes to define real exchange rates.[1]

One possible explanation for these empirical failures of PPP is that the post-Bretton Woods float period is not long enough for the long run PPP to hold given that many authors use this float period in their empirical work. For example, Diebold, Husted, and Rush (1991) study the behavior of real exchange rates during the gold standard era with samples spanning from 74 years to 123 years and find evidence in favor of long run PPP. Grilli and Kaminsky (1991) also find evidence in favor of long run PPP when they use historical data. Even though this explanation is plausible, it is still of interest to study why changes in real exchange rates are so persistent.

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[1]For example, Fisher and Park (1991) reject the hypothesis that the cointegrating vector is (1, 1, -1) for the nominal exchange rate, the home country price level, and the foreign country price level even when they do not reject the null hypothesis of cointegration for many of the country pairs they study. Some authors find that the null hypothesis of no cointegration for these variables and/or the null hypothesis that real exchange rates are unit root nonstationary cannot be rejected for many country pairs (e.g., Corbae and Ouliaris 1988; Karfakis and Moschos 1989; Mark 1990; Patel 1990). Many authors (e.g., Roll 1979; Adler and Lehmann 1983; Darby 1983; Mussa 1986; Diebold 1988; Meese and Rogoff 1988; Baillie and McMahon 1989) fail to reject a stronger hypothesis that real exchange rates are martingales.

Even if we believe that real exchange rates will exhibit tendencies to revert to their means in 50 years, we would like to know the factors that have driven real exchange rates over the past two decades.

One factor that may be important in explaining persistence of changes in real exchange rates is the existence of nontradable goods.[2] For example, Kim (1990) finds more favorable evidence for long-run PPP when he uses WPI than when he uses CPI.[3] Vataja and Ogaki (1994) find evidence in favor of long-run PPP for many industries when they use industry-level price indexes. These authors argue that a large weight placed on nontradable goods in the CPI may be the main reason why long run PPP based on CPI is not empirically supported.

If PPP holds for tradable consumption goods but not for nontradable consumption goods in the long run, and if the weights placed on tradable and nontradable good prices in general price indexes are stable, then relative prices between these nontradables and tradables will move together with real exchange rates based on general price indexes. In this paper, we examine whether or not we can observe these comovements between real exchange rates and available proxies of relative prices between nontradables and tradables.

It should be noted that there are many reasons why we may not observe such comovements. First, many tradable goods across countries are not identical. It is also hard to find pure tradable goods because tradable goods are often combined with nontradable services such as retailing services. Therefore, PPP may not hold for these tradable goods in the long

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[2]See, e.g., Jones and Purvis (1983). Recent empirical studies of exchange rates in the presence of nontradable goods include Engel (1993) and Rogers and Jenkins (1993).

[3]It should be noted that Kim uses a much longer sample period than some of the papers cited earlier that provide evidence against PPP with WPI.

run. Second, the weights placed on tradable and nontradable goods may not be stable, and these weights, rather than the relative prices of nontradables and tradables, may be moving together with real exchange rates. Third, available proxies of relative prices of tradable and nontradable goods may be poor approximations of the relative prices.

For most of the long-run movements in real exchange rates, we find comovements between relative prices and real exchange rates. On the other hand, for some medium-run swings of real exchange rates, we fail to find such comovements. Because it is arguably easier to investigate the determinants of movements in relative prices than the determinants of movements in exchange rates, it seems useful to identify time periods and countries for which measures of relative prices and real exchange rates move together. For example, there are at least three possible interpretations for the comovements of relative prices and real exchange rates. By studying time periods in which relative prices move together with real exchange rates, we may be able to understand whether or not any of the following interpretations is appropriate.

One interpretation is that differences in productivity in nontradable and tradable good industries in the long run are causing changes in relative prices and then real exchange rates in the long run (see Hsieh (1982)). If stochastic or deterministic trends of the productivity in nontradable good industries are different from those in tradable good industries, then these differences will cause changes in the relative prices in the long run. These changes in the relative prices in turn will cause changes in GDP deflator-based real exchange rates in the long run. This interpretation is based on a version of Balassa (1964) and Samuelson's (1964) model.

The second interpretation is that the relative prices between nontradables and tradables follow random walks because consumers smooth tradable good consumption as suggested by Rogoff (1992).[4]

Another interpretation is that the movements in nominal exchange rates are causing changes in relative prices for nontradables and tradables. For example, when monetary factors cause changes in nominal exchange rates, the relative prices need to change as long as PPP holds for tradables.

The rest of this paper is organized as follows. In Section II, we explain the model we use to interpret our empirical results. In Section III, we describe the econometric method that we use. Section IV presents our empirical results. Section V contain our concluding remarks.

II. The Econometric Model

Real Exchange Rate Regressions

Consider a world economy with and two countries: country H is the home country and country F is the foreign country. In each country, there are two goods: good N is nontradable and good T is tradable. In terms of country j's currency ($j=H,F$), $p_j^i(t)$ is the price of good i ($i=N,T$) at time t. Suppose that the general price index (in our empirical work, either the GDP deflator or the total consumption deflator) in country j can be expressed as

$$(1) \qquad P_j(t) = c_j(t)[p_j^N(t)]^{\alpha_j}[p_j^T(t)]^{1-\alpha_j},$$

where α_j is a constant, and $c_j(t)$ is a stationary random variable that reflects any factor that causes the general price index to deviate from $[p_j^N(t)]^{\alpha_j}[p_j^T(t)]^{1-\alpha_j}$ (such as a measurement error). Let $E(t)$ be the nominal exchange rate: one unit of the home country currency is $E(t)$ units of the

[4]Also see Razin (1993).

foreign country currency. We denote the relative price between the nontradable good and the tradable good by

$$(2) \qquad q_j(t) = \frac{p_j^N(t)}{p_j^T(t)},$$

for country j. Even though this relative price is often called the real exchange rate especially in theoretical work, we reserve the words "real exchange rate" for the real exchange rate defined by the general price index:

$$(3) \qquad E_r(t) = E(t)\,\frac{P_H(t)}{P_F(t)}\,.$$

This is because most papers in the empirical literature that test the PPP are concerned with this definition of the real exchange rate.

We assume that tradable goods in these countries are not identical, so that PPP does not hold even for the tradable goods in the short run. The PPP for the tradable goods is assumed to hold in the long run:

$$(4) \qquad ln(p_H^T(t)) = ln(p_F^T(t)) - ln(E(t)) + \varphi(t),$$

where $\varphi(t)$ is a stationary random variable with mean zero. Because $\varphi(t)$ is stationary, (4) implies that a shock in the tradable good price in a country that causes a deviation from the PPP eventually dies out.

From (1) and (2), we obtain

$$(5) \qquad ln(P_j(t)) = \theta_j + \alpha_j ln(q_j(t)) + ln(p_j^T(t)) + \psi_j(t),$$

where $\theta_j = E(ln(c_j(t)))$ and $\psi_j(t) = ln(c_j(t))\text{-}\theta_j$ for $j = H,T$. Using (3)-(5), we get

$$(6) \qquad ln(E_r(t)) = \theta + \alpha_H ln(q_H(t)) - \alpha_F ln(q_F(t)) + \varepsilon(t),$$

where $\theta = \theta_H - \theta_F$ and $\varepsilon(t)=\varphi(t)+\psi_H(t)-\psi_F(t)$. Relation (6) is the starting point of our analysis.

If the relative prices $q_H(t)$ and $q_F(t)$ are stationary, then Relation (6) implies that the real exchange rate is stationary. We assume that the logs of the relative prices are first difference stationary and thus possess stochastic trends. With this assumption, Relation (6) implies that $ln(E_r(t))$, $ln(q_H(t))$, and $\alpha_F ln(q_F(t))$ are cointegrated with a cointegrating vector $(1, -\alpha_H, \alpha_F)$. If the technological progress in the nontradable good industry and the technological progress in the tradable good industry have different stochastic trends and if preferences are stable, then the relative prices will need to possess stochastic trends (see, e.g., Ogaki and Park 1989, Ogaki 1992, and Clarida 1994 for some empirical evidence that stochastic trends exist for various relative prices).

In reality, it is hard to find a price index of all pure nontradable goods. Imagine that a proxy for $p_j^N(t)$ is actually a weighted average of the price of another tradable good and the price of a pure nontradable good:

$$(7) \qquad p_j^N(t) = [p_j^{NT}(t)]^{\beta_j}[p_j^{NN}(t)]^{1-\beta_j}.$$

Then

$$\begin{aligned}(8) \qquad ln(q_j(t)) &= ln(P_j^N(t)) - ln(P_j^T(t)) \\ &= \beta_j ln(P_j^{NT}(t)) + (1-\beta_j)ln(P_j^{NN}(t)) - ln(P_j^T(t)) \\ &= \{ln(P_j^{NN}(t)) - ln(P_j^T(t))\} - \beta_j\{ln(P_j^{NN}(t)) - ln(P_j^{NT}(t))\}\end{aligned}$$

Hence the trends of $ln(q_j(t))$ reflect the trends of the relative prices of the nontradable good and the tradable goods. Therefore, Relation (6) can still be used for a real exchange rate cointegrating regression in this case. The only difference is that estimated α_H and α_F cannot be interpreted as the weights placed on the nontradable goods.

Identification

It is important to distinguish two cases in order to think about the identification of the unknown parameters, α_H and α_F.

Case 1: $ln(q_H(t))$ and $ln(q_F(t))$ are not stochastically cointegrated.

For the identification of the parametrers in the cointegrating regression applied to (6), it does not matter whether or not the deterministic cointegration restriction is satisfied by these relative prices (see Ogaki and Park 1989 for definitions of stochastic cointegration and the deterministic cointegration restriction). Case 1 is relevant if the technological progress in the nontradable good industry in the home country has a different stochastic trend from that in the foreign country. On the other hand, if the technological progress in the nontradable good industry is transmitted at least in the long run, we have

Case 2: $ln(q_H(t))$ and $ln(q_F(t))$ are stochastically cointegrated.

In Case 1, a cointegrating regression is applied to (6), and α_H and α_F are identified. When the real exchange rate and the relative prices have nonzero drift terms, (6) implies that the deterministic cointegration restriction is satisfied. We can test the model by testing the null hypothesis of stochastic cointegration and the deterministic cointegration restriction. We can also check whether or not point estimates are plausible in light of the restrictions that $0<\alpha_j<1$ for $j=H, F$.

It should be noted that no econometric method can identify α_H and α_F in

Case 2 unless further restrictions are imposed. This is because $ln(q_H(t))$ and $ln(q_F(t))$ have a common stochastic trend (in the terminology of Stock and Watson 1988) and do not have distinct stochastic trends. Because the only restrictions imposed by the model are on the trend components of variables and no restrictions are imposed on the stationary component of the variables, the coefficients for two variables with one common stochastic trend cannot be identified.

In Case 2, there are two important subcases.

Case 2A: $ln(q_H(t))$ and $ln(q_F(t))$ satisfy the deterministic cointegration restriction.

and

Case 2B: $ln(q_H(t))$ and $ln(q_F(t))$ do not satisfy the deterministic cointegration restriction.

In Case 2A, let

(9) $$ln(q_F(t)) = \theta_q + \beta_q ln(q_H(t)) + \varepsilon_q(t),$$

where (1, $-\gamma$) is a cointegrating vector, $\varepsilon_q(t)$ is stationary with zero mean. From (6) and (9), we obtain

(10) $$ln(E_r(t)) = \theta_2 + \beta ln(q_H(t)) + \varepsilon_2(t),$$

where $\theta_2 = \theta-\alpha_F\theta_q$, $\beta = \alpha_H - \alpha_F\beta_q$, and $\varepsilon_2(t) = \varepsilon(t) - \alpha_F\varepsilon_q(t)$. Assuming that $\alpha_H \neq \alpha_F\beta_q$, (10) implies that $ln(E_r(t))$ and $ln(q_H(t))$ are cointegrated with the deterministic cointegration restriction. A cointegrating regression is applied to (10) to estimate $\beta = \alpha_H - \alpha_F\beta_q$ in this case. As in Case 1, we can test the model by testing the null hypothesis of stochastic cointegration and the deterministic cointegration restriction.

In Case 2B, let

$$(11) \qquad ln(q_F(t)) = \theta_q + \mu_a t + \beta_q ln(q_H(t)) + \varepsilon_q(t),$$

where $\varepsilon_q(t)$ is stationary with zero mean and $\mu_a \neq 0$ because $ln(q_F(t))$ and $ln(q_H(t))$ are stochastically cointegrated without the deterministic cointegration restriction in this case. From (6) and (11), we obtain

$$(12) \qquad ln(E_r(t)) = \theta_2 + \mu t + \beta ln(q_H(t)) + \varepsilon_2(t),$$

where $\mu = -\alpha_F \mu_a$. Assuming that $\alpha_H \neq \alpha_F \beta_q$, (12) implies that $ln(E_r(t))$ and $ln(q_H(t))$ are stochastically cointegrated without the deterministic cointegration restriction. In this case, a cointegrating regression is applied to (12) to estimate $\beta = \alpha_H - \alpha_F \beta_q$. We can test the model by testing for stochastic cointegration.

Time-Varying Weights

The weight placed on the nontradable good changes over time. If the weight changes with trends, the real exchange regression (6) is misspecified. In this case, the general price index is

$$(1') \qquad P_j(t) = c_j(t)[p_j^N(t)]^{\alpha_{j(t)}}[p_j^T(t)]^{1-\alpha_{j(t)}},$$

and $\alpha_j(t)$ is the weight placed on the nontradable good. If data on the weight are available, we can construct a regressor $\alpha_H(t)ln(q_H(t))$ - $\alpha_F ln(t)(q_F(t))$ and run a regression

$$(6') \qquad ln(E_r(t)) = \theta + \beta\{\alpha_H(t)ln(q_H(t)) - \alpha_F ln(t)(q_F(t))\} + \varepsilon(t),$$

where β=1.

III. Econometric Methodology

This section describes the econometric procedure that we use. This procedure allows us to test the null hypothesis of stochastic cointegration and the deterministic cointegration restriction. One word of caution is in order: even though concepts of unit roots and cointegration are formally about the infinite future, we never learn about infinite future from finite samples (see, e.g., Campbell and Perron 1991 and the references therein for this near observational equivalence problem). Therefore, from our empirical results, we only learn about "long-run" relative to our sample period. This is not a problem as long as it is understood that we use terminologies regarding the infinite future for convenience. Our purpose is to study why GDP-based real exchange rates are so persistent.

Let $y(t) = ln(E_r(t))$. We apply Park's (1992) Canonical Cointegrating Regressions (CCR) procedure[5] to

(13) $$y(t) = \gamma_d' d(t) + \gamma' X(t) + \varepsilon_c(t),$$

where $d(t) = 1$ in Case 1 and Case 2A, $d(t) = (1,t)'$ in Case 2B, $X(t)=[ln(q_H(t)),\ ln(q_F(t))]'$ in Case 1, $X(t)=ln(q_H(t))$ in Case 2A and Case 2B, $\varepsilon_c(t) = \varepsilon(t)$ in Case 1, and $\varepsilon_c(t) = \varepsilon(t) - \alpha_F \varepsilon_q(t)$ in Case 2. The CCR procedure requires us to transform the data before running a regression and corrects for endogeneity and serial correlation. Let $v(t) = [\varepsilon_c(t),\ \Delta X(t)-E(\Delta X(t))]'$. Define $\Phi(i) = E(v(t)v(t-i)')$, $\Sigma=\Phi(0)$, $\Gamma = \sum_{i=0}^{\infty}\Phi(i)$, and $\Omega =$

[5] See Ogaki (1993b) for a more detailed explanation of CCR-based estimation and testing. In this paper, notations for test statistics follow Ogaki (forthcoming). Park and Ogaki's (1991a) Seemingly Unrelated Canonical Cointegrating Regression (SUCCR) applied to (7) and (8) in Case 2A and SUCCR applied to (9) and (10) in Case 2B are asymptotically more efficient than CCR. We do not apply SUCCR because Park and Ogaki's Monte Carlo results suggest that the efficiency gain in small samples tend to be small.

$\sum_{i=-\infty}^{\infty}\Phi(i)$. Here Ω is the long run covariance matrix of v_t. Define

$$\Omega_{11.2} = \Omega_{11} - \Omega_{12}\Omega_{22}^{-1}\Omega_{21} \tag{14}$$

and $\Gamma_2 = (\Gamma'_{12}, \Gamma'_{22})'$, where Ω_{ij} and Γ_{ij} are the *ij*th component of Ω and Γ, respectively. We make an additional assumption that $\Omega_{11.2}$ is positive.[6] Consider transformations

$$y^*(t) = y(t) + \Pi'_y v(t) \tag{15}$$

$$X^*(t) = X(t) + \Pi'_x v(t). \tag{16}$$

Because $v(t)$ is stationary, $y^*(t)$ and $X^*(t)$ are cointegrated with the same cointegrating vector $(1, -\gamma)$ as $y(t)$ and $X(t)$ for any Π_y and Π_x. The idea of the CCR is to choose Π_y and Π_x, so that the OLS estimator is asymptotically efficient when $y^*(t)$ is regressed on $X^*(t)$. This requires

$$\Pi_y = \Sigma^{-1}\Gamma_2\gamma + (0, \Omega_{12}\Omega_{22}^{-1}) \tag{17}$$

$$\Pi_x = \Sigma^{-1}\Gamma_2. \tag{18}$$

In practice, long run covariance parameters in these formulas are estimated, and estimated Π_y and Π_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 be essentially considered as normal distributions, so that their standard errors can be interpreted in the usual way.[7] An important property of the

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[6]This additional assumption is satisfied if $\Delta^{-1}\varepsilon(t)$ is not stochastically cointegrated with $X(t)$ as discussed in Park (1992).

[7]The CCR estimators are asymptotically efficient, but there are other asymptotic efficient estimators by Phillips and Hansen (1990), Phillips (1991), Saikkonen (1992), and Stock and Watson (forthcoming), among others.

CCR procedure is that linear restrictions can be tested by χ^2 tests which are free from nuisance parameters. We use χ^2 tests in a regression with spurious deterministic trends added to (14) to test for stochastic and deterministic cointegration. For this purpose, the CCR procedure is applied to a regression

$$(19) \qquad X_1(t) = \theta_c + \sum_{i=1}^{q} \eta_i t^i + \gamma X_2(t) + \varepsilon_c(t).$$

Let $H(p,q)$ denote the standard Wald statistic to test the hypothesis $\eta_p=\eta_{p+1}= \dots =\eta_q=0$ with the estimate of the variance of $\varepsilon_c(t)$ replaced by $\Omega_{11.2}$ (see Park 1990 for more explanation). Then $H(p,q)$ converges in distribution to a χ^2_{p-q} random variable under the null hypothesis of cointegration. In particular, the $H(0,1)$ statistic tests the deterministic cointegrating restriction. On the other hand, the $H(1,q)$ tests stochastic cointegration.[8]

Johansen's (1988, 1991) estimators are often used, but Johansen assumes Gaussian VAR structure. The CCR does not require this Gaussian VAR assumption. Monte Carlo experiments in Park and Ogaki (1991b) show that the CCR estimators have better small sample properties in terms of the mean square error than Johansen's estimators. Following Monte Carlo-based recommendations of Park and Ogaki (1991b) and Han and Ogaki (1991), we use Park and Ogaki's VAR prewhitening method and report the third step CCR estimates and fourth step CCR test results.

[8]Monte Carlo results in Han and Ogaki (1991) suggest that the H(0,1) test for the deterministic cointegration does not have a severe size distortion problem and has fairly high size adjusted power against no cointegration. Efficiency gains in estimating the cointegrating vectors from imposing the deterministic cointegration cointegration restriction was discussed by West (1988) for the one regressor case and by Hansen (1992) and Park (1992) for the general multiple regressors case.

IV. Empirical Results

A. Results for CPI and WPI

In this section we describe our empirical results which are presented in the first panels of Tables 2 through 4 and Table 5 for WPI and CPI.[9] As our proxies for tradable and nontradable good price indexes, we use WPI and CPI, respectively. As mentioned in the Introduction, Kim (1990) finds evidence in favor of long-run PPP for WPI based real exchange rates but not for CPI based real exchange rates. Kim argues that his finding can be explained by the fact that CPI gives much larger weight to nontradable goods than WPI does. This suggests that WPI and CPI should serve as reasonable proxies of tradable and nontradable goods prices, respectively for the purpose of investigating long-run comovements between real exchange rates and relative prices.

We present results for both the full sample period (which is from 1929-1988) and the pre free float period of 1929-1972 as the free float period of 1973 onwards is quite different from the earlier era of fixed exchange rates.

Table 2 reports the test results for the assumption that the log of relative prices are first difference stationary and possess stochastic trends. If this assumption does not hold, then relative prices cannot explain why real exchange rates possess stochastic trends.

In Table 2 we present the results of the various tests for the null hypothesis of difference stationarity of the log of relative prices. Along with the conventional Dickey-Fuller (DF) and Said and Dickey (SD 1984)

[9] All empirical results are obtained by applying Ogaki's (1993a) GAUSS CCR programs.

tests, we also use Park's (1990) *J*(1,5) test. The *J*(1,5) test does not require the estimation of the long run variance of the disturbance of the regression and thus has an advantage over the SD and Phillips-Perron (1988) tests in that neither the order of the autoregression nor the lag truncation number needs to be specified. The *J*(1,5) test does not reject the null hypothesis of difference stationarity for any of the countries in any of the periods at the 10% level. The DF test also does not reject the null hypothesis of a unit root for any of the countries in any of the periods at the 5% level. The same is true of the SD tests .

Thus the results of the first panel of Table 2 can be summarized by saying that there is evidence in favor of the hypothesis that the log of the relative price for each country is first difference stationary and thus possesses stochastic trends.

Tables 3 and 4 report the results of the tests conducted to see whether or not the relative prices of the home and foreign countries are cointegrated, i.e., to see whether data support Case 1 or Case 2.

Table 3 presents Park's (1990) *H*(p,q) tests for the null hypothesis of stochastic cointegration and the deterministic cointegration restriction. In the first two panels of Table 3, the coefficient on the Italian relative price is significant and none of the H(p,q) tests is significant at the 10% level. In the third and fourth panels of Table 3 we find that the coefficient on the British relative price is significant and that once again none of the tests is significant at the 10% level. In the fifth and sixth panels of Table 3 we find that the coefficient on the relative price of Italy is not significantly different from zero and that the H(1,2) and H(1,3) tests are significant at the 5% or 10% level, although H(0,1) is not. Thus the H(p,q) tests seem to favor the null of cointegration for USA/ITALY and USA/UK but not for UK/ITALY.

In Table 4 we present the results of the various tests for the null hypothesis of no stochastic cointegration between relative prices. Along with the conventional Dickey-Fuller (DF) and Said and Dickey (SD 1984) tests, we also use Park's (1990) $I(1,5)$ test. The $I(1,5)$ test basically applies the $J(1,5)$ test to the residuals from an OLS cointegrating regression. Thus, like the J-test, it has an advantage over augmented Dickey-Fuller tests and Phillips-Ouliaris's (1990) tests in that neither the order of autoregression nor the bandwidth parameter needs to be specified.

The $I(1,5)$ test does not reject the null hypothesis of no stochastic cointegration for any of the countries in any of the periods at the 10 % level. With the exception of the SD(1) statistic for USA/ITALY in the pre float sample and USA/UK in the full sample, none of the DF or SD statistics are significant at the 5% level. Thus Table 4 seems to provide evidence in favor of Case 1 relative to Case 2.

The results of Tables 3 and 4 can be summarized by saying that the evidence between cases 1 and 2 is mixed for (USA/ITALY) and (USA/UK), but for (UK/ITALY) we have unambiguous evidence in favor of Case 1.

Table 5 reports the results of the estimation of Relation (6) under Cases 1 and 2. A war time dummy variable (for the second world war) was included for all real exchange rate regressions involving Italy.

The first part of Table 5 gives the results for the U.K.-Italy real exchange rate regressions, for both samples, under cases 1 and 2. The first two panels report the results under Case 1. Both relative prices are significant and we estimate their coefficients with the theoretically correct signs and magnitudes. The H(0,1) statistic does not reject the model, but the H(1,2) and H(1,3) statistics are significant, at the 5% level, for the full sample. We also report results under Case 2 in the third through sixth panels as a sensitivity analysis because Case 1 and Case 2 may not be well discriminated in small samples. Under Case 2, the H(0,1) statistic in panel 3 is significant at the 5 % level and the H(0,1) statistics in panels 4 and 6 are significant at the 10% level but not at the 5% level. None of the other test statistics are significant at the 10% level. Thus, for UK/ITALY, we find evidence against the model at the 5% level for the full sample but not for the pre-float sample.

The second part of Table 5 gives the results for both periods for the USA-U.K. real exchange rate under cases 1 and 2. Under Case 1, we estimate α_H and α_F with the theoretically correct signs but not the theoretically correct magnitudes. The H(0,1) and H(1,3) statistics for the full sample period are significant at the 5 % level although none of the H(p,q) statistics are significant for the pre float period. Under Case 2, we find that at least one of the H(p,q) statistics is significant at the 5 % level for the full sample; whereas for the pre float sample none of the statistics is significant at the 5 % level.

The last part of Table 5 gives the results for both sample periods and both cases for the US-Italy real exchange rate. The first two panels report the results under Case 1. In the first panel, we estimate the parameters with the theoretically correct signs and magnitudes but the coefficient of the US relative price is not significant. For the full sample, the H(0,1) test statistics do not reject the model, but the H(1,2) and H(1,3) statistics are significant at the 1% level. In the second row, the coefficient of the US relative price is not significant. Under Case 2, we find that the H(0,1) statistic in panel three and the H(1,2) statistic in panel five are significant at the 5% level, and the H(1,3) statistic in panel five is significant at the 10% level.

Thus, for all three country pairs, we find that for the full sample, at least one H(p,q) statistic rejects the model under Cases 1 and 2. For the pre float sample, we find that none of the H(p,q) statistics is significant at the 5 % level under either case. Given that it is difficult to discriminate Cases 1 and 2 in small samples, the results of Table 5 can be summarized as saying that there is evidence in favor of the model for the pre-float sample period but not for the full sample period, when we use WPI/CPI as a proxy for the relative price of nontradables.

We plot figures of the actual real exchange rate and the regression of the real exchange rate on relative prices estimated by CCR to see how well our model can explain the long run movements of the real exchange rate for the three country pairs we consider.

Figure 1 shows plots of the UK-Italy real exchange rate and the estimated regression corresponding to panel 1 of the first part of Table 5. The regression captures the general movements of the real exchange rate but misses out on the short and medium term swings.

Figure 2 shows plots of the U.K.-US real exchange rate and the estimated regression corresponding to panel 1 of the second part of Table 5. The regression seems to capture quite well the tendency of the real exchange rate to rise from 1945 to 1950 and to then to fall from 1950 to the late seventies.

Figure 3 shows a plot of the US-Italy real exchange rate and the regressions corresponding to the first two panels of the third part of Table 5. The estimated regression seems to capture the long run tendency for the real exchange rate to fall from 1948 to 1988 although it misses out on the short and medium term swings.

Thus, on the whole we find evidence that, for the pre float regime, real exchange rates are significantly affected by relative prices between nontradables and tradables in the long run for all the country pairs. However, more evidence against our model is found when we extend our sample to include the free float period.

B. Results for Consumption Deflators

In this section we describe our empirical results which are presented in the second panels of Tables 2 through 4 and Table 6.

Table 2 reports the test results for the assumption that the log of relative prices are first difference stationary and possess stochastic trends. If this assumption does not hold, then the relative prices cannot explain why real exchange rates possess stochastic trends.

Along with the conventional Dickey-Fuller (DF) and Said and Dickey (SD 1984) tests, we also use Park's (1990) $J(1,5)$ test which was described earlier in subsection *A*. The $J(1,5)$ test does not reject the null hypothesis of difference stationarity for any of the countries in any of the periods at the 10% level. The DF test also does not reject the null hypothesis of a unit root for any of the countries in either period. The same is true of the SD tests -with the exception of Canada in the free float period, for which we find that the null hypothesis of difference stationarity is rejected at the 5% level of significance for SD(4) and

SD(7).

Thus the results of Table 2 can be summarized by saying that there is evidence in favor of the hypothesis that the log of the relative price for each country is first difference stationary and thus possesses stochastic trends, with the possible exception of Canada in the free float period.

The second panels of Tables 3 and 4 report the results of the tests conducted to see whether or not the relative prices of the home and foreign countries are cointegrated,i.e., to see whether data support Case 1 or 2.

Table 3 presents Park's (1990) *H*(p,q) tests for the null hypothesis of stochastic cointegration and the deterministic cointegration restriction. The first five rows (of the second panel) present the results with the US as the home country and U.K., Canada, and Japan as the foreign countries.

In the first row, the coefficient on the U.K. relative price is not significantly different from zero although none of the H(p,q) tests is significant at the 5% level. In the second row, the H(1,2) test statistic is significant at the 5% level. In the third row, the coefficient on the Canadian relative price is not significant although none of the H(p,q) statistics are significant. In the fourth row, the H(1,2) and H(1,3) statistics do not reject stochastic cointegration, but the H(0,1) statistic rejects the deterministic cointegration restriction at the 1% level. In the fifth row, the H(0,1) and H(1,2) statistics are not significant, but the H(1,3) statistic rejects stochastic cointegration at the 1% level.

The sixth through ninth rows (of the second panel) in Table 3 present the results for Canada/UK, Canada/Japan and UK/Japan.

For all three combinations we find that the H(1,2) and H(1,3) statistics do not reject the null of stochastic cointegration at the 5%

level, although the coefficient on the UK and Japanese relative prices is not significant in the seventh and eighth rows.

The results of Table 3 can be summarized by saying that for most country pairs the H(p,q) tests do not provide strong evidence, either in favor of, or against, the null of stochastic cointegration.

In Table 4 we report the $I(1,5)$, Dickey-Fuller, and Said-Dickey tests for the null hypothesis of no stochastic cointegration between the relative prices of the home and foreign countries.

The $I(1,5)$ test does not reject the null hypothesis of no stochastic cointegration for any of the countries for any of the periods at the 10 % level. The Dickey-Fuller and SD tests are also not significant for any country pair in either period. Thus, Table 4 provides strong evidence in favor of Case 2 for all country pairs and for both time periods.

Overall, with the exception of US/Japan, we don't have unambiguous evidence in favor of Case 1 or Case 2.

Table 6 reports the results of the estimation of Relation (6) under Cases 1 and 2.

The first panel of Table 6 presents the results of the real exchange rate regressions for USA/UK, for both sample periods, under cases 1 and 2. In the first row, we estimate both coefficients with the theoretically correct signs and magnitudes. The coefficient on the US relative price is not significant and none of the H(p,q) statistics is significant at the 5% level. In the second row, we estimate the coefficients with the theoretically correct signs but incorrect magnitudes, although the standard errors for these estimates are large. Both coefficients are significantly different from zero and none of the H(p,q) statistics is significant. We also report results under Case 2 in the third through sixth rows as a sensitivity analysis. None of H(0,1),H(1,2) and H(1,3) is significant at the 5% level under Case 2, although the coefficient on the relative price is insignificant in the fourth and sixth rows. Thus, for US/UK, we have evidence in favor of the model under both cases and in both samples.

The second panel of Table 6 presents the results for the US/Canada real exchange rate, for both sample periods, under cases 1 and 2. Under Case 1, we estimate α_H and α_F with the theoretically correct signs and magnitudes for the full sample period. The coefficients of both relative prices are significant. The H(0,1) test is significant and rejects deterministic cointegration at the 5% level but not at the 1% level. For the free float sample under Case 1, both coefficients have the correct signs and are significant. The point estimate of the coefficient for the relative price in Canada is larger than one in absolute value, but is not significantly so. None of the H(p,q) test statistics are significant. Under Case 2, none of the relative price coefficients is significant and the H(0,1) statistics in the fifth and sixth rows are significant at the 5 % level. Thus, for US/Canada, we have evidence in favor of the model under Case 1 for the free float period.

The third panel of Table 6 presents the results for the US/Japan real exchange rate, for both sample periods, under cases 1 and 2. Under Case 1, we estimate the coefficients with the theoretically correct signs although the magnitudes are implausible. The H(0,1) statistic is significant at the 5% level but not at the 1% level. Under Case 2, the H(1,3) statistic in the second row is significant at the 5% level but H(0,1) and H(1,2) are not. Thus, for US/Japan, we have mixed evidence for the model under Cases 1 and

2.

The fourth panel of Table 6 presents the results for the Canada/UK real exchange rate regressions, for both sample periods, under cases 1 and 2.

Under Case 1, we estimate the coefficients with the theoretically correct signs. The magnitudes of the coefficients exceed 1, but not significantly so. The H(0,1) statistic does not reject the null of deterministic cointegration under either period and both relative prices are significant. However, the H(1,3) statistic in the second row is significant at the 1% level. Under Case 2, none of the H(p,q) statistics is significant at the 5% level. Thus, for Canada/UK, under Case 1, the model is rejected for the free float period, but not for the full period. Under Case 2, we do not reject the model for either sample period.

The fifth panel in Table 6 presents results for the Canada/Japan case. Under Case 1, we estimate α_H and α_F with the theoretically correct signs. The magnitudes of the coefficients exceeds one, but not significantly. Both relative prices are significant and none of the H(p,q) statistics is significant at the 5% level. Under Case 2, the coefficient of the relative price is not significant, although none of the H(p,q) statistics is significant. Thus for Canada/Japan, the model is supported by the data under Case 1.

Finally, the sixth panel in Table 6 presents the results for the UK/Japan real exchange rate regressions under cases 1 and 2. Under Case 1, we estimate α_H and α_F with the correct signs, although the magnitude of α_F is implausible even after taking the standard error into account. The H(0,1) statistic does not reject the model under Case 1, but the H(1,2) and H(1,3) statistics are significant at the 5% level. Under Case 2, the H(1,2) and

H(1,3) statistics are significant at the 10% level in the second and third rows. Thus, for UK/Japan, we do not find evidence in favor of the model under either case.

We plot figures of the actual real exchange rate and the regression of the real exchange rate on relative prices estimated by CCR to see how well our model can explain the long run movements of the real exchange rate for the countries we consider.

Figure 4 shows a plot of the U.S.-U.K. real exchange rate and its regression corresponding to the first row of Table 6. The estimated regression seems to capture the long run tendency for the real exchange rate to fall even though it misses out on the swings of the real exchange rate from 1975 to 1988.

Figure 5 shows a plot of the U.S.-Canada real exchange rate and the estimated regression corresponding to the first row of the second panel of Table 6. The regression seems to completely miss out on the short term and medium term movements of the real exchange rate, but does a reasonable job of capturing the long term tendency for the real exchange rate to fall from 1960 to 1975 and then to rise from 1975 to 1988.

Figure 6 shows a plot of the U.S.-Japan real exchange rate and the regression corresponding to the first row of the third panel of Table 6. This regression captures the tendency of the real exchange rate to fall from 1975 to 1978, to rise from 1978 to the middle of the 1980s, and then to fall.

Thus, with the consumption deflators based data, only for one country pair, out of the six that we consider, do we find that the model is rejected under both cases. For other country pairs we find support for the model under at least one case and sample period.

Because budget share data are available for consumption data, we can address the time-varying weight issue for consumption data. In Table 7, we report CCR results for (6′) when the weighted average of relative prices is constructed. We first take the average budget share over the whole sample and construct the weighted average with the average budget share as the weight for all periods. Then we run Regression (6′) with this fixed weighted average. This is to address the issue of whether or not (1) may be a good approximation. The results for this case are reported in the rows labeled "Fixed Weights." In the first four panels, the point estimate of the coefficient for the weighted average is within two standard errors from one. This result is in favor of (1). In the last two panels, the coefficient is not within two standard errors of one. In order to see the effects of time-varying budget shares, we use the budget share in each quarter to construct the weighted average and report the results in the rows labeled "Varying Weight." For U.S.A/U.K., U.S.A./Japan and Canada/UK, the results here are similar to those for fixed weights, suggesting that the time-varying budget shares are not likely to be important. On the other hand, the p-values for the H(p,q) tests are higher for U.S.A./Canada, Canada/Japan and UK/Japan, suggesting that the time-varying budget shares are likely to be playing an important role in determining the real exchange rate for these country pairs.

V. Conclusions

In this paper, we identify time periods, country pairs and relative price measures for which real exchange rates move together with relative prices in the long run. With the GDP deflator based real exchange rate, for all three country pairs studied, we find evidence of long run comovements for the pre-float sample period, but some evidence against our model for the full sample period. For the consumption deflator based real exchange rates, we find evidence in favor of the model in at least one case and time period for five out of the six country pairs that we consider. On the other hand, for some medium-run swings of real exchange rates, we failed to find such comovements.

Appendix

In this Appendix we explain the data used in this paper. The GDP deflator for the US is from the National Income and Product Accounts whereas that for the UK is from C.H. Feinstein (1972). The Italian GDP deflator was constructed from European Historical Statistics by B.R.Mitchell and the International Financial Statistics. The CPI, WPI and the exchange rate data for all the countries are from Lee (1976) for the sample period 1914 to 1972. These were extended for the US,UK and Italy to 1988 with IFS data.

For consumption deflators, we use relative price data constructed by Stockman and Tesar (1990) from the OECD quarterly national accounts. They decompose private final consumption of commodities by type (durables, semi-durables,nondurables and services) and by object (food, beverages and tobacco; clothing and footwear; gross rent, fuel and power; transport and communication; furniture and household operations; and other goods and services). Throughout this paper, we use services from the classification by type as our proxy for consumption of nontradables. The relative prices of nontradables are constructed from the implicit price deflators of the service and non-service components of consumption. The United States consumption data are from the National Income and Product Accounts: the proximate source is the Citibase. The British and Japanese data on services was not available in seasonally adjusted format. We deseasonalized them using a two sided moving average. The Canadian and U.S. data sets spanned from 1960Q1-1988Q4. The UK data set (after adjustment) spanned from 1962Q1-1988Q3. The Japanese data set (after adjustment) spanned from 1975Q3 - 1987Q3.

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TABLE 1

REAL EXCHANGE RATE REGRESSIONS

Case 1: No Stochastic Cointegration for the Relative Prices

(6) $$ln(E_r(t)) = \theta + \alpha_H ln(q_H(t)) - \alpha_F ln(q_F(t)) + \varepsilon(t),$$

Case 2: Stochastic Cointegration for the Relative Prices

Case 2A: Deterministic Cointegration

(8) $$ln(E_r(t)) = \theta_2 + \beta ln(q_H(t)) + \varepsilon_2(t),$$

Case 2B: No Deterministic Cointegration

(10) $$ln(E_r(t)) = \theta_2 + \mu t + \beta ln(q_H(t)) + \varepsilon_2(t),$$

TABLE 2

TESTS FOR TREND PROPERTIES OF (*THE LN OF*) RELATIVE PRICES

COUNTRY	SAMPLE	J(1,5)[a]	DF[b]	SD(1)[c]	SD(4)[c]	SD(7)[c]
		CPI AND WPI				
U.S.A.	1929-1988	2.5888	-1.4601	-3.0014	-2.3026	-1.7127
U.S.A.	1929-1972	3.2249	-1.3099	-2.3563	-1.6889	-0.5829
ITALY	1929-1988	1.3619	-2.4903	-3.0574	-3.1552	-2.3372
ITALY	1929-1972	1.5851	-2.0609	-2.6346	-2.6670	-2.3188
U.K.	1929-1988	4.2807	-1.4757	-1.7691	-2.1612	-1.9133
U.K.	1929-1972	3.1362	-1.3013	-1.4091	-1.3031	-0.7806
		CONSUMPTION DEFLATORS				
U.S.A.	60Q1-88Q4	6.7332	0.3643	-0.8397	-1.4564	-1.0816
	74Q1-88Q4	2.7793	-2.5736	-2.4912	-2.5875	-2.3663
U.K.	62Q3-88Q3	1.6321	-1.2477	-1.7789	-i.6336	-1.5867
	74Q1-88Q3	2.0674	-2.1379	-2.5355	-2.3968	-2.3480
CANADA	60Q1-88Q4	2.7759	-1.8040	-1.1620	-1.4470	-1.6332
	74Q1-88Q4	3.4503	-2.3262	-2.2931	-3.3246	-3.7976[d]
JAPAN	75Q3-87Q3	2.4443	-0.9872	-1.6919	-3.3682	-2.0486

[a]Critical values at the 1% , 5% and 10% levels of significance are 0.123 , 0.295 and 0.452 respectively. These are from Park and Choi (1988).

[b]DF denotes the Dickey-Fuller test (with trend). The 5% critical value for the sample 1929-1988 is -3.49 and that for the sample 1929-1972 is -3.52. The 5% critical value for the samples 60Q1-88Q4 and 62Q3-88Q3 is -3.45, while that for samples 74Q1-88Q4 and 75Q3-87Q3 is -3.49.

[c]SD(r) denotes the Said-Dickey(with trend) test with r lags. The 95% critical values are (approximately) the same as those for the Dickey-Fuller tests.

[d]Significant at the 5% level.

TABLE 3

CANONICAL COINTEGRATING REGRESSIONS FOR RELATIVE PRICES

COUNTRIES	SAMPLE	b_{CCR}^{a}	$H(0,1)^{b}$	$H(1,2)^{b}$	$H(1,3)^{b}$
		CPI AND WPI			
U.S.A./ITALY	1929-1988	0.3568 (0.1697)	1.9860 (0.1586)	1.5535 (0.2126)	1.5628 (0.4577)
	1929-1972	0.3506 (0.1797)	2.1015 (0.1471)	1.6783 (0.1951)	1.6803 (0.4316)
U.S.A./U.K.	1929-1988	0.4328 (0.1363)	0.1539 (0.6948)	0.0239 (0.8770)	2.1087 (0.3484)
	1929-1972	0.3693 (0.1069)	0.1806 (0.6709)	0.1368 (0.7114)	1.7816 (0.4103)
U.K./ITALY	1929-1988	0.1477 (0.2273)	1.2793 (0.2580)	4.3843 (0.0363)	6.0706 (0.0481)
	1929-1972	0.2119 (0.2456)	1.9537 (0.1622)	4.4612 (0.0347)	4.6395 (0.0983)
		CONSUMPTION DEFLATORS			
U.S.A./U.K.	62Q3-88Q3	0.2542 (0.3006)	0.0158 (0.8998)	0.1471 (0.7013)	0.7351 (0.6924)
	74Q1-88Q3	0.8739 (0.2466)	0.7460 (0.3877)	3.8106 (0.0509)	3.8344 (0.1470)
U.S.A./CANADA	60Q1-88Q4	0.4966 (0.3892)	0.7266 (0.3940)	1.2101 (0.2713)	2.1899 (0.3345)
	74Q1-88Q4	1.8992 (0.2689)	9.5112 (0.0020)	0.2997 (0.5840)	1.2978 (0.5226)
U.S.A./JAPAN	75Q3-87Q3	1.4548 (0.1273)	0.2405 (0.6238)	0.5003 (0.4794)	18.0539 (0.0001)

CANADA/U.K.	62Q3-88Q3	0.4605 (0.0863)	1.2267 (0.2680)	0.4361 (0.5090)	0.9403 (0.6249)
	74Q1-88Q3	0.0732 (0.0854)	0.1069 (0.7437)	0.8892 (0.3457)	3.2544 (0.1965)
CANADA/JAPAN	75Q3-87Q3	-0.0189 (0.0523)	0.1802 (0.6712)	1.6074 (0.2048)	1.6306 (0.4425)
U.K./JAPAN	75Q3-87Q3	1.2072 (0.4945)	7.1191 (0.0076)	0.0181 (0.8930)	1.2507 (0.8822)

[a]For each row, b_{CCR} is the coefficient on the log of the relative price of the country listed second. The standard errors are in parentheses.

[b]P-values are in parentheses.

TABLE 4

TESTS FOR THE NULL OF NO STOCHASTIC COINTEGRATION BETWEEN RELATIVE PRICES

COUNTRIES	SAMPLE	I(1,5)[a]	DF[b]	SD(1)[c]	SD(4)[c]	SD(7)[c]
			CPI AND WPI			
U.S.A./ ITALY	1929-1988	1.2693	-2.5954	-3.6875	-2.0573	-0.8011
	1929-1972	1.1805	-2.7753	-4.1244[d]	-2.1410	-0.6699
U.S.A./ U.K.	1929-1988	1.0624	-2.2368	-3.9654[d]	-3.2142	-2.1597
	1929-1972	0.9181	-2.5608	-3.8178	-3.7006	-1.4205
U.K./ITALY	1929-1988	3.5830	-2.0592	-2.1793	-1.8383	-1.3107
	1929-1972	1.3635	-2.5901	-2.7633	-2.1432	-1.1109
			CONSUMPTION DEFLATORS			
U.S.A./ U.K.	62Q3-88Q3	3.3047	-1.7613	-2.4943	-2.7173	-2.4873
	74Q1-88Q3	2.3934	-1.2217	-1.7266	-2.6448	-2.4652
U.S.A./ CANADA	60Q1-88Q4	5.9962	-1.6612	-1.6827	-3.0371	-3.1921
	74Q1-88Q4	7.2955	-3.1384	-3.2469	-1.5883	-1.8632
U.S.A./ JAPAN	75Q3-87Q3	10.5824	-0.8342	-1.5202	-2.0881	-2.7918
CANADA/ U.K.	62Q3-88Q3	2.1604	-2.0238	-1.6954	-2.1923	-2.3210
	74Q1-88Q3	3.4409	-2.0187	-1.8402	-1.9595	-3.2896
CANADA/ JAPAN	75Q3-87Q3	8.7787	-1.5139	-1.9800	-1.9566	-2.6749
U.K./ JAPAN	75Q3-87Q3	2.0382	-2.3578	-2.8690	-2.2756	-2.2043

NOTE: The I(p,q) tests reported here use (the ln of) the relative price of the first country as the regressand and that of the second country as the regressor .

[a]The 1%, 5% and 10% critical values for the I(1,5) test are 0.1027, 0.2506 and 0.4984, respectively. These are taken from Park, Ouliaris and

Choi (1988).

[b]DF denotes the Dickey-Fuller test (with trend). The 5% critical values for the sample 1929-1988 are -3.94, those for 1929-1972 are -4, those for 60Q1-88Q4 or 62Q3-88Q3 are -3.86, those for 75Q3-87Q3 or 74Q1-88Q4 are -3.95.

[c]SD(r) denotes the Said-Dickey test with r lags. The critical values are (approximately) the same as those for the Dickey-Fuller tests.

[d]Significant at the 5 % level.

TABLE 5

CANONICAL COINTEGRATING REGRESSIONS FOR THE REAL EXCHANGE RATE BASED ON GNP DEFLATORS

SAMPLE	$b^{a}_{1,CCR}$	$b^{a}_{2,CCR}$	$H(0,1)^{b}$	$H(1,2)^{b}$	$H(1,3)^{b}$
		U.K./ITALY			
1929-1988	0.7627	-0.8523	0.0017	5.7326	6.0517
	(0.1887)	(0.1528)	(0.9670)	(0.0166)	(0.0485)
1929-1972	0.6885	-0.4892	0.0608	1.5886	3.6573
	(0.1545)	(0.1347)	(0.8052)	(0.2075)	(0.1606)
1929-1988	0.5213	......	4.5645	0.0218	0.0499
	(0.3278)	......	(0.0326)	(0.8826)	(0.9753)
1929-1972	0.5248	......	2.9338	0.1977	1.7946
	(0.2306)	......	(0.0867)	(0.6565)	(0.4077)
1929-1988	......	-0.8695	2.2733	0.8875	3.8114
	......	(0.2285)	(0.1316)	(0.3461)	(0.1487)
1929-1972	......	-0.5597	3.6378	1.9005	2.8347
	......	(0.1964)	(0.0565)	(0.1680)	(0.2423)
		U.S.A./U.K.			
1929-1988	2.0226	-2.1062	5.2406	0.1544	13.8012
	(0.3887)	(0.2750)	(0.0220)	(0.6943)	(0.0010)
1929-1972	1.9517	-1.9110	1.4457	1.4622	2.9926
	(0.3651)	(0.2271)	(0.2292)	(0.2266)	(0.2240)
1929-1988	-1.2539	......	0.1077	4.1501	4.4355
	(0.7149)	......	(.07427)	(0.0416)	(0.1088)
1929-1972	-1.3375	......	1.6564	0.3039	1.0241
	(0.4039)	......	(0.1981)	(0.5815)	(0.5992)
1929-1988	......	-1.4286	0.2205	2.4968	9.4132
	......	(0.3155)	(0.6386)	(0.1160)	(0.0090)
1929-1972	......	-1.2870	0.0002	3.1044	3.2585
	......	(0.2582)	(0.9890)	(0.0781)	(0.1961)

TABLE 5 - *CONTINUED*

CANONICAL COINTEGRATING REGRESSIONS FOR THE REAL EXCHANGE RATE

SAMPLE	$b_{1,CCR}$ [a]	$b_{2,CCR}$ [a]	H(0,1) [b]	H(1,2) [b]	H(1,3) [b]
		U.S.A./ITALY			
1929-1988	0.2511 (0.4956)	-0.8212 (0.2763)	0.9684 (0.3251)	8.9591 (0.0028)	10.1460 (0.0063)
1929-1972	-0.3023 (0.2159)	-0.1875 (0.1150)	0.0002 (0.9885)	3.5755 (0.0586)	3.6907 (0.1580)
1929-1988	-1.0694 (0.5037)	....	5.1415 (0.0234)	1.7240 (0.1892)	2.3396 (0.3104)
1929-1972	-0.5856 (0.2748)	....	0.1583 (0.6907)	0.4846 (0.4863)	1.9932 (0.3691)
1929-1988	......	-0.9072 (0.2499)	0.3564 (0.5505)	4.4569 (0.0347)	5.8731 (0.0560)
1929-1972	......	-0.3383 (0.1257)	1.3591 (0.2437)	0.6190 (0.4314)	0.7503 (0.6872)

[a] $b_{1,CCR}$ is the coefficient on (the ln of) the first country's relative price. $b_{2,CCR}$ is the coefficient on (the ln of) the second country's relative price. Standard errors are in parentheses.

[b] P-values are in parentheses.

TABLE 6

CANONICAL COINTEGRATING REGRESSIONS FOR THE REAL EXCHANGE RATE BASED ON TOTAL CONSUMPTION DEFLATORS

SAMPLE	$b^{a}_{1,CCR}$	$b^{a}_{2,CCR}$	$H(0,1)^{b}$	$H(1,2)^{b}$	$H(1,3)^{b}$
		U.S.A./U.K.			
62Q3-88Q3	0.0212	-0.6578	0.0187	3.1050	3.1067
	(0.7678)	(0.3080)	(0.8912)	(0.0780)	(0.2115)
74Q1-88Q3	2.0852	-3.1441	0.0000	0.1557	1.4909
	(0.9083)	(0.8558)	(0.9951)	(0.6932)	(0.4745)
62Q1-88Q3	-1.7619	......	1.8676	0.8329	1.0597
	(0.8952)	......	(0.1717)	(0.3614)	(0.5887)
74Q1-88Q3	-0.9187	......	1.2141	2.0980	2.3245
	(0.7398)	......	(0.2705)	(0.1475)	(0.3128)
62Q1-88Q3	......	-0.4972	0.0358	0.4288	2.1543
		(0.2345)	(0.8500)	(0.5126)	(0.3406)
74Q1-88Q3	......	-1.0132	0.0425	1.1415	2.8578
		(0.7089)	(0.8367)	(0.2853)	(0.2396)

TABLE 6 - *Continued*

SAMPLE	$b^{a}_{1,CCR}$	$b^{a}_{2,CCR}$	$H(0,1)^{b}$	$H(1,2)^{b}$	$H(1,3)^{b}$
		U.S.A./CANADA			
60Q1-88Q4	0.3587	-0.2732	5.0720	2.4305	2.4511
	(0.1661)	(0.1153)	(0.0243)	(0.1190)	(0.2936)
74Q1-88Q4	0.4917	-1.2707	0.0675	0.3397	0.4603
	(0.1185)	(0.2281)	(0.7950)	(0.5600)	(0.7944)
60Q1-88Q4	0.3433	......	0.4929	0.2109	0.5308
	(0.2535)	......	(0.4826)	(0.6460)	(0.7669)
74Q1-88Q4	0.5292	......	1.9679	1.8281	2.0151
	(0.5492)	......	(0.1607)	(0.1763)	(0.3651)
60Q1-88Q4	......	0.0693	6.2834	0.0987	0.2662
		(0.1408)	(0.0122)	(0.7533)	(0.8753)
74Q1-88Q4	......	0.1551	5.0740	0.1482	4.5384
		(0.3647)	(0.0243)	(0.7003)	(0.1034)

		U.S.A./JAPAN			
75Q3-87Q3	2.3425	-4.0899	5.1806	0.8793	2.8708
	(0.6421)	(0.8700)	(0.0228)	(0.3484)	(0.2380)
75Q3-87Q3	-0.8891	......	0.4238	0.9575	6.2738
	(0.4829)	......	(0.5150)	(0.3278)	(0.0431)
75Q3-87Q3	......	-1.4798	2.9576	0.0012	0.7949
		(1.2827)	(0.0855)	(0.9721)	(0.6720)

TABLE 6 - *Continued*

SAMPLE	$b^{a}_{1,CCR}$	$b^{a}_{2,CCR}$	$H(0,1)^{b}$	$H(1,2)^{b}$	$H(1,3)^{b}$
			CANADA/U.K.		
62Q3-88Q3	1.8225 (0.5172)	-1.1927 (0.2848)	0.1094 (0.7408)	0.2925 (0.5886)	0.8392 (0.6573)
74Q1-88Q3	2.5340 (0.5624)	-0.8380 (0.3967)	0.5780 (0.4463)	1.2007 (0.2732)	13.8439 (0.0010)
62Q1-88Q3	-0.5500 (0.2939)	....	1.8603 (0.1726)	0.4001 (0.5270)	1.0006 (0.6063)
74Q1-88Q3	1.1699 (0.5893)	....	1.3587 (0.2437)	0.2106 (0.6463)	5.6137 (0.0604)
			CANADA/JAPAN		
75Q3-87Q3	1.4168 (1.0673)	-1.5748 (0.8383)	1.0835 (0.2979)	2.8531 (0.0912	3.5073 (0.1731)
75Q3-87Q3	-0.4542 (1.0032)	....	2.2358 (0.1348)	0.5226 (0.4697)	3.4350 (0.1795)
			U.K./JAPAN		
75Q3-87Q3	1.0664 (0.3505)	-2.1004 (0.2546)	0.0736 (0.7861)	6.3741 (0.0116)	9.1599 (0.0102)
75Q3-87Q3	-0.9632 (1.6739)	....	0.0042 (0.9482)	3.7804 (0.0518)	4.7993 (0.0907)
75Q3-87Q3	......	-1.2875 (0.2344)	1.0546 (0.3044)	6.5376 (0.0106)	6.8091 (0.0332)

[a] $b_{1,CCR}$ is the coefficient on (the ln of) the first country's relative price. $b_{2,CCR}$ is the coefficient on (the ln of) the second country's relative price. Standard errors are in parentheses.

[b] P-values are in parentheses.

TABLE 7

CANONICAL COINTEGRATING REGRESSIONS FOR THE REAL EXCHANGE RATE BASED ON TOTAL CONSUMPTION DEFLATORS

SAMPLE	b_{CCR} [a]	H(0,1) [b]	H(1,2) [b]	H(1,3) [b]
		U.S.A./U.K.		
62Q3-88Q3 Fixed Weight	3.2015 (0.9801)	1.5367 (0.2151)	0.5679 (0.4511)	1.2025 (0.5481)
62Q3-88Q3 Varying Weight	3.7697 (1.2710)	2.1412 (0.1434)	1.0657 (0.3019)	1.4696 (0.4796)
		U.S.A./JPN.		
75Q3-87Q3 Fixed Weight	2.3608 (0.9242)	1.5316 (0.2159)	1.1048 (0.2932)	5.1757 (0.1594)
75Q3-87Q3 Varying Weight	1.6228 (0.8392)	1.7209 (0.1896)	0.9868 (0.3205)	5.9210 (0.0518)
		U.S.A./CAN.		
60Q1-88Q4 Fixed Weight	0.6767 (0.2682)	6.5866 (0.0103)	4.7653 (0.0290)	8.7018 (0.0129)
60Q1-88Q4 Varying Weight	0.6016 (0.2150)	4.1228 (0.0423)	3.3938 (0.0654)	5.8161 (0.0546)
		CAN./U.K.		
62Q3-88Q3 Fixed Weight	3.0874 (0.7633)	0.0904 (0.7637)	0.1306 (0.7178)	0.6922 (0.7074)
62Q3-88Q3 Varying Weight	3.2280 (0.8373)	0.0080 (0.9287)	0.0109 (0.9167)	1.2670 (0.5307)

TABLE 7 - *Continued*

SAMPLE	b_{CCR}[a]	H(0,1)[b]	H(1,2)[b]	H(1,3)[b]
		CAN./JPN.		
75Q3-87Q3 Fixed Weight	4.9891 (1.1296)	0.3963 (0.5290)	0.9947 (0.3186)	4.4146 (.01100)
75Q3-87Q3 Varying Weight	4.7691 (1.1217)	0.8682 (0.3514)	0.7631 (0.3824)	4.5665 (0.1020)
		U.K./JPN.		
75Q3-87Q3 Fixed Weight	11.7006 (2.5962)	0.0456 (0.8310)	7.7508 (0.0054)	7.8011 (0.1594)
75Q3-87Q3 Varying Weight	1.6228 (0.8392)	1.7209 (0.1896)	0.9868 (0.3205)	5.9210 (0.0518)

[a] b_{CCR} denotes the coefficient on (the ln of the) weighted average of the relative prices. Standard Errors are in parentheses.

[b] P-values are in parentheses.

Plot of the (ln of the) UK(-Italy) Real Exchange Rate and the CCR Forecast with both Relative Prices as Regressors and a War Dummy for Italy

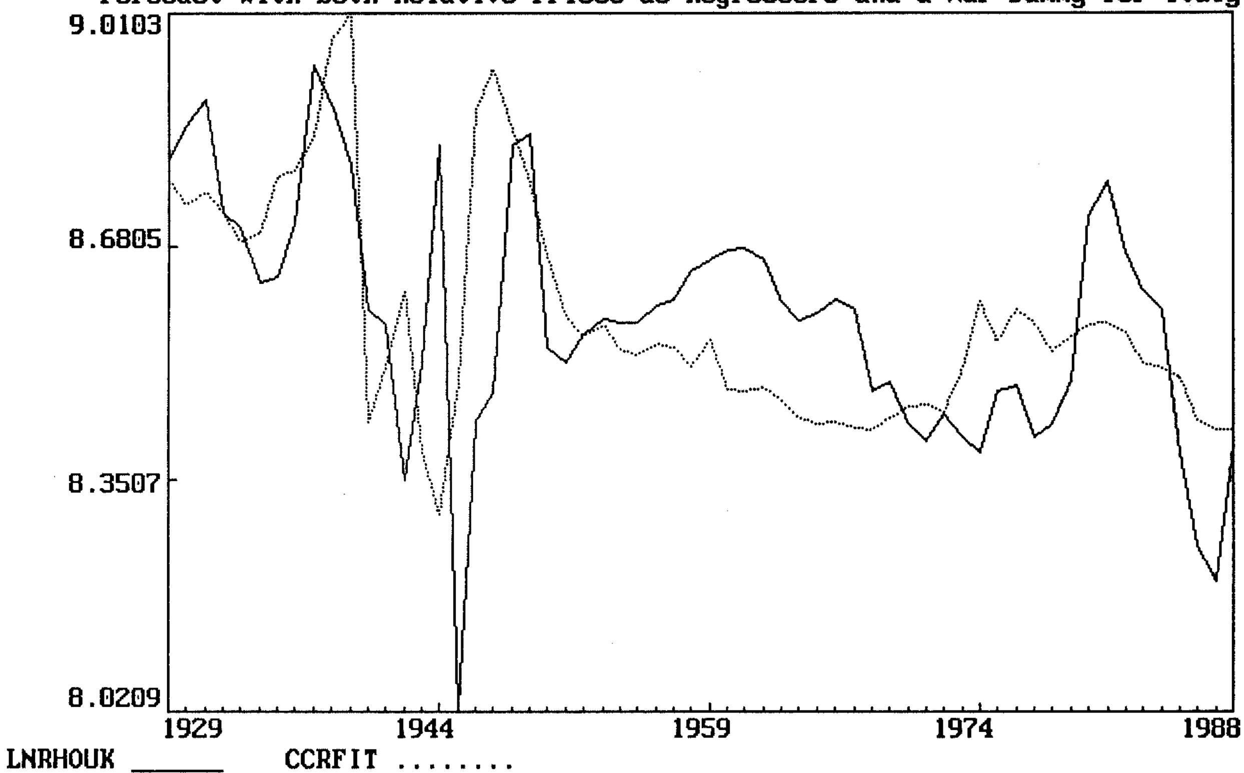

FIGURE 1

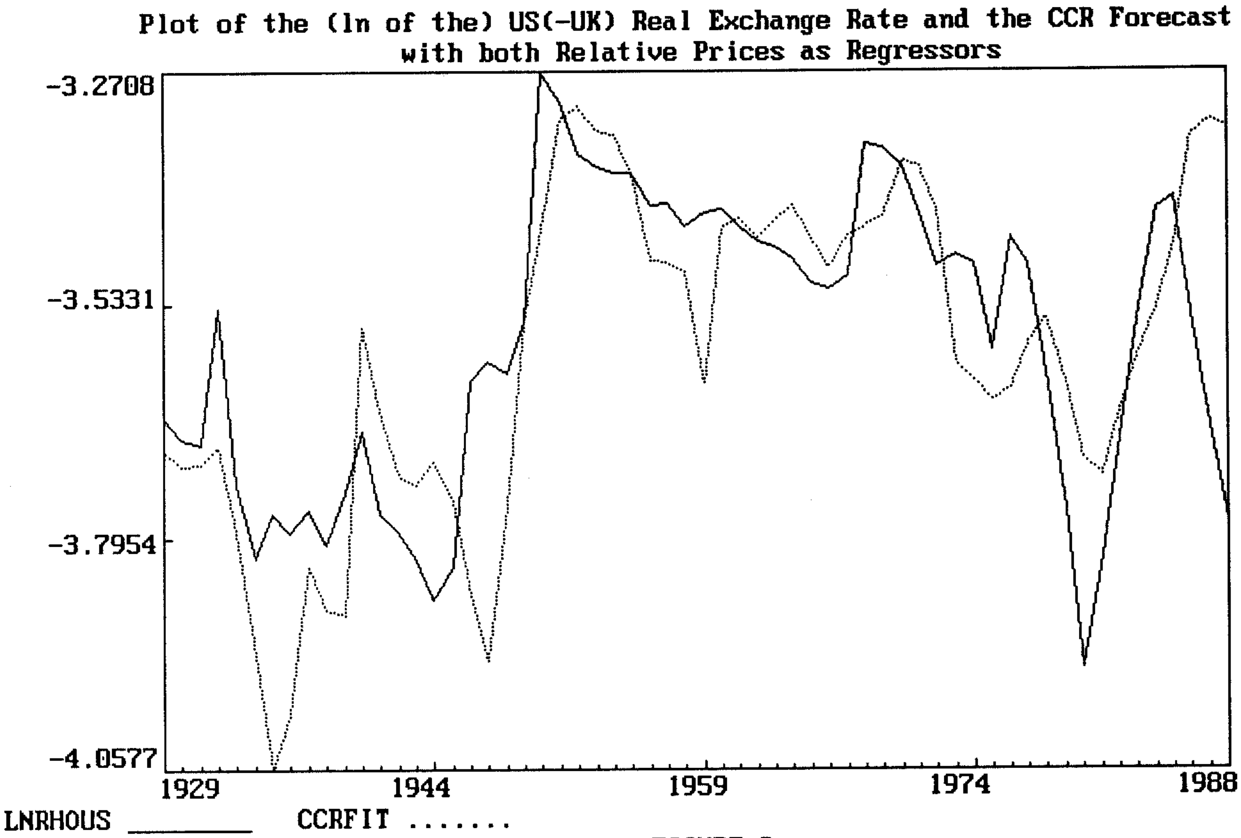
Plot of the (ln of the) US(-UK) Real Exchange Rate and the CCR Forecast
with both Relative Prices as Regressors
-3.2708
-3.5331
-3.7954
-4.0577
1929
1944
1959
1974
1988
LNRHOUS ________
CCRFIT ....

FIGURE 2

Plot of the (ln of the) US(-Italy) Real Exchange Rate and the CCR
Forecast with both Relative Prices as Regressors and a War Dummy for Italy

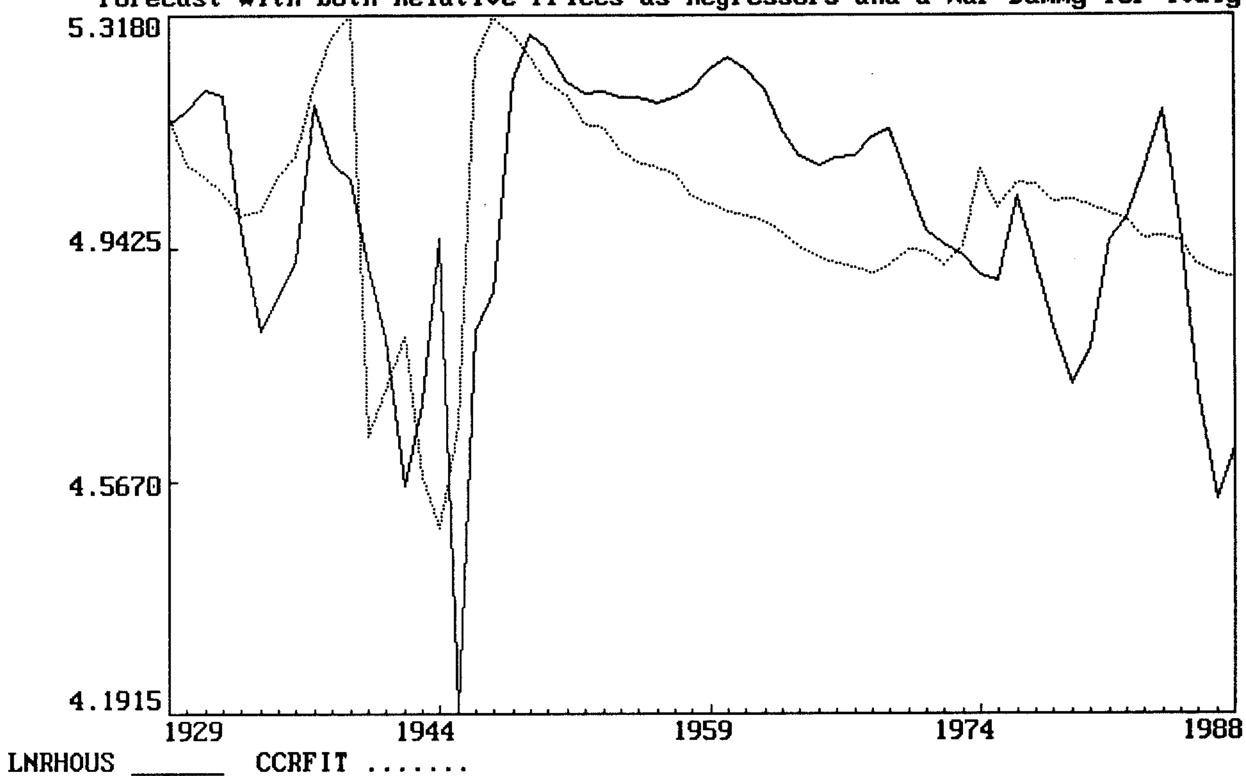

FIGURE 3

Plot of the (ln of the) US(-UK) Real Exchange Rate and the CCR Forecast with both Relative Prices as Regressors. (The UK data was deseasonalized)

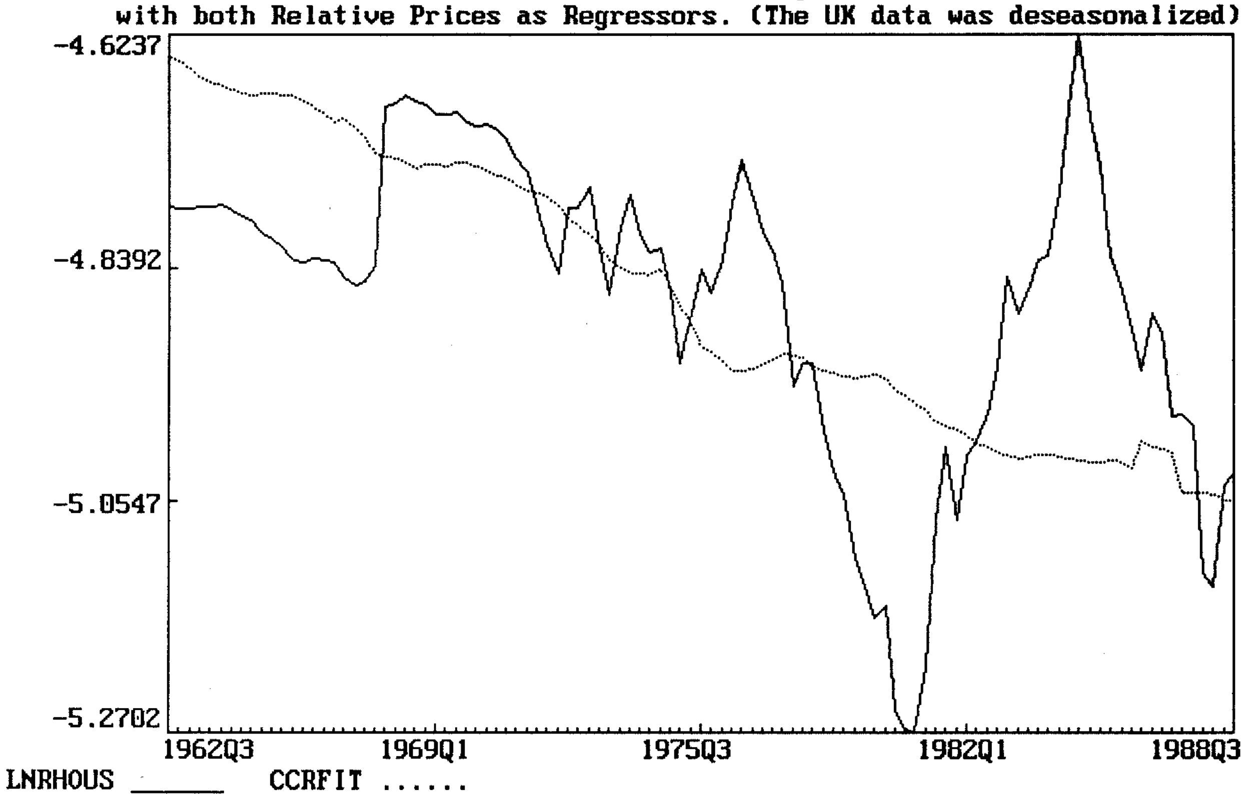

FIGURE 4

Plot of the (ln of the) US(-Canada) Real Exchange Rate and the CCR Forecast with both Relative Prices as Regressors.

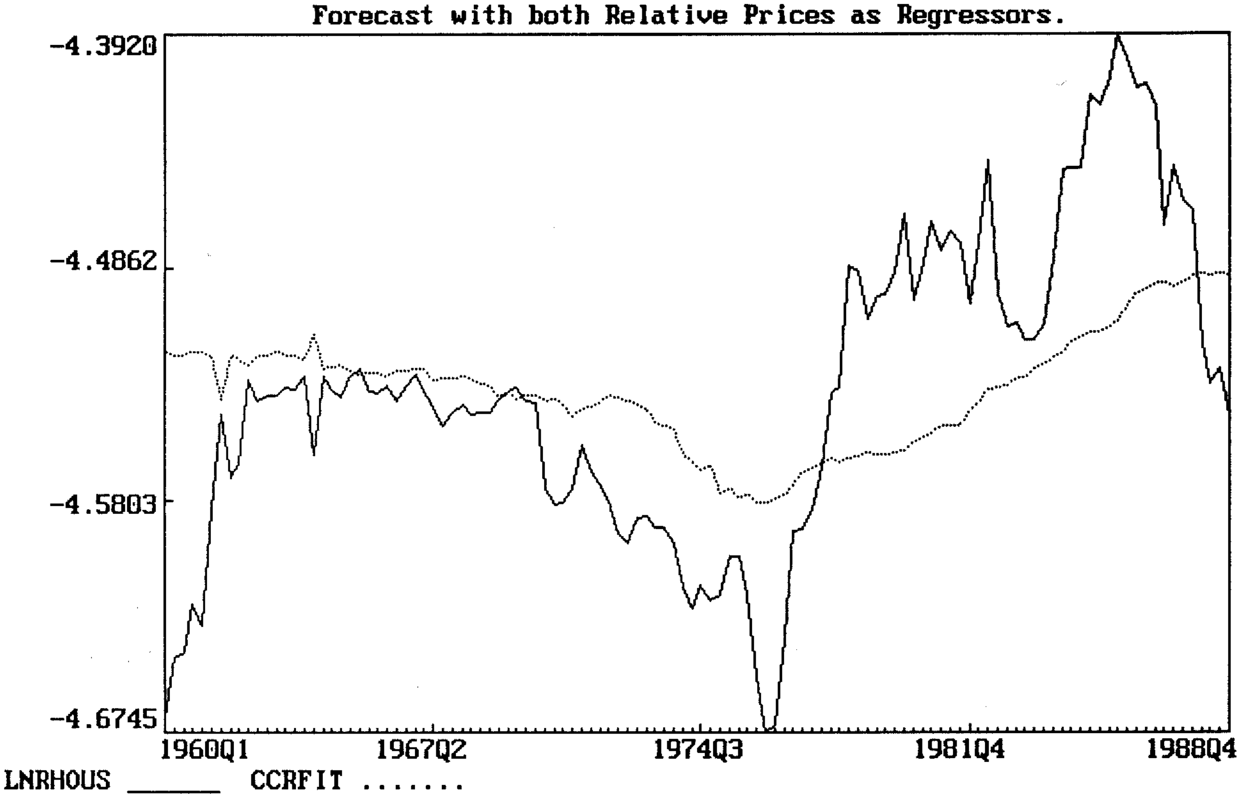

FIGURE 5

Plot of the (ln of the) US(-Japan) Real Exchange Rate and the CCR Forecast with both Relative Prices as Regressors.

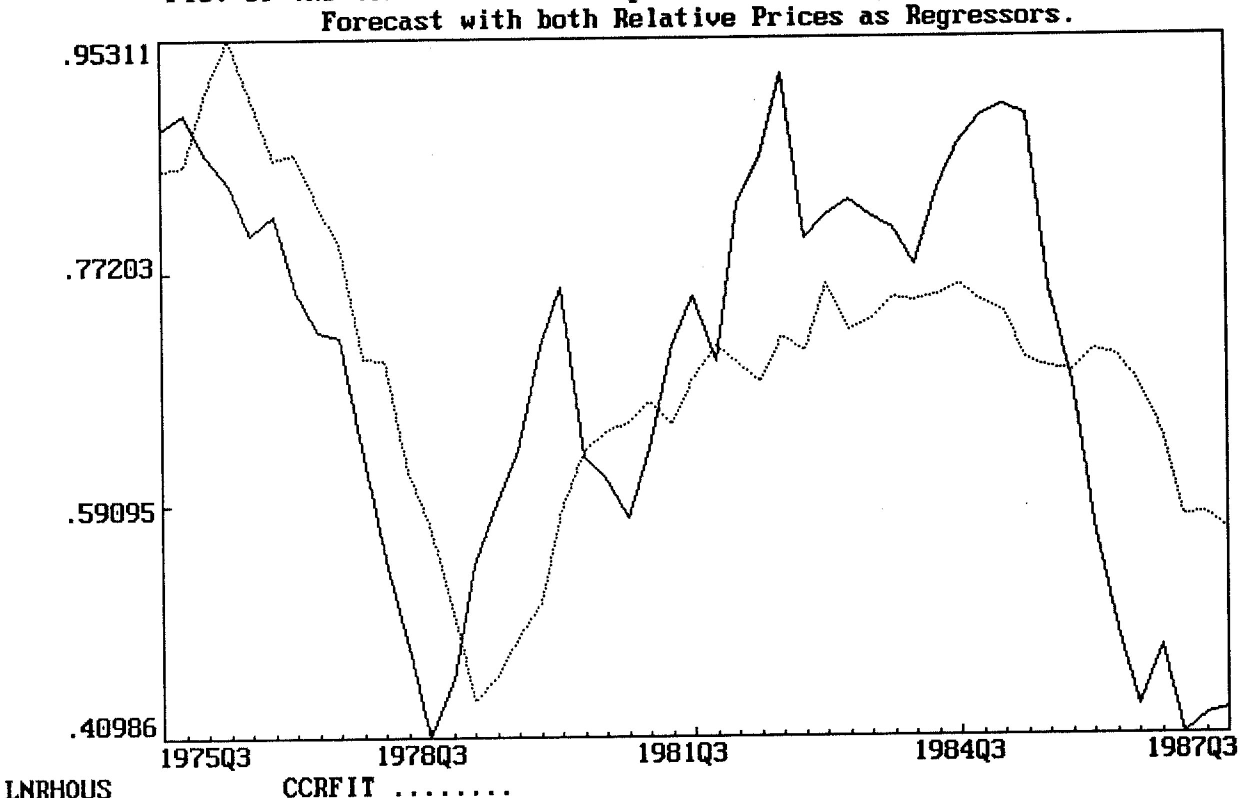

FIGURE 6