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Bank of Canada Banque du Canada

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CONTENTS

Abstract

We examine whether a link exists between oil price shocks and the U.S. real effective exchange rate. The results show that the two variables appear to be cointegrated and that causality runs from oil prices to the exchange rate and not vice versa. The single-equation error-correction model linking these two variables is stable and captures much of the in- and out-of-sample movement in the exchange rate in dynamic simulations. Finally, tests we present show that the error-correction model has significant post-sample predictive ability for both the size and sign of changes in the real effective exchange rate. The results suggest that oil prices may have been the dominant source of persistent real exchange rate shocks over the post–Bretton Woods period and that energy prices may have important implications for future work on exchange rate behaviour.

Résumé

Dans la présente étude, les auteurs cherchent à déterminer s'il existe un lien entre les chocs pétroliers et le taux de change effectif réel du dollar américain. D'après les résultats obtenus, les deux variables semblent cointégrées et la causalité s'exerce des prix du pétrole vers le taux de change et non en sens inverse. Le modèle de correction des erreurs pour le taux de change est stable et capte la plupart des mouvements à l'intérieur et à l'extérieur de l'échantillon dans les simulations dynamiques du taux de change. Enfin, les tests effectués indiquent que le modèle de correction des erreurs réussit à prévoir de façon significative l'ampleur ainsi que le sens des fluctuations du taux de change effectif réel en dehors de l'échantillon. Les résultats donnent à penser que les cours du pétrole peuvent avoir été la principale source des chocs persistants qu'a subis le taux de change réel depuis l'effrondrement du régime de Bretton Woods et que, à l'avenir, les prix de l'énergie pourraient occuper une place importante dans les modèles visant à expliquer l'évolution du taux de change.

1 Introduction

In this paper we explore the ability of real oil prices to explain movements in the U.S. real effective exchange rate. The potential importance of the price of oil for exchange rate movements has been noted by, *inter alia*, Krugman (1983a, 1983b) and Golub (1983). Although the models these authors present are intuitively appealing, to our knowledge very little work has carefully examined whether a link exists between oil prices and the U.S. exchange rate. Advances over the last few years in econometric techniques for analysing nonstationary series also make a re-examination of the stylized facts seem worthwhile. Furthermore, an examination of the relationship between oil prices and exchange rates can contribute to several recent areas of macroeconomic research.

First, such a study could contribute evidence to an empirical literature that suggests energy price changes can account for innovations in major U.S. macroeconomic variables. For instance Hamilton (1983) finds that major oil price increases have preceded almost all post-World War II recessions in the United States. Dotsey and Reid (1992) show that the measure of U.S. monetary policy proposed by Romer and Romer (1989) is coincident with several major oil price shocks, and that its explanatory power for output vanishes when oil prices are included in the analysis. Loungani (1986) finds that a significant fraction of the dispersion of employment growth across different industries may be attributed to oil price shocks. To our knowledge, comparable studies of their effects on the exchange rate have not been done.

Second, the results would also contribute to a real-business cycle (RBC) literature that aims to test the ability of simulation models to capture macroeconomic stylized facts. Recently McCallum (1989) and others have emphasized the importance of including oil price shocks in RBC models. Researchers such as Praschnik and Costello (1992a), Kim and Loungani (1992) and Finn (1991) have postulated a specific structural model driven by energy price shocks and have examined the extent to which the model captures important stylized facts. The problems of expanding such models to an international setting has been explored by Mendoza (1991), Praschnik and Costello (1992b) and Backus, Kehoe and Kydland (1992). Evidence of a systematic relationship between oil prices and the exchange rate would provide additional stylized facts for such models to capture.

Third, this study can contribute to investigations into the link between real exchange rate fluctuations and real international interest rate differentials. A series of recent papers including Campbell and Clarida (1987), Meese and Rogoff (1988), Baxter (1993), and Edison and Pauls (1993) has documented the failure of real interest parity and the lack of a cointegrating relationship between real interest rates and exchange rates. This has led some researchers to suggest that an unidentified real factor may be causing persistent shifts in real equilibrium exchange rates. Several of the above authors search for such a factor, but tests of various series including measures of fiscal policy and external indebtedness fail to produce evidence of cointegration. Evidence of a cointegrating relationship between oil prices and the exchange rate would offer a potential explanation for failure of real interest rate parity.

Finally, Meese and Rogoff (1983) note that existing monetary and portfolio balance exchange rate models have virtually no ability to forecast exchange rate changes out-of-sample (that is, they forecast no better than a random walk). We therefore test whether oil prices can forecast exchange rate changes out-of-sample better than a random walk. If so, this would suggest that examining energy prices as a determinant of the real exchange rate may a be fruitful tack.

To preview our results we find (i) there is significant evidence that both the real U.S. price of oil and the U.S. real effective exchange rate contain a unit root; (ii) these two variables are cointegrated; (iii) the price of oil Granger-causes the exchange rate and not vice versa; and (iv) a stable dynamic model using lagged real oil prices forecasts significantly better out-of-sample than a random-walk model.

The structure of the paper is as follows. In the next section we describe the data and examine its time-series properties and the long-run explanatory power of oil prices for the real exchange rate. Section 3 investigates the apparent causal relationships between exchange rates and oil prices. In Section 4 we begin with an unrestricted error-correction model that is sequentially reduced until an error-correction model with reasonable properties is derived. Section 5 presents evidence that shows that our simple model forecasts significantly better than a random walk in out-of-sample forecasting exercises. Section 6 offers some concluding remarks.

2 Data description and cointegration results

The data we use are monthly observations of the real effective (that is, trade-weighted) value of the U.S. dollar and U.S. real price of oil over the 1972M2 to 1993M1 sample period. The real U.S. dollar effective exchange rate is defined in terms of the currencies of 15 other industrial countries and deflated by wholesale price indexes, as calculated by Morgan Guaranty. The oil price series is the U.S. dollar spot price of West Texas Intermediate Crude Oil deflated by the U.S consumer price index. Both variables are used in logarithmic form. We should note that our conclusions appear robust to the use of different price deflators and measures of the U.S. real effective exchange rate. This is not really surprising, as the different measures are highly correlated. For instance, the Morgan Guaranty 40-country and the International Monetary Fund's Multilateral Exchange Rate Model (MERM) measures of the U.S. real effective exchange rate are over 98 per cent correlated with the Morgan Guaranty 15-country measure. Accordingly, we simply chose the series that gave us the longest time span.

It is now common practice to examine the time-series properties of economic data as a guide to subsequent multivariate modelling and inference. If we find that the variables are integrated of order greater than or equal to one, then it could be the case that these variables are cointegrated (see Engle and Granger 1987). This requires non-standard distributional theory in order to perform valid statistical inference. Hence, we begin by testing the null hypothesis of an autoregressive unit root using various tests, including those suggested by Dickey and Fuller (1979, 1981), and Phillips and Perron (1988). The augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) tests are based on the test regression

$$x_t = \lambda_0 + \lambda_1 x_{t-1} + \sum_{i=1}^{k} \Psi_i \Delta x_{t-i} + \varepsilon_t \qquad (1)$$

where Δ is the first-difference operator, x is the variable under consideration, λs and Ψs are parameters to be estimated, and ε is an error term. The ADF test is based on the estimated parameter $\hat{\lambda}_1$ and its corresponding t-statistic.[1] The PP approach, which allows for the presence of autocorrelation and conditional heteroscedasticity in the error term, is based on test regression (1), except that the Ψs are set equal to zero. The corresponding non-parametrically modified t-statistic of $\hat{\lambda}_1$ is then used to test the unit-root hypothesis.[2] For both tests a t-statistic larger in absolute value than the critical values results in a rejection of the null hypothesis of unit root in favour of the stationarity alternative. The test statistic distribution is non-standard, so the critical values are calculated from the response surface estimates of MacKinnon (1991).

Monte Carlo evidence by Schwert (1987) indicates that these standard unit-root tests often have weak power against persistent alternatives, so we also use the Kwiatkowski, Phillips, Schmidt and Shin (1992) test for the null hypothesis of stationarity.[3] The Kwiatkowski, Phillips, Schmidt and Shin (KPSS) test is based on the test statistic

$$\eta = \frac{T^{-2} \sum_{t=1}^{T} S_t^2}{s} \tag{2}$$

where

$$S_t = \sum_{i=1}^{t} e_i \tag{3}$$

1. The test also requires us to choose k. We follow the suggestion of Hall (1989) by beginning with an overparameterized lag structure ($k = 13$) and sequentially reducing the structure by deleting lags until the t-statistic corresponding to the last lag is statistically significant at the 5 per cent level.
2. To construct the PP test statistic we need an estimate of the long-run variance; to this end we use the vector autoregressive (VAR) prewhitened quadratic spectral kernel estimator advocated by Andrews and Monahan (1992).
3. Monte Carlo analysis by Amano and van Norden (1992) shows that this type of joint-testing procedure can substantially reduce the frequency of incorrect conclusions about a series' data generation process.

e_i is the residual from the regression of x on a constant, s is Andrews and Monahan's (1992) VAR prewhitened quadratic spectral kernel long-run variance estimator, and T is the sample size. The distribution of this statistic is non-standard and critical values are taken from Kwiatkowski et al. (1992).

The unit-root and stationarity test statistics are reported in Table 1 (p. 19).[4] For the both the U.S. real effective exchange rate and price of oil, the unit-root tests are unable to find any significant evidence of stationarity, while the KPSS test finds evidence of a unit root in the series that is significant at the 5 per cent level. We therefore conclude that the real exchange rate and the price of oil can be well characterized as I(1) processes.

Next we examine whether the real exchange rate and the price of oil form a cointegrating relationship. The tests for cointegration allow us to gauge the adequacy of specifying the long-run value of the real exchange rate simply as a function of the price of oil. If the long-run real exchange rate is determined by factors other than those associated with the price of oil, then their omission should in theory prevent us from finding significant evidence of cointegration. On the other hand, evidence of cointegration implies that the price of oil captures the dominant source of persistent innovations in the real effective exchange over this period.

To test for cointegration we use the systems approach developed by Johansen and Juselius (1990). Recent Monte Carlo work by Gonzalo (1989) suggests that the Johansen and Juselius (JJ) approach performs better than both single-equation and alternative multivariate methods in detecting cointegration. The starting point of the JJ analysis is a VAR specification for the $n \times 1$ vector of I(1) variables, namely,

$$X_t = \mu + A_1 X_{t-1} + \ldots + A_k X_{t-k} + u_t \tag{4}$$

where u_t is assumed to be an independent and identically distributed Gaussian process. Note that

4. We tested each series for statistically significant evidence of drift using the general-to-specific testing strategy proposed by Perron (1988). No evidence of drift is found for any of the variables.

we can rewrite equation (4) as

$$\Delta X_t = \mu + \Gamma_1 \Delta X_{t-1} + \ldots + \Gamma_{k-1} \Delta X_{t-k+1} + \Pi X_{t-k} + u_t \tag{5}$$

where

$$\Gamma_i = -(I - A_1 - \ldots - A_i) \qquad i = 1, \ldots, k-1$$

$$\Pi = -(I - A_1 - \ldots - A_k)$$

By rewriting equation (4) into equation (5) we are able to summarize the long-run information in X_t by the long-run impact matrix, Π, and it is the rank of this matrix that determines the number of cointegrating vectors.[5] Note that under the null hypothesis of r $(0 < r < n)$ cointegrating vectors, Π can be factored as $\Pi = \alpha\beta^T$ where α and β are $n \times r$ matrices. Therefore under the null we can write the process for X_t as

$$\Delta X_t = \mu + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \alpha\beta^T X_{t-k} + u_t \tag{6}$$

Johansen and Juselius demonstrate that β, the cointegrating vectors, can be estimated as the eigenvectors associated with the r largest, statistically significant eigenvalues found by solving the problem

$$\left|\lambda S_{kk} - S_{k0} S_{00}^{-1} S_{0k}\right| = 0 \tag{7}$$

where S_{00} represents the residual moment matrix from a regression of ΔX_t on $\Delta X_{t-1}, \ldots, \Delta X_{t-k+1}$, S_{kk} is the residual moment matrix from a regression of X_{t-k} on ΔX_{t-k+1}, and S_{0k} is the cross-moment matrix. These eigenvalues readily permit the formation of likelihood ratio tests to determine the value of r. Johansen and Juselius propose two tests with differing

5. See Johansen and Juselius (1990) for further details.

assumptions about the alternative hypothesis: (i) the Trace statistic tests the restriction $r \leq q$ $(q < n)$ against the completely unrestricted model $r \leq n$; and (ii) the λ^{max} statistic makes the alternative more precise by specifying that only one additional cointegrating vector exists $(r \leq q+1)$. The log-likelihood ratio test statistics are formed as

$$\text{Trace} = -T \sum_{i=q+1}^{n} ln(1-\hat{\lambda}_i) \tag{8}$$

$$\lambda^{max} = -Tln(1-\hat{\lambda}_{q+1}) \tag{9}$$

The critical values are non-standard and are found in Johansen and Juselius (1990). The correct critical values differ depending on whether one imposes the restriction that any cointegrating vectors must also annihilate drift terms in each variable. Since we were unable to find any significant evidence of drift in our variables, there should be no drift in our variables that is not annihilated by the cointegration vectors. Therefore we perform the test under the annihilation restriction.[6] The JJ test results, reported in Table 2 (p. 19), show significant evidence of cointegration regardless of which test statistic we use. We therefore conclude that the real effective exchange rate is cointegrated with real oil prices.

In sum the results suggest that the U.S. real effective exchange rate and the U.S. price of oil are both integrated of order one and are jointly cointegrated. This has a number of interesting implications. First, finding significant evidence of a unit root in the real effective exchange rate (rather than just failing to reject the null hypothesis of a unit root) is itself interesting evidence against the hypothesis of long-run purchasing power parity, which would require that real exchange rates be stationary. Second, the fact that the price of oil and the exchange rate have trended together over time makes it interesting to try to capture this feature of the data in RBC models. Third, our

6. Not imposing the annihilation restriction leads to the same conclusions. Furthermore, a likelihood ratio test of this restriction under the maintained hypothesis of one cointegrating vector cannot reject the restriction even at the 10 per cent level.

results also bear directly on the failure to find cointegration between real exchange rates and international real interest rate differentials. While Baxter (1993) and Meese and Rogoff (1988) suggest that their inability to find evidence of cointegration between the two variables is due to a missing real variable, the evidence presented above suggests otherwise. We find evidence of cointegration between the price of oil and the real exchange rate, which suggests that real interest rate differentials are not part of any long-run relationship with the real exchange rate, and that at most they explain only short-run exchange rate movements. We explore this possibility further in Section 4. For the time being we simply note the possibility that real interest rate differentials may simply be I(0) instead of I(1). While standard tests typically find no significant evidence of stationarity in long-run real differentials, we know of no significant evidence of unit roots in them either, which suggests that the data may simply be uninformative. This would certainly be more consistent with models of long-term international capital flows, which lead us to expect a convergence in real interest rates across nations over time.

In sum the results suggest that the U.S. real effective exchange rate and the U.S. price of oil are both integrated of order one and are jointly cointegrated. This implies that oil prices can adequately capture all the permanent innovations in the U.S. real exchange rate over our sample period and that there is a long-run relationship between these two variables.

3 Causality results

In this section we investigate the issue of causality. From Engle and Granger (1987), we know that cointegration implies that at least one of our two variables must Granger-cause the other (bidirectional causality is also a possibility). Understanding the apparent causal relationship in the data is interesting both for econometric and economic reasons.

On the econometric side, Granger-causality has important implications for inference and for evaluating the accuracy of conditional forecasts. Johansen (1992) shows that one of the conditions necessary to perform inference in a single-equation framework is weak exogeneity of the cointegrating variables with respect to the first variable under consideration. Ericsson (1992) notes that valid predictions from a conditional model require strong exogeneity (that is, both weak exogeneity and Granger-causality). If this were not the case, then it would be possible for a misspecified model to give better conditional forecasts than a correctly specified model. For example, it has become common practice to compare the forecast performance of structural exchange rate models to that of a random walk. However, if the model's forecasts use information on future values of variables that may be Granger-caused by the exchange rate (such as relative price levels, output, trade balance and so forth), then even misspecified models should be able to beat a random walk.

Causality might also shed light on the economic mechanism creating the long-run link between oil prices and the U.S. exchange rate. For example, it could be argued that the U.S. dollar exchange rate causes fluctuations in U.S. dollar oil prices, not vice versa.[7] Finding that Granger-causality runs from oil prices to the exchange rate and not the reverse would be evidence against such a mechanism. However, even if oil prices are not affected by exchange rate movements, we might still expect to find that exchange rates Granger-cause oil prices. For example, if forward-looking agents treat exchange rates as asset prices, then they should reflect all publicly available information, including future expected changes in oil prices. However, if oil prices have predictive power for subsequent exchange rate changes, this raises the question of whether exchange rates properly incorporate all available public information.[8]

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7. For a simple theoretical model, see Boughton et al. (1986, Appendix I.) The logic here is that given a fixed world supply, a rise in the real U.S. dollar exchange rate implies higher real domestic oil prices for other nations for a given U.S. dollar real oil price. If the market had previously been in equilibrium, then *ceteris paribus* the real appreciation of the U.S. dollar would bring about an excess supply of oil, since the demand in the non-U.S. market presumably falls in response to higher real domestic prices. To restore equilibrium requires lower real oil prices. The net effect of a stronger U.S. dollar is therefore lower real oil prices in U.S. dollars, but higher real oil prices in foreign currencies, and a shift in oil consumption from foreign countries to the United States. Note that this implies that higher U.S. dollar oil prices should be associated with a depreciation of the U.S. dollar, whereas the results of the error-correction model in the next section indicate the opposite.

8. Note that forecastability of real exchange rate changes using public information does not *necessarily* imply a violation of semi-strong market efficiency, even in the absence of risk premiums, since real interest rate differentials may offset the predicted exchange rate movements. However, substantial research on this question has shown that interest rate differentials typically have little or no explanatory power for exchange rate changes. See Hodrick (1987).

Our first step in testing for causality is to test for "long-run causality," or more accurately, to determine whether the real price of oil is weakly exogeneous in the sense of Engle, Hendry and Richard (1983). This can be tested using the likelihood-ratio test described in Johansen and Juselius (1990). This simply tests whether the long-run relationships captured by ΠX_{t-k} in (5) enter significantly in the equations for changes in oil price or exchange rate changes. The evidence of cointegration presented in the previous section implies that ΠX_{t-k} must enter significantly in at least one of these equations. The results, shown in Table 3 (p. 19), imply that the price of oil is weakly exogenous, while the real exchange rate is not. In other words, in the long run the level of the real effective exchange rate adjusts to the price of oil and not vice versa.

Next, we test for causality in the more general sense of standard Granger-causality, using standard tests on the vector autoregression level representation of our system. As demonstrated in Sims, Stock and Watson (1990), standard inference procedures are valid in this case under the maintained hypothesis of one cointegrating vector provided that we test the exclusion restrictions on one variable at a time. Since all results are based on asymptotic approximations, we use the limiting χ^2 critical values instead of their more common F-distributed counterparts.

The results from our tests are reported in Table 4 (p. 20). They indicate that while the price of oil Granger-causes the real exchange rate, there is no evidence to support the converse. This is consistent with the weak exogeneity results mentioned above. Therefore, the causality results suggest that, even though the price of oil is typically quoted in U.S. dollars, movements in the external value of the U.S. dollar have no significant effect on the price of oil. This conclusion is consistent with the results of Hamilton (1983), who uses Granger-causality tests with a wide range of U.S. macroeconomic variables and finds support for the proposition that oil price shocks are exogenous to the United States. Similar conclusions have been reached by Burbidge and Harrison (1984), Gisser and Goodwin (1985) and Mork (1989).

Although this result may be counter-intuitive to some researchers, a review of the behaviour of real oil prices (shown in Figure 1, p. 23) over the most recent floating exchange rate period shows that the series is dominated by major persistent shocks around 1973-74, 1979-80 and

1985-86, with another large but transitory shock in 1990-91. The historical record offers us a very plausible explanation for these shocks: they are supply-side shocks resulting from political conflicts in the Middle East.[9] Note that we are not arguing that oil prices (or the stability of price cartels) are immune to the laws of supply and demand or that they cannot be affected by shifts in the growth rates of the industrialized world. Instead, we feel that there is ample reason to believe that such demand-side factors have been small relative to the supply shocks experienced over the last twenty years, and that the supply shocks have been exogenous in the sense of most macroeconomic models.

4 An error-correction model

In this section we determine how well the dynamic process generating the U.S. real exchange rate can be captured by a single-equation error-correction model (ECM). According to the Engle and Granger (1987) Representation Theorem, the presence of cointegration in a system of variables implies that a valid error-correction representation exists. This theorem together with the evidence of weak exogeneity found above suggests that we can use a single-equation error-correction representation without the loss of either efficiency or the ability to perform proper inference.[10] In general, we can write the single-equation ECM as

$$\Delta Y_t = \alpha (Y_{t-1} - X_{t-1}\beta) + \sum_{i=0}^{k} Z_{t-i}\gamma_i + \varepsilon_t \tag{10}$$

where $Y-X\beta$ represents the error-correction mechanism and α gives the speed of adjustment towards the system's long-run equilibrium; Z represents a matrix of stationary variables, possibly

9. Specifically, the 1973-74 episode corresponds to the Organization of the Petroleum Exporting Countries (OPEC) oil embargo following an Arab-Israeli war; the 1979-80 episode was due to supply changes associated with the Iranian revolution; the 1985-86 change was due to the effect of the Iran-Iraq war on OPEC solidarity; and the temporary 1990-91 shock was the result of the Gulf War. To attribute any of these Middle Eastern conflicts in turn to macroeconomic developments in the G-7 would be incredible.

10. According to Johansen (1992), if the price of oil is weakly exogenous with respect to the real exchange rate, then estimation and inference on a single equation will be asymptotically equivalent to that on the full system. See Phillips and Loretan (1991) for a discussion of inference in an ECM under the assumption of weak exogeneity.

including contemporaneous and lagged differences of X and Y, that attempts to capture the short-run dynamics of the dependent variable; β and γ are vectors of parameters; and ε is an error term.

In our model Y is equal to the real exchange rate (RE) and X represents the price of oil (POIL). We define Z to include the lagged dependent variable, an international interest rate differential measure and the first differences of the real price of oil.[11] In preliminary work we tried three different measures of an interest rate differential, namely, (i) the United States versus Japan; (ii) the United States versus Germany; and (iii) a weighted measure constructed to reflect the make-up of the effective U.S. real exchange rate.[12]

We set k equal to 12 and estimate equation (10) over the 1972M2 to 1985M12 sample period using non-linear least squares. The estimation sample is truncated early to allow us to perform out-of-sample forecasting exercises with the remainder of the data. We reduce the dynamics of the ECM by successively omitting variables with the lowest t-statistics and re-estimating. The resulting ECM (reported in Table 5, p. 20) is simply

$$\Delta RE_t = \alpha (RE_{t-1} - \beta_0 - \beta_1 POIL_{t-1}) \tag{11}$$

Despite the remarkably simple specification, the results indicate that the price of oil can account for 6.7 per cent of the month-to-month variation and most of the longer horizon systematic movements in the real exchange rate. Specifically, the residual diagnostic tests yield little evidence of autocorrelation and autoregressive conditional heteroscedasticity (see residual diagnostic tests in Table 5).[13] The test for kurtosis, however, forces us to reject the null hypothesis of no kurtosis. We note that such a result is common in financial time series, and that given our large number of

11. Augmented Dickey-Fuller (1979) and Phillips-Perron (1988) tests find the dependent variable, the interest rate differentials and the first difference of the price of oil to be stationary. In order to capture better the stance of monetary policy, the interest rate measure for each nation is defined as the difference between their short and long rates. Results did not seem to be sensitive to this definition.

12. Precise interest rate differential definitions are available from the authors upon request.

13. Since there is evidence of non-spherical residuals at the 10 per cent level, we re-estimated our ECM with Newey and West (1987) standard errors. This did not change the significance of the parameter estimates.

observations and the absence of serial correlation, this is unlikely to affect our inference. As for the parameter estimates, the speed of adjustment term (α) is -0.028, indicating that about 28.6 per cent of adjustment toward the long-run equilibrium is completed within one year. This gives a half-life of 2.1 years. A 1 per cent increase in the price of oil will lead to a 0.513 per cent *appreciation* of the dollar in the long run. We defer discussion of the interpretation of this relationship to the latter part of this section. However, it is noteworthy that the long-run parameter estimate from the single-equation approach is within one standard error of its estimated value of 0.42 from the Johansen and Juselius (1990) system estimation approach.

Figure 2 (p. 23) presents both in-sample and out-of-sample dynamic simulations of the specification starting in 1972M2. The simulations to the right of the vertical line at 1986M1 are post-sample dynamic simulations. It is readily apparent that the ECM tracks the realized values reasonably well both in- and out-of-sample. This again suggests that the long-run value of the real exchange may adequately be specified as simply a function of the price of oil.

Another criteria of a model's adequacy is parameter stability over the sample period. As Hendry (1979) has shown, dynamic misspecification can be critical to the stability of an equation, so parameter instability may suggest that the specified dynamic structure is inadequate. Therefore, we use a sequential Chow test to determine the stability of our ECM over the 1972M2 to 1993M1 period. Note that this sample period includes data that were not used in the specification of our final ECM. The results from the sequential Chow test (Figure 3, p. 24) suggest that the specification is stable over the sample period.[14]

The estimate of β reported in Table 5 (p. 20) implies that, despite the fact that the United States is a major importer of oil, higher oil prices make the dollar appreciate. This could be due to any one of several possible mechanisms. The most commonly mentioned explanation is that the United States imports relatively less oil than its trading partners in Europe and eastern Asia, so that

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14. Andrews (1993) points out that if a break point is determined by a search over the sample, then the standard critical values are smaller than the appropriate values and inference is biased against the null of stability. To control for this effect we use critical values calculated by Andrews (1993).

the dollar depreciates less than their currencies in response to higher oil prices, which implies an appreciation of the dollar in effective terms. Alternatively, if the net import demand for oil is elastic in the long run then oil price increases will tend to improve the trade balance and require an appreciation to restore external equilibrium.[15] Also, Krugman (1983a, 1983b) and Golub (1983) note that in a three-country world (including Europe, America and the OPEC countries), higher oil prices will transfer wealth to the oil exporter (OPEC) from the oil importers (America and Europe). If America is a relatively small share of OPEC's export market but a large share of OPEC's import market, then the transfer of wealth from the industrial countries to OPEC would tend to improve the U.S. trade balance. McGuirk (1983) and Golub (1983) both find that this kind of effect helps to explain the appreciation of the U.S. dollar in response to higher oil prices.[16]

The results of this section show that consideration of the effects of oil price changes can lead to a simple, stable model of exchange rate changes with good apparent explanatory power, particularly at lower frequencies. Taken together with the results on cointegration and causality presented previously, this suggests that oil price effects have the potential to improve the performance of structural exchange rate models, a possibility that we explore further in the next section. We were surprised, however, by our failure to find any significant role for the real interest rate differentials in explaining even transitory exchange rate changes. It is possible that this is due to our particular measures of the real interest rate and that the use of other proxies for unobserved expectations of inflation might alter this conclusion. On the other hand, it is broadly consistent with the limited explanatory power of this variable reported by Campbell and Clarida (1987), Meese and Rogoff (1988) and Edison and Pauls (1993). It is also possible that the limited explanatory power of real interest rates at medium-to-low frequencies that Baxter (1993) reports may be

15. Note that the elasticity of net import demand is greater than that of either domestic supply or demand, and that this elasticity increases as the ratio of net imports to domestic demand falls. For example, if the U.S. imports 20 per cent of demand and a 50 per cent price increase cuts domestic demand only 5 per cent and boosts domestic supply only 5 per cent in the long run, then the balance of trade deficit in energy improves by 17.5 per cent. Strictly speaking, effective exchange rates should appreciate if U.S. net import demand for oil is *relatively* more elastic than that of other industrial nations. Since the United States is relatively less dependent on imports, this condition will be satisfied *ceteris paribus*.

16. Note, however, that Golub argues shifts in market share and changes in portfolio preferences may have changed the expected relationship between oil prices and the U.S. exchange rate around the time of the second OPEC shock. In contrast, we found no significant evidence of instability in the relationship.

captured more effectively by the oil price variable.

5 Can we forecast exchange rate changes?

In Section 3 we established the strong exogeneity conditions required to perform valid out-of-sample forecasting comparisons. Therefore, we now consider the ECM's ability to forecast out-of-sample exchange rate changes. First, we follow the methodology used by Meese and Rogoff (1983) and compare the out-of-sample forecasts produced by the ECM to those generated by a random walk, using Theil's U statistic. Specifically, we begin by estimating the specifications on data up to 1985M12 and then generating forecasts for all quarterly horizons up to 24 months, using realized values of POIL. Another month of data is then added, the equation is re-estimated and new forecasts are generated for all horizons. This process is repeated until the end of our data set (1993M1). Finally, forecast errors are then calculated for all estimation periods and all horizons. Table 6 (p. 21) reports Theil's U statistic, which is the ratio of the model forecast's root-mean-squared-error (RMSE) to the RMSE of a random walk forecast (that is, no change). Values less than 1 therefore imply that the model performs better than a random walk, while values greater than 1 indicate the reverse. The U statistics suggest that the ECM performs better than the random walk over the forecasting sample, with the model's performance improving slowly as the forecasting horizon increases.

While the results for the ECM are suggestive, it would be useful to determine whether the model forecasts *significantly* better than a random walk. To address this issue we now turn to the concept of forecast encompassing, which allows us to test formally the forecast performance of different models.[17] A model is said to encompass a competing formulation if its forecasts help to predict the forecast errors of its competitor, while its competitor's forecasts give no information about the first model's forecast errors. Tests of forecast encompassing are easily calculated, based on an artificial regression involving only forecasts and their errors:

17. For general expositions on forecast encompassing, see Chong and Hendry (1986).

$$u_t^A = \alpha + \beta\,(F_t^A - F_t^B) + \varepsilon_t \tag{12}$$

where u_t^A is the forecast error from model A, F_t^B and F_t^A are the forecasts of model B and A respectively, and α and β are parameters. If β is statistically significant, this implies that model B forecast-encompasses model A, which suggests that model B provides information not admitted by model A.[18] In Table 7 (p. 21) we report the marginal significance level associated with the t-statistic on β from ordinary least squares estimation of (12). Since fully efficient multistep forecasts are linear combinations of the one-step-ahead forecasts, only the one-step-ahead forecast need be considered here. The results imply that the ECM forecast-encompasses the random walk, which suggests that our simple specification provides us with additional information relative to a random walk.

One potential criticism of these forecast-encompassing tests is that they rely on inference in small samples where non-normality is suspected, yet inferences are based on the assumption of normality. To avoid this problem, we consider a non-parametric test recently proposed by Pesaran and Timmermann (1992).[19] The test examines the predictive accuracy of a set of forecasts to predict the direction of the change in a variable. Since the test is non-parametric, it does not require assumptions about the distribution or the parameters of the sampled population. Pesaran and Timmermann (PT) derive the test statistic and find it to be distributed normally with zero mean and unit variance. A rejection of the null hypothesis suggests that the model generating the forecasts has a statistically significant ability to predict changes in a variable. We dynamically simulate the ECM starting in 1973M3 and calculate the PT test statistic. We calculate the test statistic to be equal to 2.14, which allows us to reject the null at the 3 per cent level and conclude that the ECM is able to significantly predict changes in the U.S. real exchange rate.

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18. Since we are comparing our model (A) to a random walk (B), equation (12) simply suggests regressing the ECM's forecast errors on its predicted exchange rate changes. If β were significantly different from zero, this would imply that the ECM's forecasts are systematically over- or under-predicting the amount of exchange rate movement. When we reverse the test to see if the ECM encompasses the random walk, we simply change the dependent variable from the ECM's forecast errors to actual exchange rate changes.

19. We also used Spearman's rank correlation coefficient as a non-parametric test of whether $\beta=0$ in (12). This gave the same conclusions as the parametric tests reported in Table 7.

6 Conclusions

We have explored whether a link exists between the price of oil and the U.S. real exchange rate. The results presented above show that the U.S. real exchange rate appears to be cointegrated with the real price of oil. This suggests that oil prices may have been the dominant source of persistent real shocks over the post-Bretton Woods period. Causality tests also indicate that causality runs only from oil prices to exchange rates and not vice versa. The single-equation ECM relating these two variables is stable and captures much of the in- and out-of-sample movement in the exchange rate in dynamic simulations. Tests show that the ECM has significant out-of-sample predictive ability for both the size and sign of changes in the real effective exchange rate.

Our results make several useful contributions. They show that real oil prices can account for innovations in another important U.S. macroeconomic variable and thereby add to the literature that documents the influence of oil price shocks on the U.S. economy. They also provide support for McCallum's (1989) conjecture that oil price shocks should be incorporated into models of real business cycles and present another interesting stylized fact that models of international business cycles will need to capture. In addition, they advance the research on the failure of real interest rate parity by identifying a real factor that can account for the nonstationarity in real exchange rates.

The results also suggest a potentially important role for energy prices in future research on exchange rate modelling. For instance, Amano and van Norden (1993a, 1993b) find energy price shocks also to be an important long-run determinant of the Canada-U.S. real exchange rate and they find that their ECM forecasts better than a random walk out-of-sample. These results stand in contrast to previous work using monetary and portfolio-balance models of exchange rate determination, which generally fail to produce significant evidence of cointegration or specifications that out-perform a random walk in post-sample forecasting.

Tables

Table 1: Tests for unit roots
Augmented Dickey-Fuller (ADF), Phillips-Perron (PP) and Kwiatkowski, Phillips, Schmidt and Shin (KPSS) Tests

Sample: 1972M2 to 1993M1

Variable	ADF lags	ADF t-statistic	PP t-statistic	KPSS
U.S. dollar	1	-1.667	-1.673	0.550*
Price of oil	2	-2.103	-2.503	0.686*

Note: * indicates significance at the 5 per cent level.

Table 2: Johansen-Juselius test for cointegration estimation under assumption of restricted drift

Number of lags = 4[a]

Number of cointegration vectors	Trace statistic	λ Maximum statistic
1	23.511*	17.448*
0	6.063	6.063

a. Appropriate lag lengths were determined through standard likelihood ratio tests with a finite-sample correction. We began with 13 lags and tested down.

Note: * indicates significance at the 5 per cent level. The critical values are taken from Johansen and Juselius (1990).

Table 3: Johansen weak exogeneity tests

Variable	Test statistic	Significance level
U.S. dollar	10.680	0.3E-03
Price of oil	0.799	0.371

Table 4: Granger-Causality tests

Dependent variable	Number of lags[a]	Independent variable	Significance level
U.S. dollar	4	Price of oil	0.001
Price of oil	4	U.S. dollar	0.906

a. Lag lengths were selected on the same basis as in footnote a, Table 2. However, the conclusions are robust for lag lengths 1 to 13.

Table 5: Error-correction model results

$$\Delta RE_t = \alpha(RE_{t-1} - \beta_0 - \beta_1 POIL_{t-1})$$

Sample: 1972M2 to 1985M12

	Parameter estimate	Standard error	t-statistic
Adjustment	-0.028	0.013	-2.087
Constant	3.033	0.707	4.290
Price of oil	0.513	0.230	2.233
Residual diagnostic tests[a]			
LM(1)[b]	0.073	ARCH(1)	0.056
LM(4)	0.707	ARCH(4)	0.101
LM(12)	0.179	ARCH(12)	0.804
Kurtosis	0.12E-4	Skewness	0.328

a. Reported values are the marginal significance levels.
b. Lagrange multiplier tests for autocorrelation of order (k).

Table 6: Theil's U statistic for forecast comparisons: model-based versus random walk

Starting period: 1986M1

Horizon	1	2	4	12	24
	0.973	0.957	0.927	0.806	0.694

Table 7: Forecast-encompassing tests

Starting period: 1986M1

Hypothesis	$\alpha = 0$	$\beta = 0$	$\alpha = \beta = 0$
Model[a]	0.903	0.044	0.012
Random walk[b]	0.902	0.884	0.943

a. Testing the null hypothesis that the ECM does not forecast-encompass the random walk.
b. Testing the null hypothesis that the random walk does not forecast-encompass the ECM.

Note: Reported values are the marginal significance levels.

Figure 1: The real U.S. dollar price of oil
based on 1983 U.S. dollars

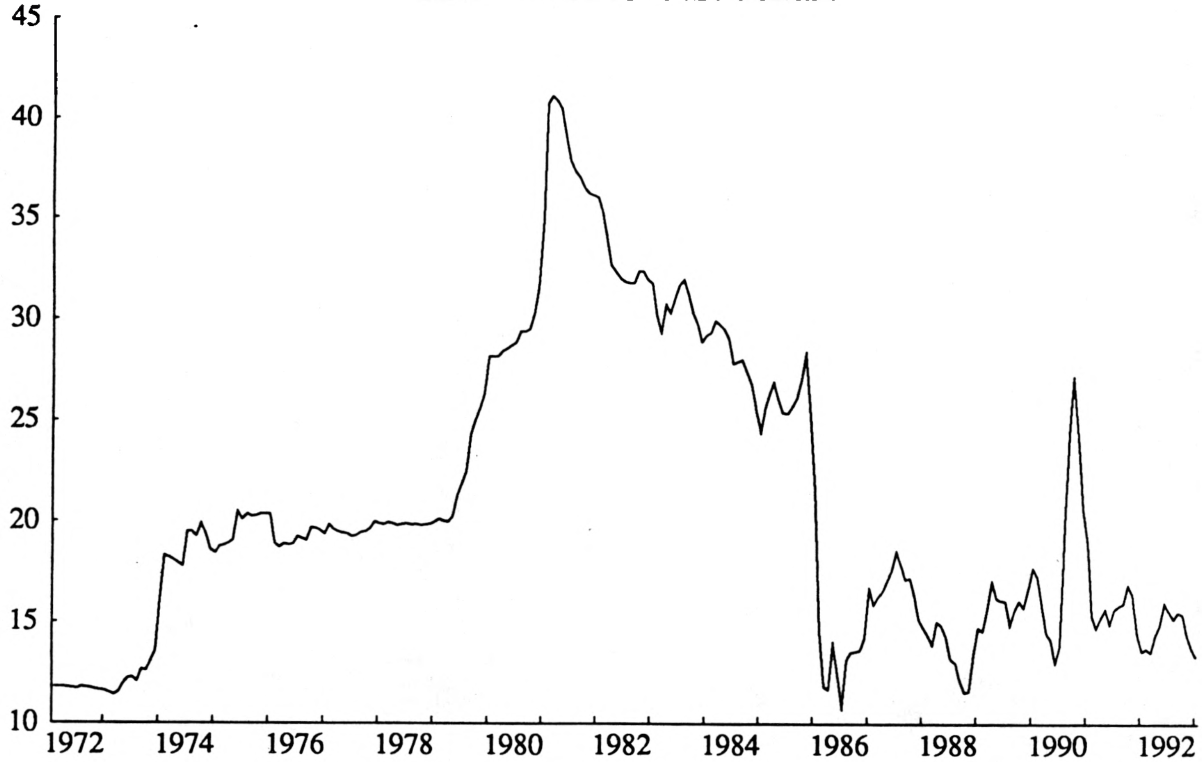


Figure 2: Dynamic simulation of the ECM

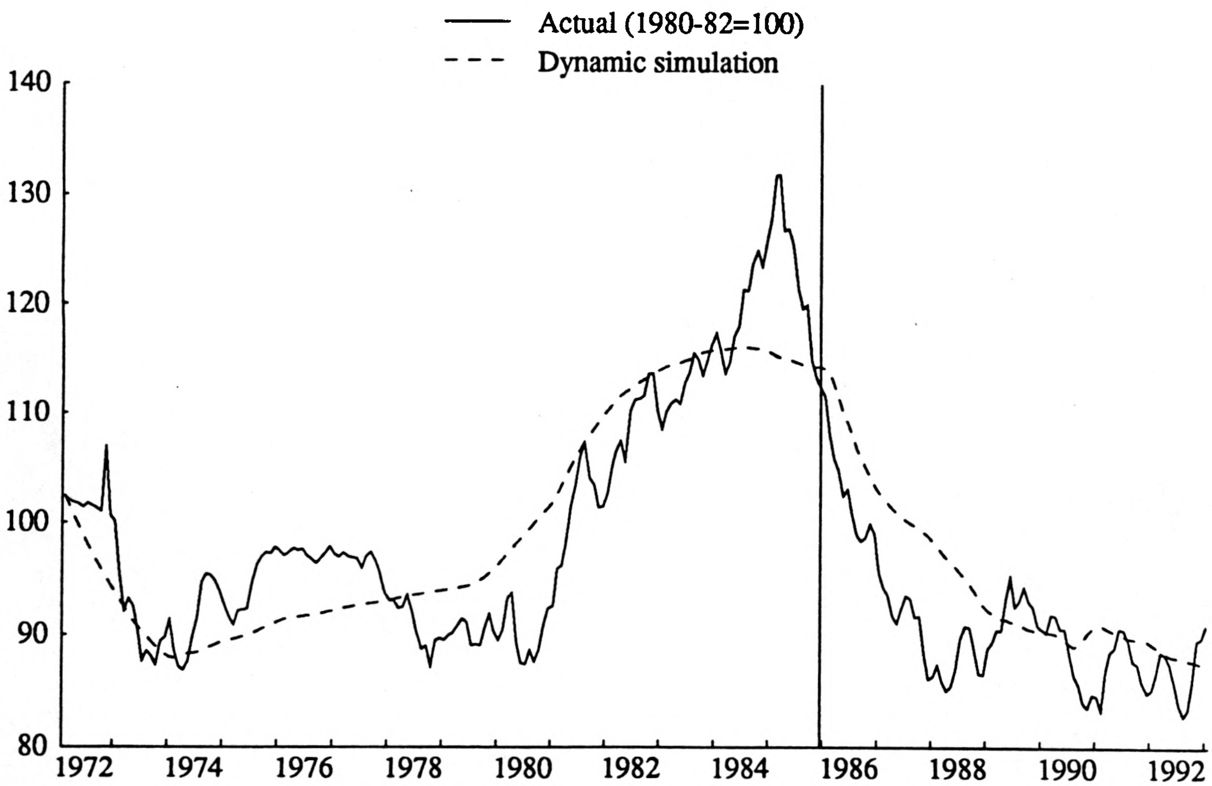
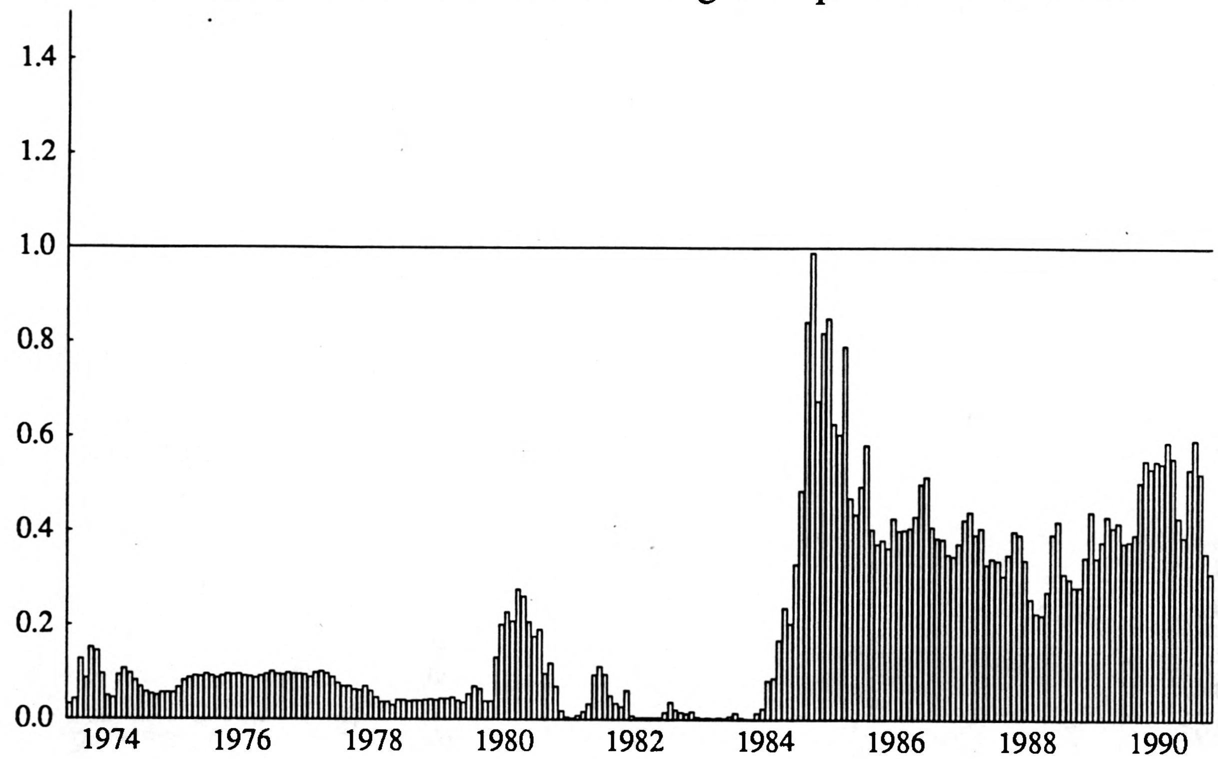


Figure 3: Sequential Chow tests
Test statistics normalized to 1.0 using the 5 per cent critical value



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