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Terms of Trade and Real Exchange Rates: The Canadian Evidence

by
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Bank of Canada Banque du Canada

Terms of Trade and Real Exchange Rates: The Canadian Evidence

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CONTENTS

Abstract

This paper presents empirical evidence linking the Canada-U.S. real exchange rate with terms-of-trade variables. Using recently developed econometric methods for nonstationary data, we find that the real exchange rate is cointegrated with terms-of-trade variables, and that causality runs from the terms of trade to the exchange rate. We then construct a simple exchange rate equation that satisfies several specification tests and performs better than a random walk in post-sample forecasting experiments. The results suggest that much of the variation in the real exchange rate is attributable to movements in the terms of trade and that the influence of monetary factors is secondary.

Résumé

La présente étude fait état de résultats empiriques liant l'évolution du taux de change réel du dollar canadien par rapport au dollar américain à celle des termes de l'échange du Canada. En appliquant des méthodes économétriques récentes propres à l'analyse de données non stationnaires, les auteurs sont portés à conclure que le taux de change réel est cointégré avec les termes de l'échange et que la causalité va de ces derniers au premier. Puis ils mettent au point une équation simple de taux de change qui résiste à différents tests de spécification et prédit mieux qu'un modèle de marche aléatoire au delà de la période d'estimation. D'après les résultats, une grande partie des variations du taux de change réel est attribuable aux mouvements des termes de l'échange, et l'incidence des facteurs d'ordre monétaire est secondaire.

1 Introduction

Recent surveys of exchange rate models, such as Meese (1990), Mussa (1990) and MacDonald and Taylor (1992), tend to agree on only one point: that existing exchange rate models are unsatisfactory.

The failure of these models has come to be recognized in a number of ways. First, monetary models that appeared to fit the data for the 1970s were rejected when the sample period was extended to the 1980s (see Frankel 1984 and Backus 1984). Later work on the monetary approach, such as that of Campbell and Clarida (1987) or Meese and Rogoff (1988), found that even quite general predictions about the comovements of real exchange rates and long-term real interest rates did not seem to conform with the evidence. One interpretation they suggest is that shifts in factors determining the long-term real exchange rate play a significant role in exchange rate movements. Further evidence to support this view is offered by Lastrapes (1992), who finds that much of the variance of nominal exchange rate changes, even at monthly horizons, seems to be due to permanent real shocks.

A second important factor in the rejection of existing exchange rate models was the shift in focus from estimated coefficient values to broader criteria that included out-of-sample performance. The seminal study in this vein is Meese and Rogoff (1983), which showed that several monetary and portfolio-balance models which appeared to fit well in-sample could not explain out-of-sample exchange rate movements better than a random walk, particularly at short horizons. It also helped lead to the characterization of exchange rates as a random walk (for example, see Lothian 1987). Given a growing appreciation of the complications that nonstationary processes (such as random walks) can create for model estimation and hypothesis testing, this suggested that applied researchers should use the newer unit-root and cointegration methods. These new techniques have tended to confirm the existing results, however, by showing that monetary and portfolio-balance models generally do not provide a cointegrating relationship for the exchange rate (for example, see Adams and Chadha 1991).

Both of these factors suggest that existing exchange rate models omit some important source of shocks that alter long-run real exchange rates. This paper explores one such source: exogenous shifts in the terms of trade. Specifically, we consider whether terms-of-trade shocks can explain much of the historical variation in the Canada-U.S. real bilateral exchange rate. The Canada-U.S. experience is particularly interesting for several reasons. First, since these countries share the largest bilateral volume of trade in the world, one might reasonably expect to find terms-of-trade effects on the exchange rate. Second, since both nations have long-standing policies of flexible exchange rates and open capital markets, one might also reasonably expect monetary and portfolio-balance effects to be important.[1] Third, Canada is small relative to world markets, so its terms of trades might be expected to behave exogenously, clarifying the link between the exchange rate and the terms of trade. Fourth, the dominance of the United States in Canada's trade relations means that Canada's effective exchange rate is well approximated by the bilateral exchange rate, which simplifies our analysis.

While models of the long-run effects of terms-of-trade shocks on the exchange rate are well established, there is little empirical work on this question. Our contribution is that we use up-to-date data on exchange rates and terms of trade, and we analyse them using a variety of recently developed econometric methods for nonstationary data. We find both a cointegrating relationship between the real exchange rate and terms-of-trade variables, and a very simple exchange rate equation that forecasts well out of sample. We also confirm that causality runs from the terms of trade to the exchange rate, and not vice versa. Our equation shows that much of the variation in the Canada-U.S. real exchange rate since 1973 can potentially be attributed to terms-of-trade shocks, and that monetary factors seem to play only a secondary role.

In Section 2, we review the theory linking terms-of-trade movements to equilibrium exchange rates in a small open economy. Section 3 introduces the set of variables we investigate. Section 4 establishes the presence of a unit-root in the real exchange rate and shows the existence of a cointegrating relationship between the terms-of-trade variables and the real exchange rate. Section 5 then goes on to show how the cointegrated system may be reduced to a simple, single-equation error-correction model. Various diagnostic tests are used to assess the model's adequacy, and we investigate some alternative interpretations of the model. Section 6 examines the model's ability to forecast exchange rate changes out of sample and explores its implications for market efficiency. Section 7 considers the relative importance of monetary and terms-of-trade shocks in explaining movements in the real exchange rate. The final section summarizes the results and offers conclusions.

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1. Note that the Bank of Canada moved to flexible exchange rates in 1970, well before the generalized breakdown of the Bretton Woods system began in 1973.

2 A model of the terms of trade and the equilibrium exchange rate

In the tradition of small open-economy models, the terms of trade are exogenous variables that play a key role in determining not only the exchange rate, but the whole distribution of resources and activity through the economy.[2] However, it is important to note that economic theory cannot unambiguously specify the effect of a terms-of-trade movement on the real exchange rate. The literature on the effects of terms-of-trade shocks on exchange rates is too vast to review here, but a particularly useful and general analysis is that of Neary (1988).

Neary considers the case of a small open economy producing arbitrary numbers of traded and non-traded goods under competitive conditions and examines the real exchange rate that is required to bring about a balance of payments equilibrium. His results are neatly summarized in the equation[3]

$$\hat{\pi} = \frac{\omega'}{\bar{S}}(\mu - \sigma)\, d\Phi \tag{1}$$

where $\hat{\pi}$ is the proportional change in the real equilibrium exchange rate (defined such that an appreciation implies $\hat{\pi} > 0$), ω is the vector of individual commodity weights in the non-traded goods price index, $\bar{S}$ is a re-normalization of the matrix of the responses of the demand for non-traded goods to a change in their prices, μ is a vector of the marginal propensities to consume non-traded goods in response to a change in utility, σ is the vector of cross-price effects of a change in the terms of trade on the excess supply of non-traded goods, and $d\Phi$ is the change in utility arising from a change in the terms of trade.

An improvement in the terms of trade normally implies $d\Phi > 0$. Since $1/\bar{S}$ is a positive definite matrix, $\hat{\pi}/d\Phi$ will be positive (that is, imply a real appreciation), provided the elements of ω and $\mu - \sigma$ are positively correlated. In other words, a terms-of-trade improvement leads to an appreciation, so long as it increases utility and the weighted average of its effect on the excess demand for non-traded goods $(\mu - \sigma)$ is positive. In that case, the real appreciation is required to eliminate the excess demand for non-traded goods created by the shift in the terms of trade.

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2. The extension of real-business-cycle models to the open economy has stimulated interest in terms-of-trade shocks in recent years. See, for example, Macklem (1992), Ostry (1988), Turnovsky (1991) or Mendoza (1992). However, this literature is much broader in scope and focusses on the behaviour of calibrated rather than directly estimated models.

3. Neary (1988, eq. 11).

However, $(\mu - \sigma)$ may have many negative entries, which means that increases in utility can create an excess supply of non-traded goods, requiring a depreciation to clear the market. Hence, the effect of a terms-of-trade movement on the real exchange rate depends critically on which response dominates. Furthermore, $(\mu - \sigma)$ may itself vary, depending on the nature of the terms-of-trade shock. For example, increases in the price of two exported goods have equal effects on the terms of trade but very different effects on the cross-elasticities of demand and supply. It may therefore be important to decompose terms-of-trade movements in some way to find a stable relationship.

While this comparative-statics analysis shows the expected long-run response of the exchange rate, it gives no insight into the short-run dynamics of the exchange rate. However, the most general monetary exchange rate models, associated with Shafer and Loopesko (1983), Campbell and Clarida (1987) and Meese and Rogoff (1988), show that the real interest rate differential should reflect the deviation of the real exchange rate from its expected long-run level.[4] More generally, these monetary models predict a relationship between the deviation of the real exchange rate from its expected long-run value and relative monetary policy. However, they also predict that monetary variables should have no permanent effects on the real exchange rate. Therefore, while we should include monetary variables in a real exchange rate equation, we should expect their effects to be transitory. As we see below, this fits naturally into the cointegration methodology we adopt. The fact that monetary variables must have transitory effects also does not limit their importance in explaining exchange rate behaviour over any fixed horizon.

3 Data description

To determine the potential role of terms-of-trade effects, we examine the post-Bretton Woods data for a relationship between Canada's real exchange rate and the terms of trade. For simplicity, our measure of the real price of foreign exchange (RPFX) is Canada's bilateral real exchange rate with the U.S. (C$/U.S.$), deflated using consumer price indices (CPIs). Strictly

4. These authors note that if uncovered interest parity holds, then

$$q - E_0(q_{+1}) = r^* - r$$

where q is the real price of foreign exchange, $E_0(X_{+1})$ is the expectation of X next period conditional on information available now, r is the domestic real interest rate and r^* is the foreign real interest rate.

speaking, terms-of-trade shocks should be expected to influence a nation's effective exchange rate rather than any particular bilateral rate. However, given the predominant weight of the U.S. component in any effective index for Canada, we focus on the bilateral rate to be more consistent with the monetary model and with the literature on exchange rate forecasting.[5]

For the terms-of-trade index, we use the price of exported commodities divided by the price of imported manufactured goods. This is consistent with the conceptual view of Canada as a net exporter of resource-based commodities and a net importer of manufactures. Some previous studies have tried to find an empirical link between Canada's real exchange rate and the overall terms of trade, without much success (for example, see Lafrance and Longworth 1987). Early in our investigation, however, we found signs that energy prices seemed to play a role that was qualitatively different from that of other commodities. Accordingly, the overall terms-of-trade index was split into two components: the price of exported energy and the price of exported non-energy commodities, each divided by the price of imported manufactured goods. These series are called TOTENRGY and TOTCOMOD respectively.[6]

As noted above, the small open economy model assumes that the country under study is a price-taker on world markets, implying that terms-of-trade movements are exogenous shocks and that any relationship between exchange rate and terms-of-trade movements can be interpreted as a causal link. We note that these measures of the terms of trade are ones historically used within the Bank of Canada to capture price movements in sectors where Canadian firms are thought to behave as price-takers. By limiting ourselves to commodity and energy export prices, we focus on markets where homogeneity should reduce the market power of any exporter.[7] Below, we test for signs of any reverse causality from the exchange rate to our terms-of-trade variables. Separate tests for both long-run effects (weak exogeneity) and short-run effects (Granger causality) support the exogeneity assumption.

To capture the influence of monetary policy, we include a measure of the Canada-U.S. interest rate differential, RDIFF:

5. Nonetheless, we checked the distinction between bilateral and effective real exchange rates by including the U.S. real effective exchange rate against overseas countries as an additional explanatory variable. It failed to enter significantly, suggesting this distinction plays only a minor role.
6. Aside from the interest rate measure, all variables in the model are in natural logarithms. Further detail on definitions and data sources may be found in Appendix 1.
7. In addition, empirical studies suggest that U.S. exporters tend to set their manufacturing export prices in U.S. dollars, ignoring the effects of exchange rate movements on their prices in foreign currency. See Knetter (1989) and Gagnon and Knetter (1990).

$$RDIFF \equiv (i^{short}_{Canada} - i^{long}_{Canada}) - (i^{short}_{US} - i^{long}_{US}) \tag{2}$$

As discussed below, this particular formulation is not critical to our results. However, it has the advantage of having two alternative interpretations. First, the gap between short and long nominal interest rates may be thought of as a proxy for the stance of monetary policy. By capturing the relative tightness of monetary policy, while avoiding the problems caused by unstable money demand functions when money supplies are used directly, the inclusion of RDIFF should help capture the usual short-run dynamics implied by standard monetary models of the exchange rate. The forward-looking nature of the long-term interest rate also allows us to capture expectations of future monetary policy. Second, RDIFF is essentially the same variable Frankel (1979) uses to proxy the real interest rate differential, which is the key variable suggested by a broad class of monetary models.[8] Regardless of which interpretation is preferred, both suggest RDIFF should capture short-run exchange rate dynamics. They are also both consistent with our finding, below, that RDIFF has no long-term effect on the real exchange rate. To keep our terminology consistent, however, we will refer to RDIFF as the real interest rate differential, without intending to prejudice its interpretation.

Graphs of these data series are shown in Figure 1 (page 32). As with any empirical equation, the choice of how to measure a given variable may be open to debate. As an aid to interpreting our results, we therefore note at the outset that our results are robust to some changes in variable definition, but not to others. Given the volatility of commodity prices, TOTCOMOD and TOTENRGY are dominated by the variations in energy and non-energy commodity prices, with changes in the price of manufactured goods acting as little more than a deflator. Minor changes in their construction do not alter our conclusions. As we will see, however, it is essential to enter commodity and energy terms of trade as separate variables. Using only an aggregate measure of the terms of trade gives very different results. The choice of frequency and deflator for the real exchange rate is not critical.

Our initial work in this area used unit-labour cost indices and GDP deflators on quarterly data and produced similar results. The change to the CPI was made to facilitate the transition to

8. Under the assumption that international real interest rates converge in the long run, the nominal long-term interest rate differential captures expected inflation differentials and the short-long differential captures the real interest rate.

monthly data. Some have suggested that the relationship between energy prices and the real exchange rate is due to the fact our measure of price contains energy prices. However, the same results are found when we deflate using the CPI excluding food and energy.

Finally, the role of RDIFF is not very sensitive to the choice of the nominal interest rate series; using either the long or short nominal interest rate differential gives similar conclusions. However, using a real long interest rate differential with static inflationary expectations, as in Meese and Rogoff (1988), destroys the variable's explanatory power.

4 Unit-root and cointegration tests

An important question to be determined at the outset is whether the real exchange rate contains a unit root. If so, then we should use cointegration methods to model its behaviour, and we have restrictions on how it can be influenced by variables of different orders of integration. However, a unit root in the real exchange rate is usually taken to be evidence against the hypothesis of long-run purchasing power parity (PPP). While tests on most floating exchange rates in the post-Bretton Woods period are unable to find significant evidence of PPP, tests over a century or more of data usually do find such evidence (see Johnson 1990). This seems to be a particularly confusing result, since it is precisely over such long periods of time that we would expect gradual shifts in terms of trade, industrial structure, productivity growth, and so on to alter real equilibrium exchange rates. Therefore, our first step is to test for a unit root in RPFX.

We use a variety of tests to answer this question. First, we use the Augmented Dickey and Fuller (1979) and Phillips and Perron (1988) tests.[9] However, it is well known that such tests may lack power against persistent alternatives if the data do not span a long enough time period (see Perron 1991). Therefore, we also use a test developed by Kwiatkowski, Phillips, Schmidt and Shin (1992) that allows us to test the null hypothesis of stationarity against a unit-root alternative. In this way, we may to be able to distinguish between a series that follows a unit root process and one that is simply uninformative as to the data-generating process. The results shown in Table 1 (page 28) all point to the same conclusion. Both unit-root tests are unable to reject the unit-root null, while the test for stationarity rejects the null hypothesis of stationarity in favour of the unit-

9. Before the application of any of these tests, we pre-tested each of our series for statistically significant evidence of drift. No evidence of drift is found in any of the variables.

root alternative.[10] We therefore conclude that the real exchange rate contains a single unit root, suggesting that long-run PPP should not hold.[11]

An immediate implication of this result is that in the long run, RPFX should only be determined by other integrated variables. Therefore, our next step is to test TOTCOMOD, TOTENRGY and RDIFF for evidence of unit roots. The results are also shown in Table 1 (page 28). They show that while TOTCOMOD and TOTENRGY appear to be nonstationary processes, RDIFF appears to be stationary. Consequently, the real interest rate differential should at most have only a transitory effect on the real exchange rate, while the effects of changes in the terms of trade could be permanent, which is what the theory presented above leads us to expect.

Our next step is to determine whether the nonstationary variables identified above are cointegrated. This has two purposes. First, if we wish to specify an easily interpreted form for an exchange rate equation, such as an error-correction model, we must first determine the existence and number of the cointegrating relationships in the data. Second, this step acts as a check on the adequacy of our model. If the long-run real exchange rate is determined by factors other than those associated with the terms of trade, then their omission should prevent us from finding evidence of cointegration. Evidence of cointegration, on the other hand, implies that TOTCOMOD and TOTENRGY can adequately capture all the permanent innovations in RPFX over our sample period.

We test for cointegration using both the single-equation method proposed by Hansen (1990) as well as the more familiar systems approach developed by Johansen (1988) and Johansen and Juselius (1990).[12] The Hansen test simply examines whether RPFX is cointegrated with our terms-of-trade variables, while the Johansen-Juselius approach examines the total number of cointegrating vectors in the system and gives additional information on the system's dynamics. As the results in Table 2 (page 29) show, both tests provide significant evidence of

10. In an earlier paper we have shown, using Monte Carlo experiments (Amano and van Norden 1992b), that this joint testing approach leads to substantially fewer incorrect conclusions.

11. Additional tests (not shown) rejected the presence of more than one unit root.

12. Engle and Granger (1987) and Phillips and Ouliaris (1990) tests for cointegration fail to yield evidence that would allow us to reject the no cointegration null.

cointegration.[13] If there is only one cointegrating vector, the Johansen and Juselius test results need not imply that it is a cointegrating relationship for the exchange rate. In particular, it could be the case that TOTCOMOD and TOTENRGY are cointegrated and unrelated to RPFX. However, in repeating the test on the bivariate system of TOTCOMOD and TOTENRGY, we found no evidence of cointegration that was significant at the 5 per cent level. Hence, we conclude that the real exchange rate is cointegrated with our terms-of-trade variables. Our next step, then, is to formulate and estimate an error-correction model for the exchange rate that can capture both the long-run effects of terms-of-trade shocks and the transitory effects of monetary shocks.

5 The error correction model

The Engle and Granger (1987) representation theorem implies that any system of cointegrated variables that have ARIMA representations may be written in the form of an error-correction model (ECM). In particular, if we have a system of variables X that are cointegrated with cointegration vectors given by a matrix β, we may write the dynamics of the system as[14]

$$\Delta X_t = \alpha \cdot (X_{t-1} \cdot \beta) + Z_t \cdot \gamma + \varepsilon_t \qquad (3)$$

where Z is a vector of stationary variables, which may include lagged differences of X. In this form, the vector $X\beta$ gives the deviation of the system from its long-run or "equilibrium" relationships, while α gives the speeds of adjustment towards those relationships for each variable and $Z \cdot \gamma$ captures other short-run dynamics.

Our goal in estimating this model is to understand the forces that appear to be driving RPFX. We are interested in modelling the determination of TOTCOMOD and TOTENRGY only to the extent that they aid us in this primary goal. In this case, we might choose to estimate only the single ECM equation for RPFX rather than the full ECM system. In general, this would not be an advisable approach. For example, Johansen (1992) demonstrates that this will normally cause a loss in efficiency and complicates inference. More intuitively, it would be hazardous to give

13. The Johansen-Juselius test also admits some weaker evidence of a second cointegrating vector when using the λ_{max} test statistic. Since any linear combination of cointegrating vectors must also be a cointegrating vector, this implies that RPFX could be modelled as having a long-run value determined solely by TOTCOMOD or by TOTENRGY. In practice, we found that both variables are essential to capturing the behaviour of RPFX. We therefore find the evidence of a second cointegrating vector, which is statistically weak, to be unconvincing.

14. For notational clarity, we use a representation that does not necessarily imply that ε is an innovations process. This has no effect on the interpretation of α or β.

structural interpretations to a single equation for RPFX without considering the effects of changes in RPFX on TOTCOMOD and TOTENRGY.

There is, however, a special case where such a single-equation approach can be justified. Johansen (1992) demonstrates that if (i) all other cointegrating variables are weakly exogenous (in the sense of Engle, Hendry and Richard 1983) with respect to the first variable under consideration, and (ii) if there is only a single cointegrating vector, then estimation and inference on the single equation system will be equivalent to that on the full system. The weak exogeneity condition can be tested on the full system using the simple likelihood-ratio test described in Johansen and Juselius (1990).[15] We calculate this test statistic for our system of RPFX, TOTCOMOD and TOTENRGY to be 0.395, which, given a χ^2 distribution under the null of weak exogeneity, implies a marginal significance level of 82.1 per cent. This suggests that weak exogeneity is a valid working assumption and so the single-equation model we estimate below should be efficient and allow valid inference. It also implies that there is no evidence that the terms of trade respond to the deviation of the exchange rate from its long-run level, which is consistent with our desired interpretation of this long-run relationship as an exchange-rate equation.[16]

We now proceed to estimate our exchange rate equation using the non-linear least-squares methodology described by Phillips and Loretan (1991). Unlike the Engle and Granger (1987) two-step procedure, this approach simultaneously estimates both the long and short-run relationships in the form

$$\Delta Y_t = \alpha(Y_{t-1} - X_{t-1}\beta) + Z_t\gamma(L) + u_t \tag{4}$$

where Y is the dependent variable (in this case, RPFX), X and Z are vectors of integrated long-run and stationary short-run explanatory variables respectively, $\gamma(L)$ is a polynomial in the lag operator, α and β are defined analogously as in equation (3) and u_t is an independently and identically distributed mean zero error term. Phillips and Loretan (1991) find that, given the

15. This is simply a test of whether the elements of α in the equations for TOTCOMOD and TOTENRGY in equation (3) are significantly different from zero.

16. This tells us nothing about Granger-causality among these variables, however, which is investigated below. For a discussion of exogeneity and causality, see *Journal of Policy Modeling* (June 1992), particularly the introduction by Neil Ericsson.

presence of cointegration, standard inference can be used to test hypotheses about β and γ (but not α.) In general, (4) may include any number of lags in the variables.

We estimate (4) using the contemporaneous value and first four lags of Z_t over the 1973M1 to 1992M2 sample period.[17] Given the results of the previous section, we use $Y = RPFX$, $X = [Constant, TOTCOMOD, TOTENRGY]$ and Z is defined to include both the first-differences of the terms-of-trade series as well as our measure of the real interest rate differential RDIFF. By successively omitting variables with insignificant t-statistics, the equation eventually reduces to the specification reported in Table 3 (page 30), namely,

$$\Delta \mathrm{RPFX}_t = \alpha(\mathrm{RPFX}_{t-1} - \beta_0 - \beta_C \mathrm{TOTCOMOD}_{t-1} - \beta_E \mathrm{TOTENRGY}_{t-1}) + \gamma RDIFF_{t-1} \qquad (5)$$

In order to ensure that the standard errors presented in Table 3 are valid and to check that there is no additional structure in the data that may improve the model's performance, we subject the ECM to a variety of residual diagnostic tests.[18] LM and ARCH tests up to order 12 are unable to find significant evidence of autocorrelated or heteroscedastic residuals. The Jarque and Bera (1980) test, however, forces us to reject the null hypothesis of normality in favour of the non-normal alternative. Further investigation shows that the rejection is largely due to skewness of the residuals. We note that such non-normality is common in financial time series, and that given our large numbers of observations and the absence of serial correlation, this is unlikely to affect our inferences.

The specification of the error-correction model is remarkably simple, with all transitory factors captured by the single lagged value of the real interest rate differential. The speed of adjustment α is -0.038, implying that about 37.1 per cent of adjustment is completed within one year, or equivalently, a half-life of 17.9 months. The estimated long-run effects suggest that a 1 per cent improvement in commodity terms of trade leads to a 0.811 per cent appreciation of the Canada-U.S. real exchange rate, while a 1 per cent improvement in energy terms of trade results in a 0.233 per cent depreciation of the real exchange rate. It is noteworthy that the long-run parameter estimates from the ECM are similar to those given by the Johansen-Juselius full-system estimator, which gives long-run coefficients on the commodity and energy terms of trade of

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17. The Canada-U.S. dollar exchange rate began to float on 1 June 1970. Therefore, while the sample period is not strictly speaking "post-Bretton Woods," it uses data from only the most recent floating period.
18. Complete diagnostic test results are reported in Amano and van Norden (1992a).

-1.283 and 0.175, respectively.[19] Both are within two standard errors of the estimates in Table 3 (page 30). Finally, as monetary models predict, increases in Canadian short-term interest rates, *ceteris paribus*, lead to an appreciation of the Canadian dollar. It is also interesting to note that despite the inclusion of only three variables, all of which are lagged, the results imply that these economic fundamentals can account for a surprising 12 per cent of the month-to-month changes in the real exchange rate.

Although not inconsistent with economic theory, the depreciating effects of increases in the energy terms of trade run counter to the conventional view of the Canadian dollar as a petro-currency (that is, one whose value increases with higher energy prices.)[20] One explanation for this result, which we examined, was that it simply captures the effects of Canadian domestic energy policies. The Canadian government kept domestic energy prices well below world levels during the 1970s, finally deregulating them around the time of the 1985 drop in energy prices.[21] In an attempt to control for such policy initiatives, we added a variable to (5) that measured the ratio of domestic Canadian crude oil prices to world (West Texas) crude oil prices. Its coefficient was only marginally significant, and the variable's inclusion very slightly increased both the estimated coefficient and the t-statistic on TOTENRGY. We also added a series of dummy variables to the ECM that corresponded to the introduction of major Canadian energy policy initiatives. These variables were generally insignificant, and their inclusion again failed to change either the sign or significance of the coefficient on TOTENRGY. We therefore conclude that the effects of energy price shocks on RPFX do not seem to have been strongly influenced by Canadian energy policies.[22]

To give a better notion of the ECM's fit, Figure 2 (page 33) compares the behaviour of RPFX to the dynamic simulations of the model.[23] The upper panel presents the simulation when the equation is estimated using the full set of data (1973M1 to 1992M2), whereas the lower panel

19. These are the normalized values of the eigenvector corresponding to the largest eigenvalue of the estimated coefficient matrix for the level variables.
20. We note that Macklem (1992) works with a calibrated model of the Canadian economy and produces a similar long-run effect for some terms-of-trade shocks.
21. The imposition and removal of these price controls were accompanied by the imposition and removal of energy export licensing (to limit the diversion of supplies to the U.S. market) and a system of oil import subsidies.
22. The complete results are available from the authors upon request.
23. Dynamic simulations use predicted value of RPFX at t to calculate the prediction at $t+1$, so forecast errors can quickly accumulate. In comparison, the dynamic simulation of a random walk model would simply give a constant.

re-estimates (5) using data only up to 1985M12. There are several points to note. The ECM fits the exchange rate much better than a random walk in sample. While there are occasionally large forecast errors, these tend to be in periods of rapid exchange rate movement and are usually short-lived. Not only does the equation capture the major fluctuations in the exchange rate over this period, it predicts major turning points in advance of the actual occurrence, as we see both around the peak in early 1986 and the low in 1991. Furthermore, as we see in the lower panel, we can obtain similar results using estimates based only on the first two-thirds of our sample. This suggests that virtually the whole of the appreciation of the Canadian dollar after 1986 can be explained on the basis of factors which were already operational during its previous long depreciation. It also suggests that the equation's parameter estimates may be stable over time, a point we now examine more closely.

Figure 3 (page 34) gives another indication of the parameter stability of our model. We perform a sequential Chow test and find no evidence of a shift in coefficient estimates that is nominally significant at the 10 per cent level.[24] We also use a more general approach to examine the stability of the parameter estimates, estimating (4) over the period 1973M1-1976M1, then adding observations one by one to the end of the sample period and updating the parameter estimates at each step. Figure 4 (page 35) shows how these estimates and their nominal 99 per cent confidence intervals evolve as observations are added. Starting with the estimates of the constant term, we find evidence of instability around the time of the second OPEC oil shock as our parameter estimates drift upwards. After that period, however, the estimates are quite stable. Estimates of the coefficients of TOTCOMOD and TOTENRGY behave similarly, shifting somewhat around 1979 but remaining within the calculated confidence intervals thereafter. The coefficient of RDIFF tends to drift towards zero slightly over time, but remains within its confidence intervals for the entire sample. Accordingly, this suggests that there is some evidence of a shift in parameter estimates around the time of the 1979 oil shock, but that the estimates have

24. Andrews (1990) 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 critical values and inference is biased against the null of stability. To control for this effect we use the critical values calculated by Andrews (1990).

remained fairly stable since.[25]

As noted in the Introduction, post-sample prediction ability has become an important yardstick in assessing the validity of exchange rate models. The stability of the equation's parameter estimates over the last decade suggests it might forecast well out of sample, a question we consider in the next section.

6 Forecasting the real exchange rate out of sample

If we wish to properly assess the model's out-of-sample forecasting ability, we need to establish stronger exogeneity conditions than we have used to this point. As Ericsson (1992) notes, valid prediction from a conditional model like (5) requires strong rather than just weak exogeneity. In practical terms, we need to ensure that RPFX does not Granger-cause TOTCOMOD, TOTENRGY or RDIFF, in addition to the maintained hypothesis of weak exogeneity. If this non-causality were not the case, then even a misspecified model might give better conditional forecasts than a random walk out of sample. This is because we can still gain information about the evolution of our dependent variable by conditioning our forecasts on variables that are Granger-caused by our dependent variable.[26]

Valid testing for Granger-causality in cointegrated systems is a subject at the forefront of time-series econometrics (for example, see Toda and Phillips 1991a, 1991b). Fortunately, our model has a simple structure with a single cointegrating vector, so appropriate testing procedures have been established (see Sims, Stock and Watson 1990). In particular, we may use standard Granger-causality tests on a vector autoregression level representation of our system, provided we have a maintained hypothesis of one cointegrating vector and that we test the exclusion restrictions on one variable at a time.[27] The results from the Granger-causality tests are reported in Table 4 (page 30). We first examine the system consisting of the real exchange rate and the commodity and energy terms of trade. Next we examine the causal relationship between the exchange rate and the real interest rate differential by including it in the system we considered

25. As a practical matter, we do not consider the limited evidence of instability around the 1979 oil shock to be very serious, taken by itself. Estimates of β_C and β_E are highly correlated, and TOTENRGY is dominated by major oil-price shocks in 1974, 1979 and 1985. The instability that we find simply reflects the fact that at least two of these shocks are needed to accurately estimate their effects, a conclusion that we do not consider to be unreasonable.

26. Note that the presence of cointegration implies Granger-causality should exist in at least one direction between our long-run variables.

27. Since inferences in this case are based on asymptotic approximations, we use the limiting χ^2 critical values instead of their more common F-distributed counterparts.

above. We do not find significant evidence that any of these variables are Granger-caused by RPFX. These results and the earlier evidence of weak exogeneity allow us to conclude that the restriction of strong exogeneity is satisfied, so we can go on to perform valid conditional forecasting.

We now follow the methodology used by Meese and Rogoff (1983) and compare the out-of-sample forecasts produced by (5) to those generated by a random walk. Specifically, we begin by estimating (5) on data up to January 1986 (the peak of the RPFX), and then generate dynamic conditional forecasts for one-, three-, six- and twelve-month horizons.[28] Another month of data is then added, the equation is re-estimated and new forecasts are generated. This process is repeated until the end of our data set (February 1992). Finally, forecast errors are then calculated for all estimation periods and all horizons. Table 5 (page 31) reports Theil's U statistic, which is the ratio of the ECM forecast's root-mean squared error (RMSE) to the RMSE of a random walk forecast. Values less than one therefore imply that the model performs better than a random walk while values greater indicate the reverse.[29] We also show how the U statistic changes as we vary the starting period of our experiment.

Starting in January 1986, the exchange rate model generates a smaller RMSE than the random walk model at all horizons under consideration, with our model's performance improving as the forecasting horizon increases. The same is true for most other starting dates as well, although the results from the experiment beginning in January 1988 show much less difference across forecast horizons than the rest. The key exceptions are the forecasts which begin in January 1989, in which the random walk tends to perform better than the exchange rate model, with the difference again increasing as the forecast horizon lengthens. Presumably, this implies that the model did not explain exchange rate changes very well around the 1989-90 period, but improved again thereafter. Therefore, we conclude that while our exchange rate equation usually performs better than the random walk, it is possible to find a period in which it does worse.

While the former result is encouraging, it would be useful to determine whether our forecasting model performs *significantly* better than a simple random walk. To address this issue we now turn to the concept of forecast encompassing, which allows us to test formally the

28. Note that the realized values for TOTCOMOD, TOTENRGY and RDIFF are used to generate all forecasts.
29. Amano and van Norden (1992a) also report the mean absolute error for each forecast. The conclusions remain the same.

forecast performance of different models.[30] A model is said to encompass a competing formulation if its forecasts help to predict the forecast errors of its competitor, but its competitor's forecasts give no information about the first model's forecast errors.[31] Tests of forecast encompassing are easily calculated, based on an artificial regression involving only forecasts and realized outcomes

$$u_t^A = \alpha + \beta F_t^B + \varepsilon_t \tag{6}$$

where u_t^A is the forecast error from model A, F_t^B is the forecast of model B, and α and β are parameters. If β is statistically significant, this implies that model B's forecast encompasses model A, suggesting that model B provides information not admitted by model A. Since fully efficient multi-step forecasts are linear combinations of the one-step-ahead forecasts, only the one-step-ahead forecast need be considered here.

Table 6 (page 31) reports the marginal significance level of β in (6), again as a function of the starting date.[32] In every case we have evidence significant at the 10 per cent level that our model's projections help to predict random walk forecast errors out of sample, and in 2 of these cases it is significant at the 1 per cent level. There is no evidence of the reverse that is significant at the 10 per cent level. Therefore, we conclude that our ECM seems to forecast significantly better than the random walk out of sample.

The model's out-of-sample forecasting performance leads naturally to the question of whether this implies a violation of semi-strong market efficiency. The answer is - not necessarily. Such a violation would imply that excess returns are predictable; we have simply shown that real exchange rates changes are predictable. As Levine (1989) notes, however, expected excess returns can be shown to be the sum of expected movements in the real exchange rate and expected

30. For general expositions on forecast encompassing, see Chong and Hendry (1986) and Marquez and Ericsson (1990).

31. Marquez and Ericsson (1990) state a weaker condition, namely that the first model's forecasts help to predict the forecast errors of its competitor. In infinite samples, this will be the same as the condition we stated. However, these conditions may differ in finite samples. In the applied work that follows, we use the stricter condition, which follows Marquez and Ericsson's discussion of their empirical results.

32. By allowing for a non-zero constant term, a more powerful encompassing test may result if the forecast errors of model A are systematically biased. For compactness we report only the tests of $\beta = 0$. Complete tests of both $\alpha = 0$ and $\beta = 0$, and $\alpha = \beta = 0$, which give similar conclusions, may be found in Amano and van Norden (1992a).

real interest rate differentials between the two currencies.[33]

$$x_{t+1} = E_t(\Delta RPFX_{t+1}) + E_t(\tilde{r}_{t+1}) + \varepsilon_{t+1} \tag{7}$$

Given the evidence above that real exchange rate movements seem to be predictable, excess returns x_{t+1} should be predictable unless there are offsetting movements in expected real interest rate differentials $\tilde{r}_{t+1}$.

We can test this indirectly with the regression

$$x_{t+1} = \kappa + \varpi \cdot (F_{t+1}^{\mathrm{ECM}} - RPFX_t) + v_t \tag{8}$$

where $(F_{t+1}^{\mathrm{ECM}} - RPFX_t)$ is the model's forecast of the change in the real exchange rate from t to $t+1$ conditional on information available at t. The error term v_t is composed of three parts: the expectational error ε_{t+1} from (7), a measurement error that reflects the difference between $E_t(\Delta RPFX_{t+1})$ and $(F_{t+1}^{\mathrm{ECM}} - RPFX_t)$, and an omitted variable $E_t(\tilde{r}_{t+1})$. If we estimate (8) by ordinary least squares, we should expect the measurement error to bias our estimate of ϖ towards zero. Furthermore, since (7) predicts that $E_t(\tilde{r}_{t+1})$ will be highly negatively correlated with $(F_{t+1}^{\mathrm{ECM}} - RPFX_t)$, this should bias our estimate of ϖ still further towards zero. When we estimate (8), we find that our estimate of ϖ is statistically significant (see Table 7, page 31), and we are unable to reject the joint hypothesis that ϖ equals unity and κ is zero (the marginal significance level for the joint hypothesis is 0.721). The model therefore seems to imply a rejection of semi-strong market efficiency.[34] The fact that the estimated coefficient on forecast real exchange rate changes is very close to unity also suggests that expected real interest rate differentials show little tendency to offset expected real exchange rate movements, and that the ECM's forecast is a good proxy for the expected real exchange rate.

In sum, the out-of-sample forecasting results reported above are in contrast to those of other researchers such as Meese and Rogoff (1983), who find that no model of exchange rate determination is able to consistently out-perform a simple random walk. Our ECM usually

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33. Levine (1989, 164-66). He defines excess returns as the prediction error of the forward exchange rate.

34. Violating market efficiency also requires that prediction is possible using public information that was available before the beginning of the period over which returns are calculated. Our model uses terms-of-trade data, which are available only after a lag of one quarter and which may be subsequently revised. On the other hand, its movements are dominated by shifts in energy and other commodity prices, which are instantaneously available to the public.

produces smaller forecast errors than a random walk over several forecast horizons, and it passes a variety of tests stricter than those that are typically used in other studies (see for example Schinasi and Swamy 1989). We find that the results are generally robust to the starting date selected, that the superior performance of the ECM is statistically significant, and that the results cannot be attributed to Granger-causality running from RPFX to our explanatory variables. We also find that the relationship between the real exchange rate and the terms of trade seems to imply a violation of semi-strong market efficiency.

7 Decompositions of real exchange rate movements

Having established that the terms of trade and real interest rate differentials together can give a stable and reasonable model for the Canada-U.S. real exchange rate, we now consider the relative importance of these two kinds of variables in explaining exchange rate behaviour over the last twenty years. Using our ECM, we can show that any observed change in the exchange rate can be expressed as the sum of five factors: the prediction error of a dynamic simulation of the error-correction model, gradual adjustment towards some initial conditions, and shifts due to changes in TOTCOMOD, TOTENRGY or RDIFF.[35] By performing this decomposition for dynamic simulations with different starting points and horizons, we can build series of decompositions, showing the contribution of a given factor to exchange rate changes over a certain time period. We performed this decomposition for all possible starting points in the 1973-92 period and for all horizons from 1 to 60 months. This gave us a set of series that represented the fraction of the change in RPFX that is attributable to a given source (as a function of the starting date of the simulation).

Examining the variance of these series then gives us a historical average decomposition of the variance of changes in RPFX over a given horizon. Unfortunately, since these series are correlated, their variances do not sum to the variance of changes of RPFX. Interpreting these results therefore requires that we orthogonalize our components (that is, that we apportion their covariances.) To do so, we use the Choleski orthogonalization commonly used in the variance decompositions of VAR models. Specifically, we assume a Granger-causal ordering of

35. See Appendix 2.

TOTENRGY, TOTCOMOD, RDIFF, the lagged adjustment terms, and the dynamic forecast error.[36]

The resulting orthogonalized decomposition is shown in Figure 5 (page 37). Since the simulation error captures the fraction of changes in RPFX not explained by the ECM, its steady decline shows that the model explains an increasing degree of the variance of these changes as we lengthen the horizon over which changes occur. We see that the model explains 22 per cent of the variance over 1-month horizons, 50 per cent over 12-month horizons, and 93 per cent over 5-year horizons. This high explanatory power for long-term changes in exchange rates seems remarkable. Of this, the lagged adjustment term is the dominant factor at all horizons less than three years, accounting for up to 64 per cent of the variance. Changes in any of the three explanatory variables never account for more than 10 per cent of the variance at these shorter horizons. This presumably reflects the smoothness of TOTCOMOD and TOTENRGY, which means they vary relatively little over such horizons, although the exchange rate is still considerably influenced by these factors. Over longer horizons, however, the importance of lagged adjustment again diminishes and is eventually surpassed by changes in TOTCOMOD and, to a slightly lesser extent, TOTENRGY. At a 5-year horizon, these account for 38 per cent and 25 per cent, respectively. In contrast, RDIFF never explains more than 11 per cent, and its explanatory power falls to 5 per cent at the longest horizons.

It is also interesting to consider what interpretation the ECM offers for historical movements in the Canadian dollar's real exchange rate. To do so, we can dynamically simulate the ECM starting in January 1973 and then decompose changes in the simulated values as before into those arising from changes in RDIFF, TOTCOMOD and TOTENRGY.[37] Since we are now decomposing past movements rather than variances, there is no need to orthogonalize the components. The top half of Figure 6 (page 38) shows the actual value of RPFX and its predicted value from the dynamic simulation shown in the top half of Figure 2 (page 33). The dotted

36. Given weak exogeneity, we would expect the forecast error to come last in this ordering, and we should place lagged adjustment terms after changes in the independent variables. Furthermore, we would expect changes in world energy and commodity prices to cause changes in Canadian monetary policy, rather than vice versa, so the terms-of-trade measures should precede the interest rate differential. Results were not very sensitive to the ordering of the two terms of trade measures.

37. We omit the contribution of the deviation from the long-run equilibrium at the start of the simulation, since this is roughly constant after the first few years of the simulation. It would account for roughly another 5 per cent depreciation of the simulated value in the first subperiod considered below.

vertical lines divide the sample period into seemingly distinct periods of exchange rate behaviour: the strong Canadian dollar from January 1973 to October 1976, the rapid depreciation leading to a plateau around June 1981, the further depreciation to the record level of February 1986, and the subsequent appreciation that brings us to virtually the end of our sample. The bottom half of the graph shows the contribution of each of the explanatory variables to the change in the simulated values over of these four subperiods.

There are several interesting facts to take away from this historical analysis. First, the role of monetary policy factors captured through RDIFF appears to be small compared to that of terms-of-trade shocks. This is consistent with the monetary model's inability to explain much of the floating exchange rate behaviour. It is also consistent with Baxter (1992), who examines evidence for other countries and finds that real interest rate differentials can only account for a small portion of the variation in the real exchange rate. Second, the net movement of the real exchange rate in response to shocks seems to have varied greatly depending on the extent to which different shocks re-enforced or cancelled one another. Thus the first OPEC energy price shock had little net effect on the exchange rate at first, as it was largely offset by a major commodity price shock, whereas the second OPEC shock was associated with a major depreciation of the Canadian dollar. Third, the largest of the three components of exchange rate movements now seems to be that resulting from energy price shocks, which have the largest single effect in three of the four subperiods we consider. Finally, it is interesting to note the substantial time it takes for the full effects of a shock to be felt. In particular, large persistent energy price shocks still had significant effects on the exchange rate four years after their initial impact. However, very short-lived shocks, even if large, had little noticeable effect. For example, during the Gulf War crisis in 1990-91, TOTENRGY rose by 57 per cent in four months, only to fall 51 per cent in the following four months.[38] Yet its effect on either the actual exchange rate or the equation's simulated value is difficult to detect.

8 Conclusions

This paper set out to see whether terms-of-trade shocks can explain much of the historical variation in the Canada-U.S. real exchange rate. Along the way, we have uncovered several interesting findings. First, there is a cointegrating relationship for the real exchange rate that

38. In comparison, the second OPEC energy shock caused TOTENRGY to rise 66 per cent in 16 months.

appears to be stable over more than a decade. Second, tests of causality in this relationship have shown that while terms-of-trade shocks have a significant effect on the exchange rate, the reverse is not true. This is consistent with the standard assumptions of the small open economy model. Third, a simple error-correction model that uses terms-of-trade variables seems to have an empirically significant ability to forecast exchange rate changes out of sample, in the sense of Meese and Rogoff (1983). So far as we know, this is the only model for the Canada-U.S. exchange rate to make such a claim. As for our original question, we find that terms-of-trade changes can potentially explain much of the variation in the real exchange rate over the last 19 years, particularly over longer horizons. While the monetary factors we considered occasionally made important contributions to real exchange rate movements, they generally accounted for less than 10 per cent of the exchange rate's overall variance.

While we believe our results demonstrate the potentially important role terms-of-trade shocks may play, it should be noted that these results also require certain restrictions on the role of these shocks in order for us to arrive at this conclusion. For example, in our model the effects of such shocks accumulate over a period of a few years, so that short-lived shocks (such as the Gulf War energy price shock) have little noticeable effect. More strikingly, however, the results imply that while increases in non-energy commodity prices tend to strengthen the Canadian dollar, increases in energy prices tend to weaken it. While not inconsistent with standard models of a small open economy, this result runs counter to the view of the Canadian dollar as a "petro-currency," and does not appear to be explained by Canadian domestic energy policies. We leave this for future research.

Appendix 1: Data definitions

The data used in our analysis include[1] the monthly average of the noon daily spot Canada-U.S. exchange rate (B3400); 30-day Canadian prime corporate paper rates in PCPA (B14039); secondary market yields on long-term Canadian corporate bonds in PCPA (B14016); 30-day U.S. commercial paper in PCPA (B54416); secondary market yields on long-term AAA U.S. industrial bonds (Moody's Bond Survey); consumer price index (1986=100) all items (P484000-CANSIM); U.S. consumer price index (1982-84=100) all items (m.cusa0-Data Resources, Incorporated). Commodity and energy terms of trade are defined as

TOTENRGY = (0.06138*B1503+0.04104*B1504+0.07613*B1505)/PM

TOTCOMOD = (PX/PM) - TOTENRGY

where

PX = 0.4664*B1501+0.06138*B1503+0.04104*B1504+0.07613*B1505
+0.77484*(B1502+B1506+B1507+B1508+B1509)

PM = 0.2986*B1558+0.7014*(B1554+B1555+B1551+B1556+B1557
+B1559)

The following are export prices: B1503, crude petroleum; B1504, natural gas; B1505, other energy products; B1501, wheat; B1502, other farm and fish products; B1506, lumber and sawmill products; B1507, pulp and paper; B1508, other metals and minerals; B1509, chemicals and fertilizer. The following are import prices: B1558, machinery and equipment; B1554, construction materials; B1555, industrial materials; B1551, food; B1556, motor vehicles and parts from the U.S.; B1557, motor vehicles and parts from the rest of the world; B1559, other consumer goods.

1. All data are from the *Bank of Canada Review* unless otherwise specified.

Appendix 2: Decomposition of the error-correction model

Using our error-correction model, we can show that any observed change in the exchange rate can be expressed as the sum of five factors: the prediction error of a dynamic simulation of the error-correction model, gradual adjustment towards some initial conditions, and shifts due to changes in TOTCOMOD, TOTENRGY or RDIFF. To see this, start with a simple error-correction model written in the form

$$y_t - y_{t-1} = \alpha \cdot (y_{t-1} + X_{t-1}\beta) + \varepsilon_t$$

then solved back to some initial period to give

$$y_t = \left(\sum_{j=1}^{t} (1+\alpha)^{t-j} \cdot \{\alpha X_{j-1}\beta + \varepsilon_j\} \right) + (1+\alpha)^t \cdot y_0$$

this can then be partitioned as

$$y_t - y_0 = \left(((1+\alpha)^t - 1) y_0 + \sum_{j=1}^{t} (1+\alpha)^{t-j} \cdot \alpha X_0 \beta \right) + \left(\sum_{j=0}^{t-1} (1+\alpha)^{t-j-1} \cdot \alpha (X_j - X_0) \beta \right)$$
$$+ \left(\sum_{j=1}^{t} (1+\alpha)^{t-j} \cdot \varepsilon_j \right)$$

The first term in parentheses is that arising from lagged adjustments to past values, the second term is that due to changes in the independent variables, and the third term is the forecast error of our dynamic simulation.

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Table 1:
Tests for unit roots and stationarity
Sample Period: 1973M1 to 1992M2

Variable[a]	ADF Lag Length[b]	ADF	PP[c]	KPS[d]
RPFX	13	-1.790	-1.342	0.834**
TOTCOMOD	21	-2.578	-2.217	0.521*
TOTENRGY	20	-1.429	-2.129	0.517*
RDIFF	18	-3.212*	-5.233**	0.054

a. Henceforth, * and ** represent significance at the 5 and 1 per cent levels, respectively. The unit-root and cointegration critical values are calculated from the response surface estimates of MacKinnon (1991).

b. The ADF test uses the data-dependent lag selection procedure applied by Hall (1989). We start the testing-down with the ADF lag length set equal to twice the seasonal frequency or 24.

c. The Phillips-Perron test statistic is calculated using the long-run variance estimator developed by Andrews (1991) and Andrews and Monahan (1992).

d. The long-run variance is estimated using the Newey and West (1987) procedure. The critical values are taken from Kwiatkowski, Phillips, Schmidt and Shin (1992).

Table 2:
Tests for cointegration

Hansen ADF and PP Tests		
H-ADF Lag Length[a]	H-ADF	H-PP[b]
12	-3.517**	-13.369**
Johansen and Juselius Test[c] (20 Lags)[d]		
No. of Cointegrating Vectors under the Null hypothesis	Trace Statistic	λ^{max} Statistic
Less than 1	47.752**	28.536**
Less than 2	19.216	16.066*
Less than 3	3.150	3.150

a. See Table 1, footnote b (page 28).

b. See Table 1, footnote c (page 28).

c. The Johansen-Juselius test statistics and their distributions differ, depending on whether one imposes the restriction that any cointegrating vectors must also annihilate drift terms in each variable. Since we were unable above to find any significant evidence of drift in our variables, there should be no drift in our variables that is not annihilated by the cointegrating vectors. Therefore, this table presents the results of the test that imposes the annihilation restriction. Not imposing this restriction leads to the same conclusions. Furthermore, the likelihood-ratio test of this restriction under the maintained hypothesis of one cointegrating vector is only 0.146, so the restriction cannot be rejected at even the 10 per cent significance level. Finally, results from the Johansen (1988) tests, which do not allow for drift in any of the integrated variables, gave evidence of exactly one cointegrating vector for all reasonable lag lengths.

d. Appropriate lag lengths for the Johansen and Juselius test are determined using standard likelihood ratio tests with a finite-sample correction. However, depending on the exact critical values used, this test suggested using 15, 20 or 23 lags. Fortunately, the cointegration results were not sensitive to the choice of lag length.

Table 3:
Error correction model estimates

$$\Delta RPFX_t = \alpha(RPFX_{t-1} - \beta_0 - \beta_C TOTCOMOD_{t-1} - \beta_E TOTENRGY_{t-1}) + \gamma RDIFF_{t-1}$$

Sample: 1973M1 to 1992M2 (230 Obs.)

R-squared: 0.123

Variable	Parameter Estimate	Standard Error	t-Statistic	Significance Level
Speed of Adjustment α	-0.038	0.011	-3.446	-[a]
Constant	0.552	0.097	5.681	0.000
TOTCOMOD	-0.811	0.296	-2.736	0.006
TOTENRGY	0.223	0.060	3.700	0.000
RDIFF	-0.187	0.043	4.390	0.000

a. The t-statistic for this parameter does not have the standard distribution under the null, so conventional significance levels do not apply.

Table 4:
Granger causality results[a]

Dependent Variable	Test Statistic	Significance Level
TOTCOMOD	15.200	0.765
TOTENRGY	13.419	0.859
RDIFF	27.102	0.132

a. The results are based on tests with 20 lags. The conclusions, however, are robust for the lag range (15 to 23) suggested by the above likelihood ratio tests (see Table 2, footnote d, page 29).

Table 5:
Theil's U statistic
Forecast comparisons: Model-based versus random walk

Horizon (Months)	1	3	6	12
Starting Period:				
January 1986	0.962	0.885	0.829	0.676
January 1987	0.957	0.901	0.800	0.645
January 1988	0.980	0.965	0.974	0.940
January 1989	0.999	1.043	1.120	1.456
January 1990	1.000	0.958	0.883	0.819
January 1991	0.985	0.922	0.851	0.565

Table 6:
Forecasting encompassing tests[a]

Starting Period:	H_0: ECM does not encompass RW	H_0: RW does not encompass ECM
January 1986	0.013	0.280
January 1987	0.003	0.113
January 1988	0.000	0.151
January 1989	0.026	0.898
January 1990	0.071	0.581

a. Reported values are the marginal significance levels of the hypothesis that $\beta = 0$ in (6).

Table 7:
Market efficiency test regression results
Sample: 1977M9 to 1992M1 (173 Obs.)
R-squared: 0.073 Durbin-Watson: 2.028

	Parameter Estimate	Standard Error	t-statistic
Constant	-0.001	0.001	-0.671
Predicted	0.879	0.239	3.670

Figure 1: Raw data

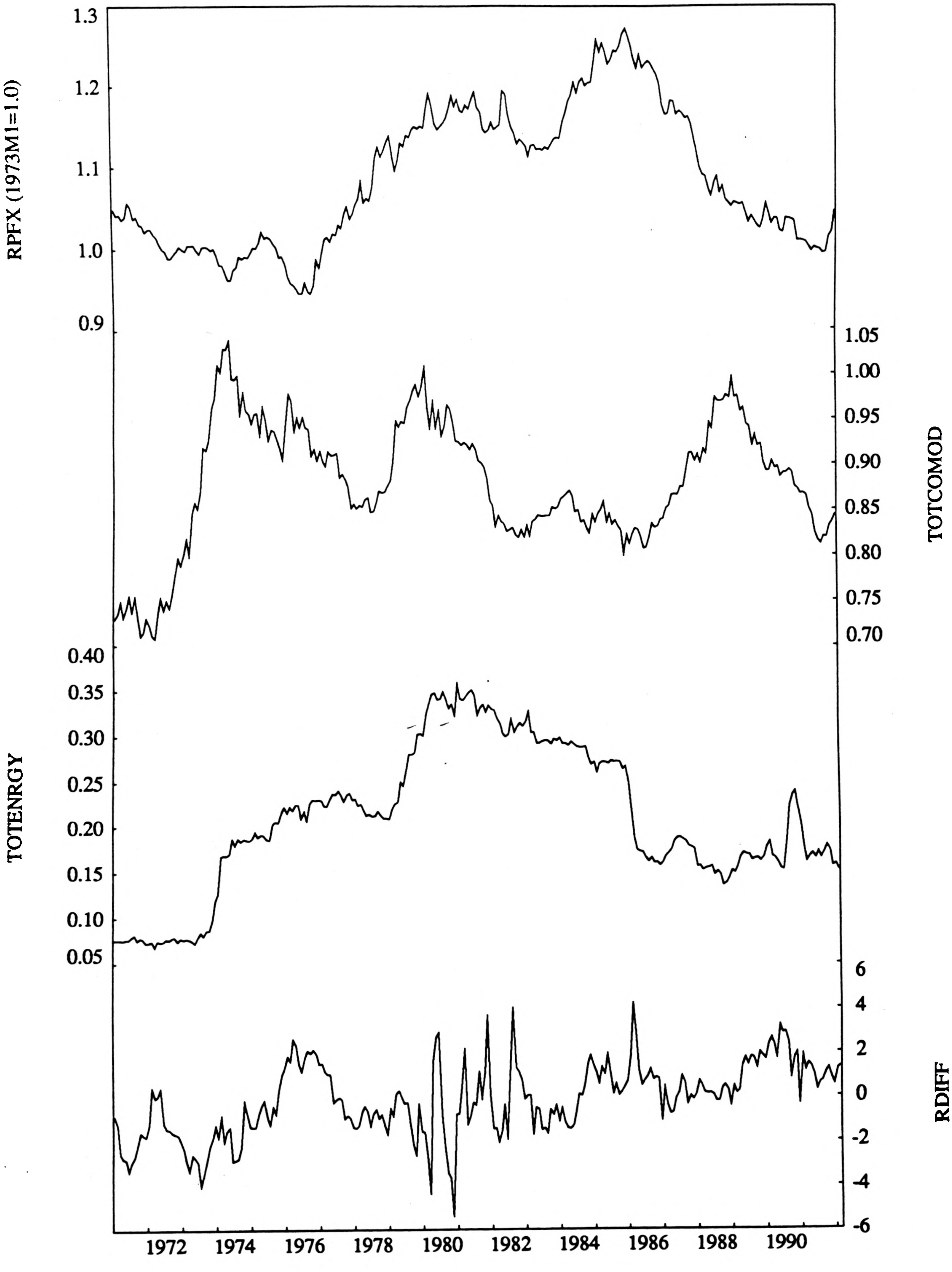
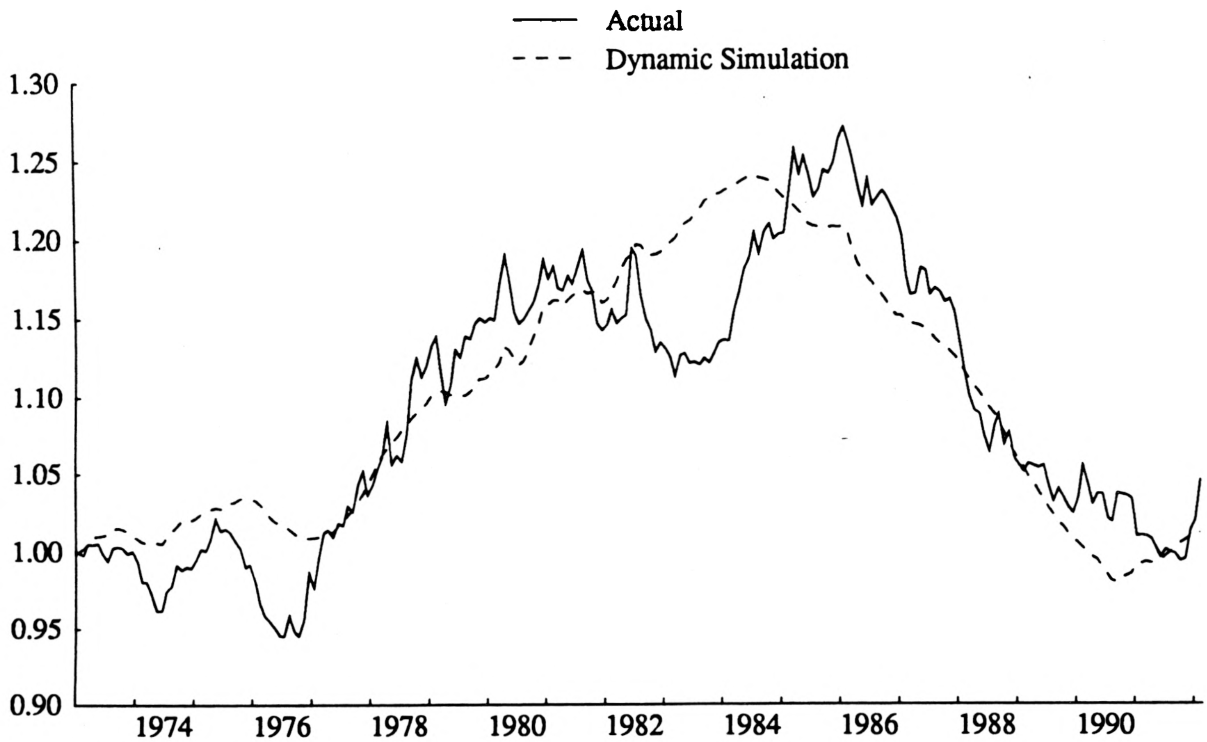


Figure 2: Dynamic simulations of the RPFX model

Simulations based on the full sample



Simulations based on the sample ending in 1985M12

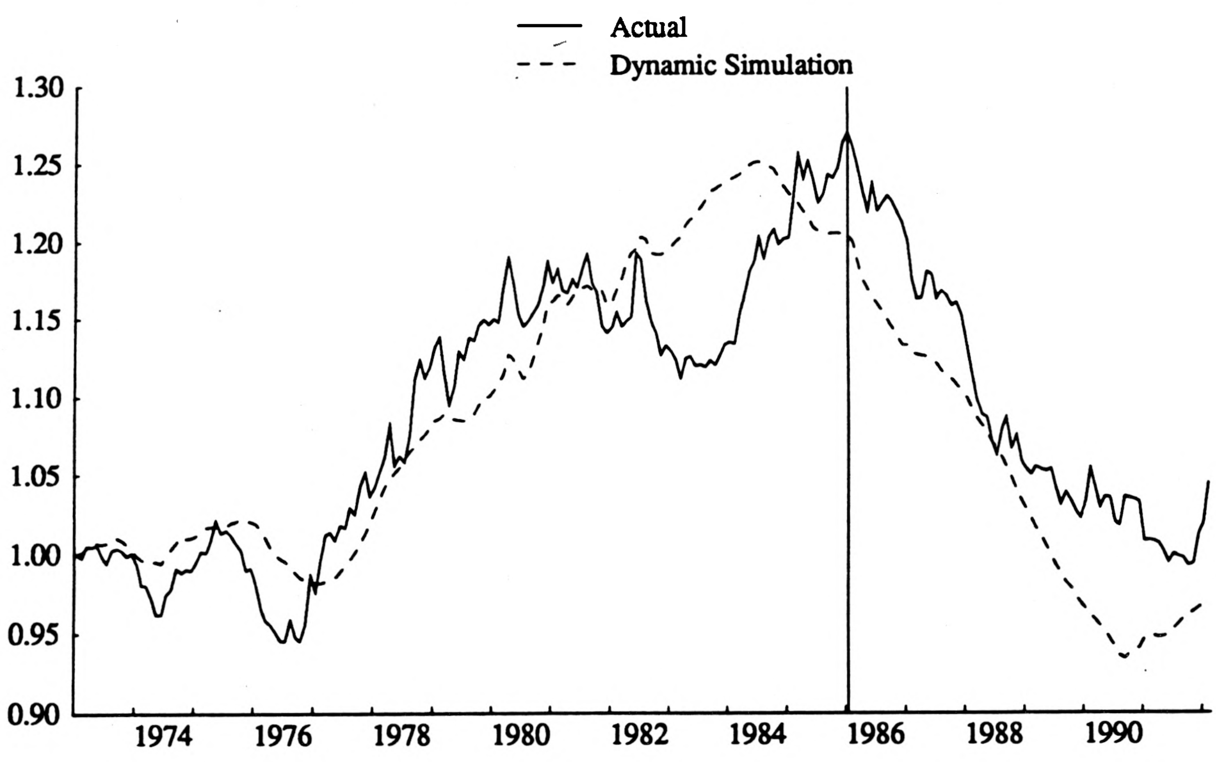
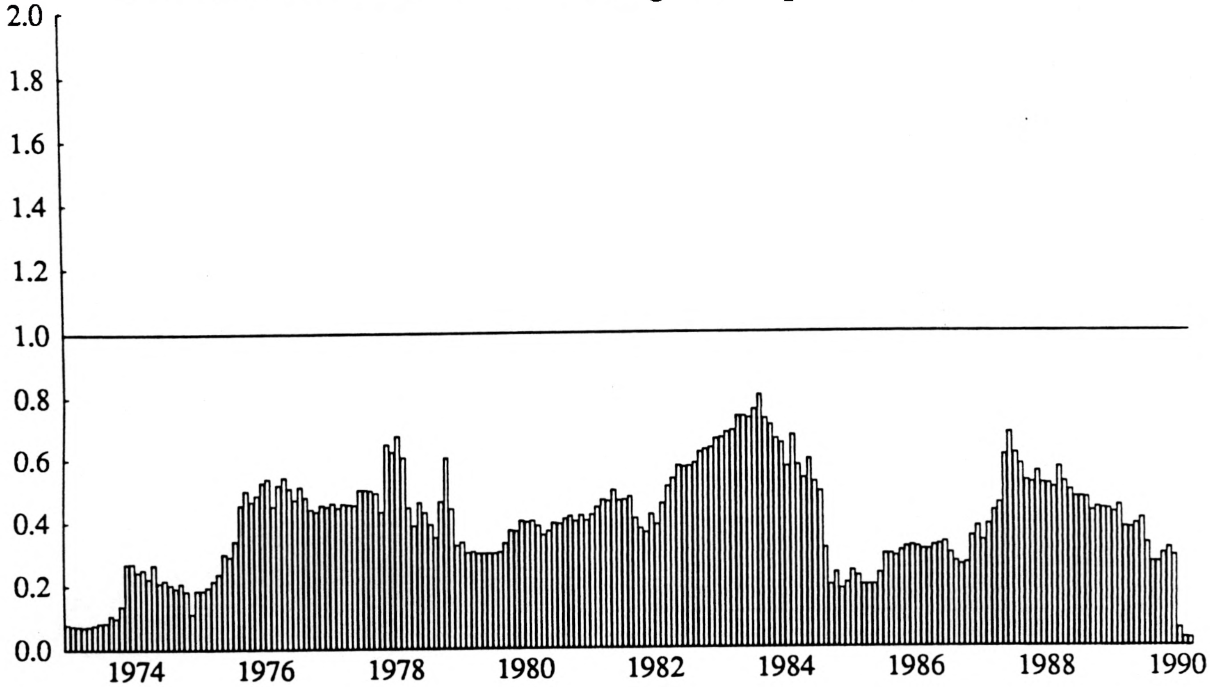


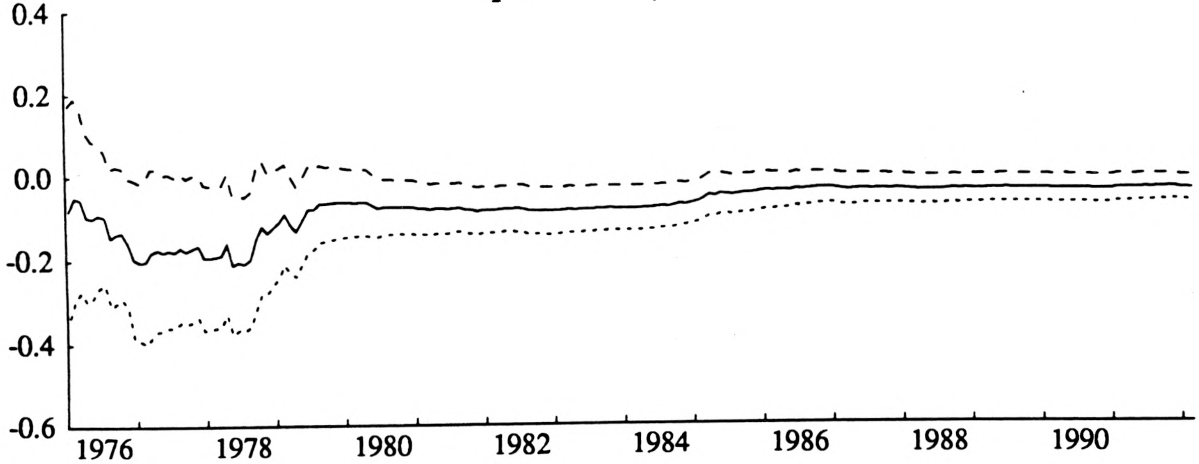
Figure 3: Sequential Chow tests
Test statistics are normalized using the 10 per cent critical value



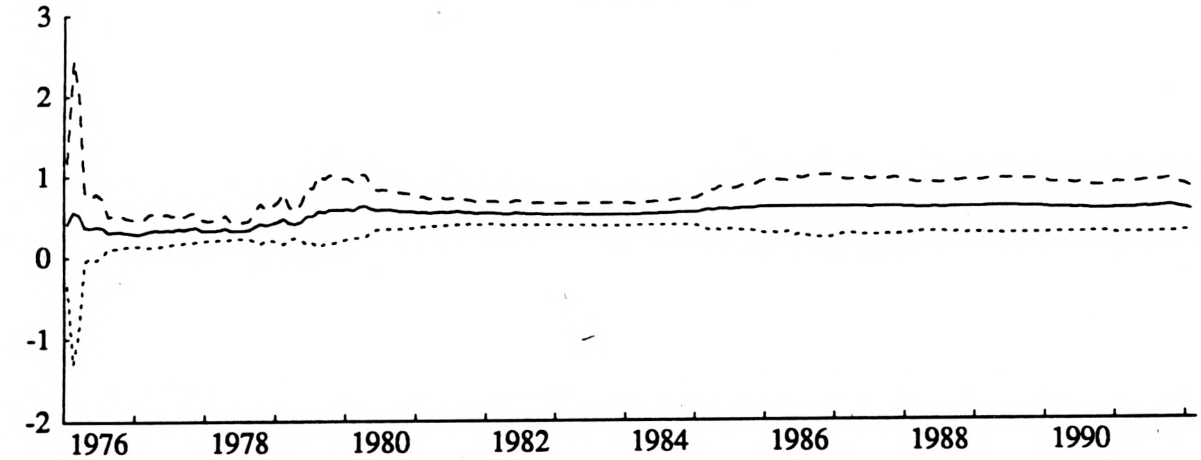
10 per cent critical value = 14.4

Figure 4: Recursive parameter constancy

Speed of adjustment



Constant



Commodity terms of trade

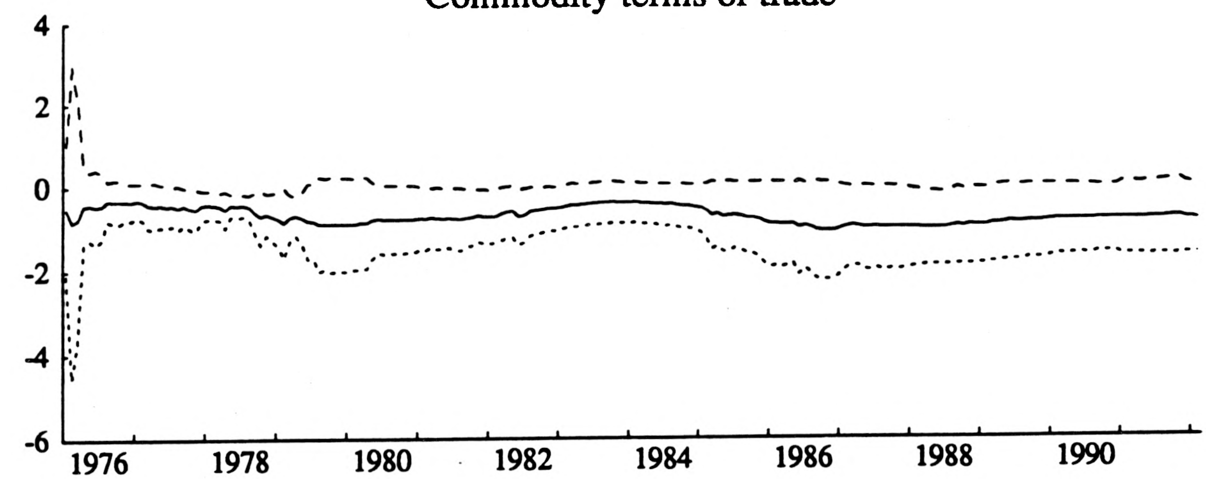
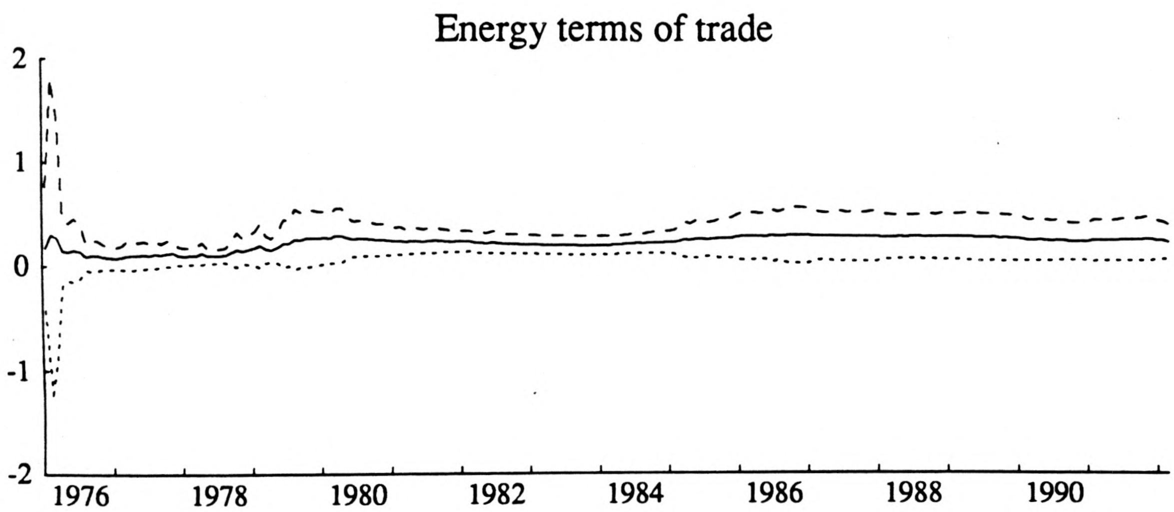


Figure 4: Recursive parameter constancy (continued)



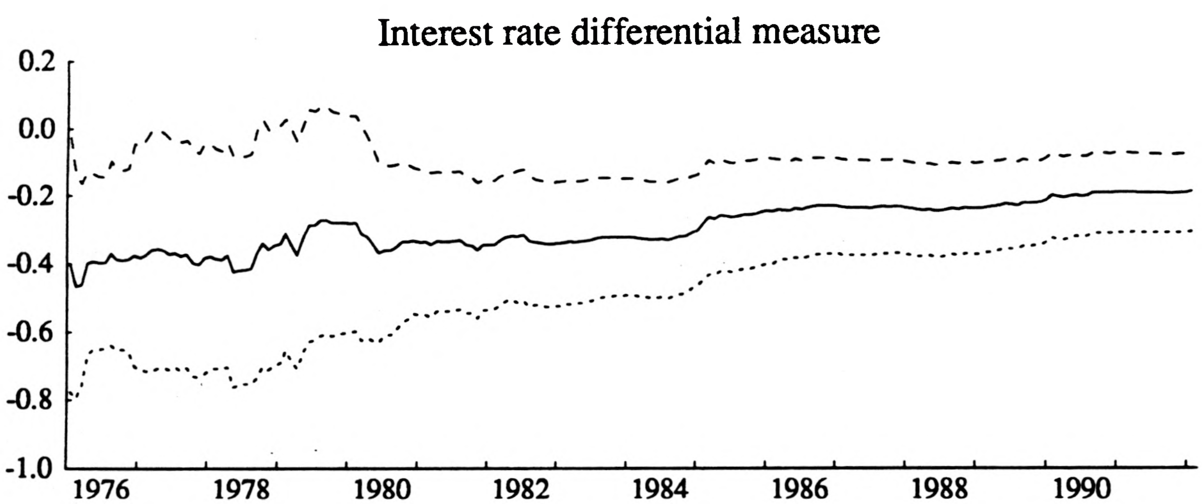


Figure 5: Variance decomposition of exchange rate movements

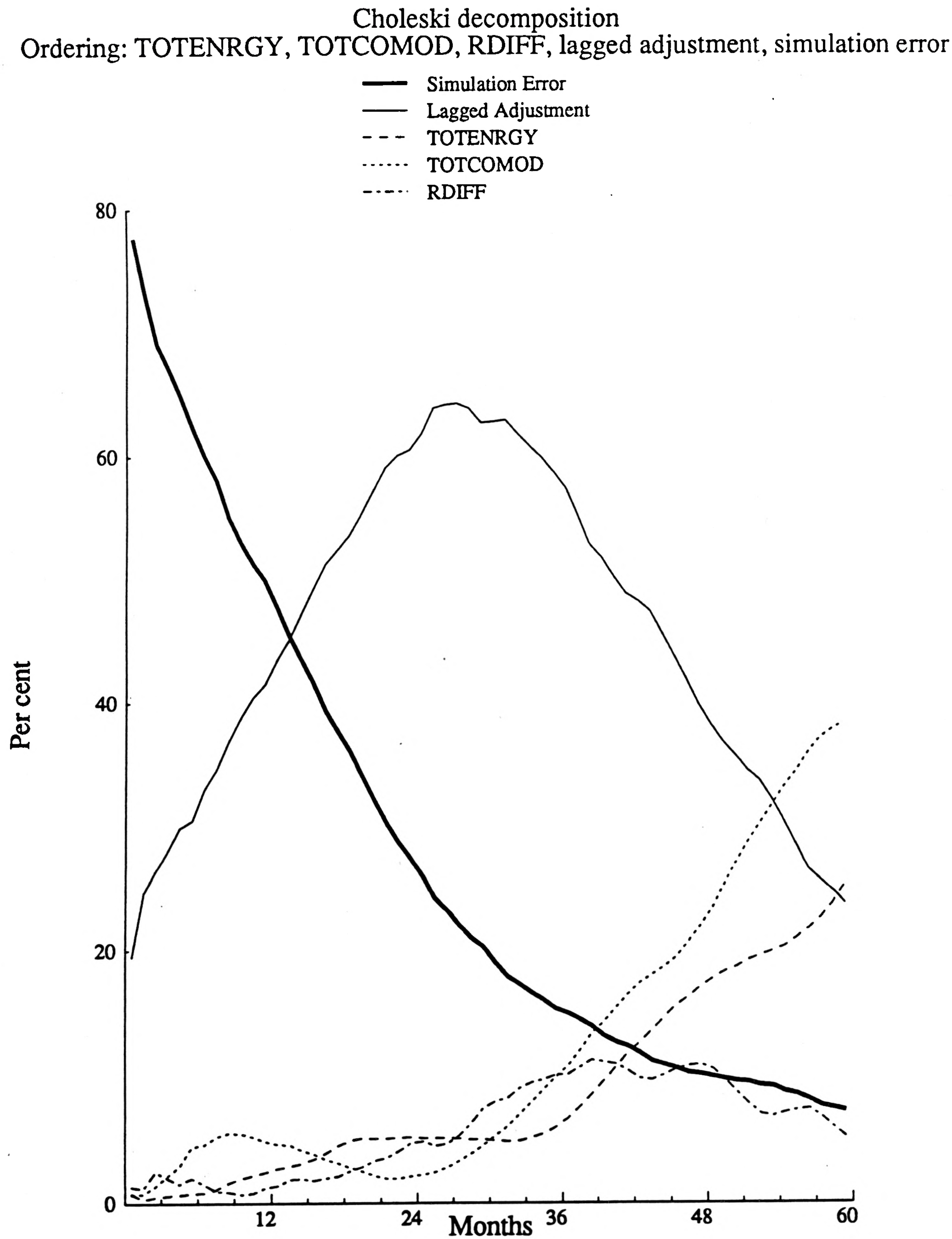
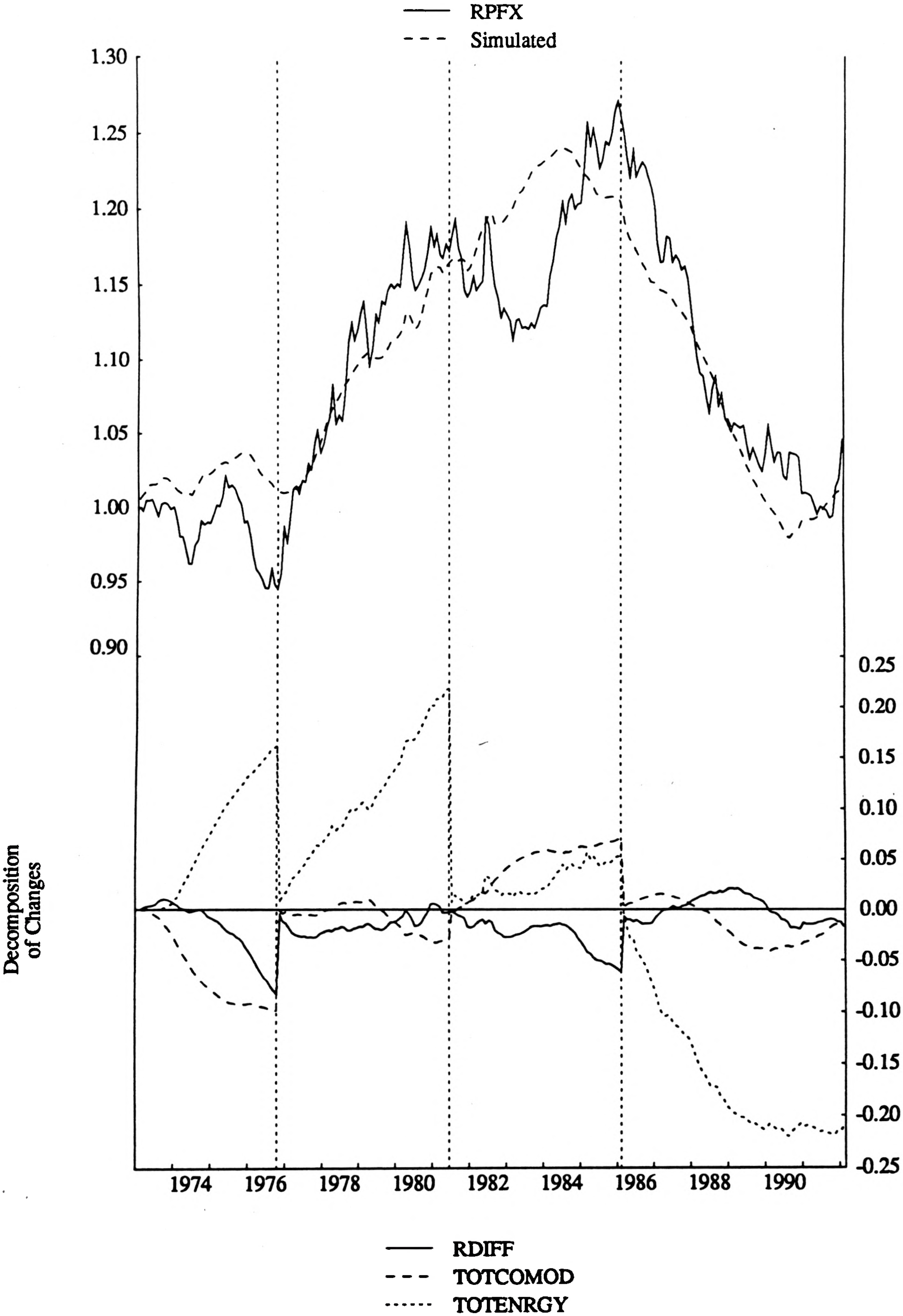


Figure 6: Historical decomposition of exchange rate movements



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